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Does the balance of power within a family matter? The case of the Retirement Equity Act

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Abstract

This paper studies within-family decision making regarding investment in income protection for surviving spouses using a simple and tractable Nash-bargaining model. A change in US pension law (the Retirement Equity Act of 1984) is used as an instrument to derive predictions from the bargaining model about the household demand for survivor annuities and life insurance and to contrast these with the predictions of the classical single-utility-function model of the household. In the empirical part of the paper, the predictions of the classical model are rejected in favor of the predictions of the Nash-bargaining model.

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1. Introduction

Most economic theory assumes that household behavior is determined by a rational agent maximizing a single household utility function. This means that the behavior of multiperson household can be described as decisions made by a (possibly benevolent) dictator within a household. For most purposes, this assumption has been proven to be

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powerful way of describing actual behavior, but in certain applications it is important to consider explicitly the multiperson nature of many households. This paper considers one such application: the analysis of a government policy intended to redistribute resources within a family.

The specific issue analyzed in this paper is a married couple's choice of the amount of survivor protection to be provided to a surviving spouse after the death of her partner.[1] The potential conflict of interest between spouses rises from the fact that providing protection to a surviving spouse is costly (e.g. life insurance is not free). This means that the more survivor protection is provided, the less resources the household has available in other states of the world. Thus there is the potential for conflicting interests between spouses.

The application studied in this paper is the spousal signature requirement of the Retirement Equity Act (REA) of 1984. This requirement mandated that a married pension plan participant, when retiring, must choose his pension payment in a form of a joint-and-1/2 survivor annuity unless his spouse signs a notarized consent form waiving her right to this survivor protection.[2] The mandate affected only pension plan participants who started receiving their pensions after January 1, 1985.

In the theoretical part of the paper, a Nash-bargaining model of family decision making is used to analyze the specific effects of this law change for the selection of survivor annuities and life insurance holdings. The law change is interpreted as having changed spouses' relative outside options. The model predicts that the law change would increase the selection of the survivor annuities and increase life insurance holdings for most households. These predictions of the Nash-bargaining model are contrasted with the stark prediction of the classical model that the law would have had no effect since the household budget set is unchanged. Thus this exogenous law change provides a well-identified empirical strategy for testing the predictions of the bargaining model against the predictions of the classical model.[3]

In the empirical part of the paper, several cross-sectional datasets are used to study these predictions. The effect on the survivor annuity selection is studied using the Current Population Survey (CPS) December 1989 Pension Benefit Survey and a combination of

[1] From now on, we will use convention that the husband is the spouse who, having been the primary earner, is more likely to die earlier. While the reverse situation is relevant for some couples, this is still (especially for the cohorts used in the empirical analysis) overwhelmingly more typical. The law change that is studied in this paper, while written in gender-neutral terms, was explicitly targeted to increase the protection of widows after the death of their husbands.

[2] A joint-and-1/2 survivor annuity is an annuity that pays a fixed income stream as long as the primary annuitant (the pension plan participant) is alive and 50% of this stream as a survivor benefit for his spouse after his death as long as she is alive. A typical alternative to the survivor annuity is a single life annuity that pays a higher fixed income stream during the participant's lifetime. The terms "joint annuity" and "survivor annuity" are used interchangeably in this paper.

[3] Most of the existing literature that tries to test between alternative models of household behavior use as their identification sources variables that could easily be interpreted as being endogenous to the decision (like the relative income shares of the husband and wife). Thus, the rejections of the classical model in these papers can be due to this problem of identification strategy. Two exception are Duflo (2000) and Lundberg et al. (1996). In the former, the natural experiment was an expansion of pension benefits in South Africa. In the latter, the natural experiment was a policy change in the UK, which changed the Child Benefit from tax credits to a direct payment to the mother. Both papers reject the classical single utility function view of the household.

the Health and Retirement Survey (HRS) and the Assets and Health Dynamics Among the Oldest Old (AHEAD).[4] These results show that the law change increased the selection of survivor annuities by approximately 7 percentage points (a 10% increase). Results from HRS-AHEAD data indicate that the median increase in life insurance holdings for affected households was approximately US$5000. This corresponds to approximately 25% of median life insurance holdings of the affected group. These joint annuitization and life insurance findings support the Nash-bargaining theory over the classical single-utility maximization model.

2. Survivor protection: legislation and economic evidence

Protection of surviving spouses can be provided by several instruments: privately purchased annuities, survivor annuities from private pensions, public pensions (Social Security), life insurance and savings. It is worth noting that many of these instruments can used for motives other than survivor protection. Several authors have argued that bequest motives are important explanations for wealth accumulation (savings behavior) and for life insurance holdings (e.g. Bernheim, 1991; Brown, 1999; Kotlikoff, 1988). Most households rely substantially on Social Security, which provides a real joint annuity for married retirees. The surviving spouse in a typical married couple receives between one half and two thirds of the couple's Social Security benefits, depending on the spouses' relative earnings histories.

Prior to REA, all private sector and union pension plans in the US were affected by the Employment Retirement Income Security Act (ERISA) of 1974. ERISA required that if the pension plan's primary form of pension payout was an annuity, then the default option for married participants must be a joint-and-1/2 survivor annuity.[5] Pension plan participants were free to choose other payout options without consulting their spouses. Holden and Nicholson (1998), using New Beneficiary Survey data, show that ERISA increased selection of survivor annuities by married male pension plan participants from 48.1% to 63.9%. Unfortunately, these data cannot be used to disentangle the two effects of ERISA: the mandate that survivor benefits must be an option (increased availability) and the effect of the default choice.[6]

The Retirement Equity Act (REA) of 1984 was a major revision of the original ERISA legislation. It included two provisions that were explicitly meant to redistribute resources within a family. It mandated the provision of pre- and post-retirement survivor annuities

[4] When used together, these datasets will be referred as HRS–AHEAD data. Preliminary release data from HRS wave 1998 is used and therefore the following disclaimer applies. "This analysis uses HRS Preliminary Release data. These data have not been cleaned and may contain errors that will be corrected in the final Public Release version of the dataset."

[5] Before ERISA pension plans were not required to provide survivor annuities.

[6] A recent paper by Madrian and Shea (2001) provides evidence on the effect of the default choice on the investment decision made by 401(k) plan participants. They find that the choice of default option has a significant effect on retirement related investment decisions.

unless the spouse affected signed a consent form in the presence of a notary public or a pension plan administrator.

This paper will focus on the post-retirement survivor annuities mandate. This requirement specified that employers must provide married participants with a notice form explaining the choice (typically between a single life annuity and a joint-and-survivor annuity; in some cases lump-sum payment is also offered) and the rights of the parties involved at least 90 days before start of the pension payments. The mandated default form of joint-and-survivor annuity provided 50% of the benefit received when the participant was alive to his spouse after his death. This requirement affected all defined-benefit plans and most defined-contribution plans.[7]

For a typical retiring worker, the effect of selecting a joint-and-survivor annuity over a single life annuity is a reduction in pension benefits of approximately 10% (based on TIAA-CREF annuity pricing table, from TIAA-CREF, 1996). Pensions where the payments had started before January 1, 1985 were unaffected by these survivor annuity requirements.

3. Theoretical models of household decision making and the Retirement Equity Act

Three simple models of household decision making are compared in this section with respect to their predictions on the effects of the signature requirement of the Retirement Equity Act. The models are the classical single-utility-function model, an "almost dictatorial" model and a Nash-bargaining model. Because each of these three models gives different predictions regarding the effects of the signature requirement, they can be tested empirically.

3.1. The economic environment

There are two periods in the model. In the first period, both spouses are alive with probability one. In the second period, the husband is alive with probability $(1-p)$ and the wife is alive with probability 1. Period 1 in the model presents early retirement and period 2 late retirement. In the first period, the household must decide how much of its endowment to consume now, and how much to allocate to different states of the world in period 2. A complete market structure is assumed, so the couple can freely trade consumption between first period and two possible states in the second period with no short-selling constraints. The analysis here is presented in terms of the choice between first period consumption, purchase of single-life annuities and the purchase of life insurance.

[7] Among defined-contribution plans, all the money-purchase pension plans were affected by the provision. Under certain limited circumstances, profit-sharing and stock-bonus plans were not affected (Schechter, 1985). It is also worthwhile to notice that the defined benefit or defined contribution plans that did not provide the option to annuitize the pension wealth through the plan were not affected by this requirement. This means that many DC plans currently gaining popularity (like 401k's) are not affected by this requirement. State and local government pension plans were not affected by this law while the federal government pension plan had a similar requirement change effective at same time thanks to the Civil Service Retirement Spouse Equity Act of 1984.

This choice is motivated by the fact that these are the Arrow securities for the relevant states. The discussion relating to the models links this back to joint life annuity choice.[8]

The utility functions of household members (or the household utility function in the case of the classical model) are assumed to be of the time-separable expected utility form. Household members have no bequest motives and the household is assumed to have an exogenously given endowment W.

3.2. Classical single-utility-function maximization

In this case, the household maximizes

$$U(c_1, c_2, \tilde{c}_2) = u(c_1) + (1-p)v(c_2) + p\tilde{v}(\tilde{c}_2) \quad \text{s.t. } W = c_1 + qc_2 + \tilde{q}\tilde{c}_2, \tag{1}$$

where c_1 is the current period consumption, c_2 is the consumption when both spouses are alive (price q) and $\tilde{c}_2$ is the consumption in the widowhood state (price $\tilde{q}$). Note that the utility function over consumption in period 2 is allowed to be state-dependent.

This model makes a stark prediction with respect to the signature requirement: it should have absolutely no effect on the decisions that the household makes since the budget set remains unaffected.

3.3. "Almost dictatorial" model

A slight generalization of the classical model has a (possibly altruistic) husband making all the economic decisions in the family, while the wife has her own utility function (which is irrelevant to the household's decisions). So while she has her own preferences, we observe only the choices made according to husband's preferences. Since the husband has ample tools in this model to undo any increase in joint annuitization by canceling life insurance (or selling it short), he may choose not to get his wife's signature to forego the survivor annuity. Instead he can completely offset this increase in survivor protection by canceling his life insurance (or by short-selling life insurance). This simple model provides justification for the investigation, in the empirical part of the paper, of possible offsetting behavior on the life insurance margin.

3.4. Bargaining model

The Nash bargaining models (Manser and Brown, 1980; McElroy and Horney, 1981) and, more generally, efficient contracting models (Chiappori, 1988) of household behavior have been a topic of research in several areas of economics in past 15 years. The applications of these models or other more general models of household decision making to retirement-related topics are rare, two notable exceptions being Browning (2000) on the

[8] In this environment, the same state space is spanned also by life insurance and risk-free bonds or by single- and joint-life annuities. In the real world, the situation is more complicated, due to the lumpiness of the pension annuity selection, differences in the pricing of annuities and life insurance and rationing/non-linear pricing in the life insurance market.

theory side and Lundberg and Ward-Batts (2000) on the empirical side.[9] The basic tenets of these models are that households have at least two decision makers with separate utility functions, and that the choices households make are Pareto-efficient.

This section presents a Nash-bargaining model, but the model presented can easily be understood to be just a special case of a more general efficient contracting model. All the results presented continue to hold in these more general models. In the Nash-bargaining model, both spouses are assumed to have utility functions over their own consumptions in different states of the world. The spouses in the model make a binding decision in period 1 that is honored also in all states of period 2.

The key determinant of the Nash-bargaining solution is the outside option. This is defined as the utility level that the agent would attain should negotiations break down. In most of the Nash-bargaining literature on family decision making, the outside options are considered to be the spouses's respective utility levels in the case of divorce, given the institutional arrangement on the sharing of household resources, as in the original McElroy and Horney (1981) contribution. Lundberg and Pollak (1993) introduced the notion of a non-cooperative marriage as the outside option. In the context of this application, the non-cooperative marriage is the preferred interpretation. A non-cooperative marriage is interpreted in this context as a situation in which both spouses separately consume the income streams over which they have property rights, and do not optimally divide household chores. Although household chores are not explicitly modelled here, they represent one of the wife's bargaining chips: the threat of not providing household services to the husband is a potential instrument in her bargaining strategy.

3.4.1. Definition (Nash-bargaining problem)

Formally, the problem of the couple can be stated as

$$\max_{\bar{c}} \left(V^{\mathrm{f}}(c_1^{\mathrm{f}}, c_2^{\mathrm{f}}, \tilde{c}_2^{\mathrm{f}}) - h^{\mathrm{f}}\right)^{*}\left(V^{\mathrm{m}}(c_1^{\mathrm{m}}, c_2^{\mathrm{m}}) - h^{\mathrm{m}}\right)$$

$$\text{subject to } (c_1^{\mathrm{f}} + c_1^{\mathrm{m}}) + q(c_2^{\mathrm{f}} + c_2^{\mathrm{m}}) + \tilde{q}\tilde{c}_2^{\mathrm{f}} = W, \tag{2}$$

where V^{f} and V^{m} are the utility functions of the wife and husband, h^{f} and h^{m} are the outside options of the wife and husband, c_1, c_2 and $\hat{c}_2$ are, respectively, first-period consumption, second-period consumption in the state where husband is alive, and second-period consumption in the wife's widowhood state.

The respective utility functions of the spouses take form of time-separable expected utility:

$$V^{\mathrm{f}}(c_1^{\mathrm{f}}, c_2^{\mathrm{f}}, \tilde{c}_2^{\mathrm{f}}) = u^{\mathrm{f}}(c_1^{\mathrm{f}}) + (1-p)v^{\mathrm{f}}(c_2^{\mathrm{f}}) + p\tilde{v}^{f}(\tilde{c}_2^{\mathrm{f}}) \text{ and } V^{\mathrm{m}}(c_1^{\mathrm{m}}, c_2^{\mathrm{m}})$$
$$= u^{\mathrm{m}}(c_1^{\mathrm{m}}) + (1-p)v^{\mathrm{m}}(c_2^{\mathrm{m}}). \tag{3}$$

[9] Browning (2000) models the decisions similar as studied in this paper as a non-cooperative game. Under the assumptions used, he finds that the Nash-equilibrium of the game can be Pareto-efficient. Lundberg and Ward-Batts (2000), on the other hand, find that variables plausibly correlated with the respective bargaining powers of the spouses (such as spouses' respective education levels and age difference between spouses) affect the net worth of households in the first wave of the HRS.

In all proofs below, we assume that utility functions are concave and twice differentiable and that there is some marital surplus to be shared in the optimum (so the outside options do not bind). We also assume that outside options h^{f} and h^{m} are functions of the amount of wealth that the each spouse commands should the negotiations break down. We assume that fraction α of the total wealth of the couple is commanded by the wife, so the husband's and wife's wealth in the outside option situation are $(1-\alpha)W$ and αW, respectively.

In the bargaining context, the signature requirement of the Retirement Equity Act changes the relative outside options of the spouses by redistributing property rights on the income stream provided by the survivor annuity to the wife. Before the requirement, she does not have a claim on that income stream. This is equivalent to a redistribution from the husband's outside option to the wife's outside option, which in the context of our model is equivalent to increasing α. This interpretation of REA does not have to be taken literally in what follows, it suffices that the relative outside option of the spouses are affected by the Retirement Equity Act.

Result 1. REA increases the utility of wife and decreases the utility of husband.

This result is a direct consequence of the redistribution of outside options by REA. Since her outside option is higher when REA is in effect, her utility in the Nash-bargaining solution will be higher. Moreover, this result holds whether or not the household would have chosen the survivor annuity without REA. This is a general property of the standard Nash-bargaining solution: outside options always matter to the solution.

Result 2. REA increases the amount of money transferred to the survivor state (the sum of survivor annuities and life insurance) and increases the wife's private consumption in periods 1 and 2.

Proof. First-order conditions for the problem (2) yield:

$$u^{\mathrm{f}\prime} = \frac{v^{\mathrm{f}\prime}}{q} = \frac{\tilde{v}^{\mathrm{f}\prime}}{\tilde{q}} = \lambda u^{\mathrm{m}\prime} = \lambda \frac{v^{\mathrm{m}\prime}}{q},$$

where

$$\lambda = \frac{V^{\mathrm{m}}(c_1^{\mathrm{m}}, c_2^{\mathrm{m}}) - h^{\mathrm{m}}}{V^{\mathrm{f}}(c_1^{\mathrm{f}}, c_2^{\mathrm{f}}, \tilde{c}_2^{\mathrm{f}}) - h^{\mathrm{f}}}.$$

λ is decreased by REA due to the Result 1. For concave u and v functions and for an unchanged budget set, a decrease in λ will unambiguously increase c_1^{f}, c_2^{f} and $\hat{c}_2^{\mathrm{f}}$ and also decrease c_1^{m} and c_2^{m}. □

As the bargaining power is tilted towards the wife with REA, the family will consume more items that enter positively into her utility function. Since life insurance holdings and survivor annuities are perfect substitutes in this model, the prediction is only on the sum of these two.

In practice the choice of survivor annuity is a discrete choice, the possible choices typically being no survivor benefits in the annuity and some selected levels of survivor annuity (say, 50% or 100% survivor annuity). Given that we have an unambiguous prediction on the sum of the life insurance and survivor annuities, this allows us to make unambiguous comparative static predictions on the life insurance holdings of the households who would not change their discrete survivor annuity choice when moved from no-REA environment to REA environment. For these households, the life insurance holdings should go up. For the group that would change their discrete choice survivor annuity choice between the two environment, the prediction on life insurance is ambiguous.[10]

4. Empirical analysis

The predictions of the Nash-bargaining model are tested against the predictions of the classical single-utility maximization model and the "almost dictatorial" model in this section. Note that the goal here is modest, we are not trying to estimate any structural parameters of the Nash-bargaining model needed for, e.g. normative analysis of the policy. Instead our interest is in the reduced-form predictions of the model.

Outcomes studied from cross-sections of married couples include survivor annuity choice and life insurance holdings. The identification strategy in these regressions is based on either the husband's birth year or the start-date of his pension. Where the data permit, households in which husband is not receiving pensions (and will not receive in the future) are used as a control group. This allows us to use both standard first-difference and difference-in-differences empirical strategies.

The datasets used in this section include Health and Retirement Survey and Assets (HRS), Health Dynamics Among the Oldest Old (AHEAD), Current Population Survey (CPS) December 1989 Pension Benefit Survey and CPS March files from several years. Evidence from published tables on the annuity selections of TIAA-CREF participants is also presented.

4.1. Outcomes, population studied and models estimated

The outcomes studied in this section are:

(1) The probability (conditional on the husband having a pension) that the husband's pension provides survivor benefits;
(2) The probability that the wife receives life insurance payments should her husband die;
(3) The amount of life insurance protection.

The classical single-utility-function model gives a stark prediction that the law should not have affected any of these outcomes. The Nash-bargaining model predicts an increase in joint annuitization. Furthermore, for most households, the Nash-bargaining model

[10] The reason that the effect on life insurance holdings of this group is ambiguous is that the model makes only a prediction about the sum of life insurance holdings and survivor annuity selection. By making appropriate assumptions about the form of the outside option functions and the utility functions, this effect can be made arbitrarily small or large.

predicts increase in the life insurance holdings (so also the second and third outcomes should increase due to the legislation).[11]

The population studied is married couples whose husband was born between 1916 and 1919 (old group) or between 1924 and 1931 (young group). The empirical justification for this comes from the yearly March CPS files from 1976 to 1998. The probability of a pension income receipt from federal or private pensions was calculated for married males as a function of their age. This relationship was relatively stable across time during the whole period. The approximate probability of receiving pension income at age 54 was 4%, 15% at age 62, 38% at age 66 and around 40% at age 70. This suggest that most of pension starts happen when husband is between 62 and 65. Furthermore, this means that the members of our old group was likely to have started their pensions before REA and the members of the young group were unlikely to have started their pensions before that.[12]

The receipt of pension income by the husband is the key quantity in the analyses.[13] When annuity choice is studied, only those who had pension income are included in the sample. When other outcomes are studied, the pension variable is used to divide the data into control and experimental groups. While using non-pension holders as a control group for pension holders is not an ideal control strategy, the inclusion of a rich set of covariates (like career high earnings of the spouses) should mitigate problems related to the differences between these groups.

For all outcomes, the following regression is estimated using only households where the husband had pension income:

$$Y = \alpha + \beta * \text{young} + \eta * Z + \varepsilon, \tag{4}$$

where Y is the dependent variable (either an indicator for survivor benefits from pension, or an indicator for having life insurance, or the amount of life insurance holdings), young is an indicator for the husband being young enough to be affected by the legislation and Z is the set of potential covariates.[14]

The model estimated is the standard first-difference model (where the difference is with respect to birth cohort of the husband) and where β is the estimated effect of the legislation.[15]

[11] That is, for households that did not change their annuity choice because of the law change. According to the estimates of this section, these households constitute approximately 93% of the households in which the husband has pension income.

[12] The sample (where the data allow this) was further restricted to married couples who were married when the husband turned 60 to eliminate couples who might have not been married at the time of the annuity selection. Due to relatively low number of couples marrying after age 60, this criterion has very little effect on the sample.

[13] The cohorts studied are sufficiently old in the data (at least 66) to be very likely to have already started their pensions, which is important for our purposes, since it implies that they have already made their annuity selection.

[14] The full covariate set in all these regressions includes non-linear controls for age differences between spouses, third-order polynomials in career high earnings of the spouses (including interaction terms), an indicator for wife having no work history, educational and race indicators and an indicator for the wife having (or expecting) pension of her own.

[15] Even though the dependent variable is in many cases a binary variable, these regressions were estimated using OLS. As a robustness check, they were also estimated as Probits. All the qualitative results of the analysis were unaffected by the choice between Probit and the linear probability model estimated by OLS.

The control group did not face the annuity selection, so they cannot be used as a control group for that outcome. For other outcomes, the following model using all the observations is estimated.

$$Y = \alpha + \beta_0 * \text{young} + \beta_1 * \text{pension} + \beta_2 * (\text{young} * \text{pension}) + \eta * Z + \varepsilon, \tag{5}$$

where pension is an indicator for husband having a pension and β_2 is the program effect. Here the full covariate set also includes indicators for the husband's birth year (since the identification is no longer based solely on the birth year).[16] This estimation strategy can be seen as a standard difference-in-differences strategy, where the first-difference is between cohorts and the second is between pension statuses of the husband.

An exception to the empirical strategies described above is the annuity selection equation estimated from the CPS 1989 Pension Benefit Survey.[17] There an indicator variable for the pension starting after 1985 was used instead of the cohort proxy. This specification allows the use of the husband's birth year in the covariate set.

4.2. Data sources

The Current Population Survey (CPS) December 1989 Pension Benefit Survey is a special supplement to the CPS collected for the purpose of analyzing the effects of the Retirement Equity Act. It includes detailed pension information (including the start date of benefits) and information on whether each pension provides survivor benefits. Because this survey was collected only 4 years after the law change, it is subject to fewer sample selection problems than other datasets such as HRS–AHEAD.[18] Ideally one would like to analyze a question like selection of survivor annuities from flow data. The downside of CPS 1989 dataset is that it does not include information on life insurance holdings, but it does include information on the career high earnings, which is a key covariate in the regressions.

The Health and Retirement Survey (HRS) and the Assets and Health Dynamics among Oldest of the Old (AHEAD) panels are the main sources of data in this paper. HRS and AHEAD started as separate panels in 1992 and 1993. AHEAD started as a panel survey

[16] When birth year indicators are included in the regression they replace the variable young. However, due to the limited sample size, the interaction terms with pension are not estimated separately for each birth year even in this case. Instead the interaction young×pension is still used.

[17] The described husband's birth cohort restrictions do not apply to this case either, since we do not have to rely on the birth cohort for identification with this data.

[18] The mechanism for sample selection is the following: suppose husbands have private information on their life expectancy and this information enters into the decision whether to select survivor annuity (with the ones more likely to die soon more prone to select survivor annuities). Then any cross-section attempt to estimate the effect of the start year (or birth cohort group) on annuity selection will be biased since the earlier the start year is (or the older the birth cohort is), the higher the proportion of those who chose survivor annuities because of the private negative information on their life expectancy and have died before reaching the data collection. This leads to a bias towards the finding that among later pension starter a higher fraction choose survivor annuities.

Table 1
Age of the compared cohorts in the first wave of AHEAD and in the HRS 98. The preferred comparison is between the off-diagonal cells of this table

Dataset	Age of the young	Age of the old
AHEAD 1	62–69 (not used in the analysis)	74–77
HRS 98	67–74	79–82

of households in which at least one of the members was over 70 years old at the time of the first interview (born in 1923 or earlier). HRS started as a panel of households having at least one member born between 1931 and 1941. Before 1998 there was one additional AHEAD wave (1995) and two additional HRS waves (1994 and 1996). In the HRS 1998, these two panels were merged and additional cohorts were included in the panel to have a representative sample of households were at least one member was born before 1947. It is important to note that most individuals who are young enough to be affected by the law change are only part of the HRS 1998 data. The advantage of HRS–AHEAD data is the detailed information on the work histories, pensions and life insurance and assets holdings.

The HRS–AHEAD data were used in two separate ways in the empirical analysis. A cross-sectional estimation of the effect of legislation uses the HRS 1998. However, the mortality bias for the older group could be significant in this approach, since the members of the older group would have to have lived to be 79–82 to be included in the sample. For this reason, the effects of the law change are also estimated using data for the older group from the first wave of AHEAD and data from HRS 98 for the younger group. Although not a perfect solution, this approach could reduce the mortality bias substantially. Ideally one would like to compare similarly aged individuals at different times, but this is not possible given the existing datasets.

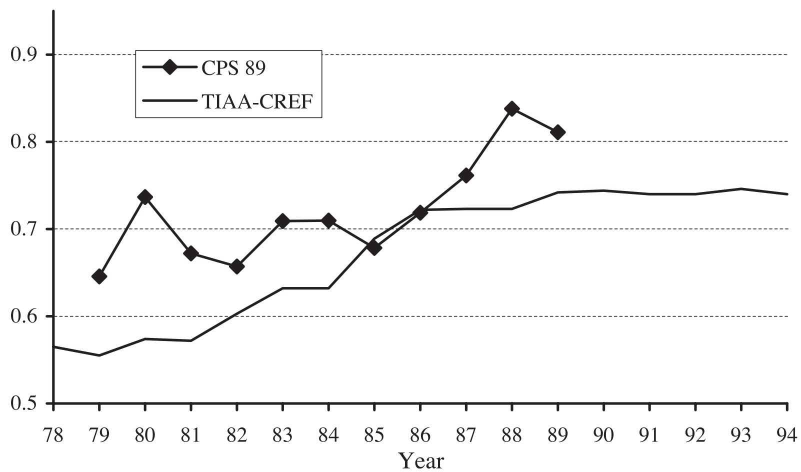

Fig. 1. Fraction of married males (CPS) and all males (TIAA-CREF) choosing survivor annuities.

Table 2
Results on probability that husband's pension provides survivor benefits (CPS 89 December Supplement)

Survivor Benefit Choice (CPS 89)			
Pension 1985 or after	0.069	0.070	0.063
	(0.025)**	(0.025)**	(0.025)*
Wife has pension			−0.081
			(0.038)*
N	1540	1540	1540
Demographics and education	No	Yes	Yes
Income variables	No	No	Yes

** and * indicate significance at 1% and 5% confidence levels, respectively. Standard errors are in the parentheses.

The ages of the older and younger group at the times of different surveys is reported in Table 1.

4.3. Results

4.3.1. Survivor annuity selection

Evidence on annuity selection are presented from three different data sources. The first data source is a published table of annuity choices by TIAA-CREF participants (TIAA-CREF, 1996), where the data are tabulated according to the start year of the pension. The other data sources are the CPS 1989 December Supplement and the HRS–AHEAD data, described above.

The evidence from TIAA-CREF flow data is presented in Fig. 1. Between 1978 and 1994 selection of survivor annuities went from 56.5% to 74%, an increase of 17.5 percentage points. More than half of the total change (9 percentage points) occurred between 1984 and 1986. This suggest that while there was an pre-existing trend in the data, the legislation had a substantial effect on the selection of survivor annuities.[19]

The estimates from CPS 1989 December Supplement are presented in Table 2, and in Fig. 1. Across different specifications these results suggest that the selection of survivor annuities went up by 7 percentage points after the law change.[20] This result is robust to the choice of covariate set and is statistically significant.

[19] Two caveats are in order here: the data includes also non-married participants and the workers from state universities (and certain church-run universities) who were not affected by the legislation.

[20] The CPS 1989 December Supplement was also used in a General Accounting Office (GAO) Report studying the effects of REA on the selection of survivor annuities(United States General Accounting Office, 1992). Based on simple tabulation, the GAO estimated that the selection of survivor annuities increased by 15 percentage points after the legislation. Two choices explain the differences in the simple REA program effect between the current study and the GAO study: (1) GAO study used all observations in the analysis regardless of how long ago the selection was made while here we restrict to more recent observations and (2) GAO study uses as the after-REA probability the highest individual year after REA while we use the average over all years after REA.

Table 3

Results on the probability that husband's pension provides survivor benefits (from HRS–AHEAD)

	Across 1993 and 1998			Cross-section in 1998		
Husband young	0.048	0.082	0.077	0.047	0.091	0.101
	(0.038)	(0.043)	(0.047)	(0.044)	(0.048)	(0.050)*
Wife pension			−0.101			−0.082
			(0.039)*			(0.040)*
N	1001	1001	999	913	913	911
Demographics and education	No	Yes	Yes	No	Yes	Yes
Income variables	No	No	Yes	No	No	Yes

Probability that husband's pension provides survivor benefits

Age in 1985		
54–61	69.1%	
66–69	64.3%	from 1993 data
66–69	64.5%	from 1998 data

** and * indicate significance at 1% and 5% confidence levels, respectively. Standard errors are in parentheses.

The underlying identifying assumption for these estimates from CPS 89 to be the effect of REA that must hold is that the selection of these benefits would have not changed across time for other reasons than REA. Given the short time-series dimension of the data available, it is impossible to completely rule out existing time trends towards more annuitization (see Fig. 1), but for the CPS 89 data, there seems to be solid break in the pattern around the law change.[21]

The results from the HRS–AHEAD data are presented in Table 3. These estimates suggest that signature requirement increased the selection of survivor annuities by 5–10 percentage points. These results range from statistically insignificant to significant at the 5% level depending on the data used and the specification estimated.

Among the covariates, it is interesting to note that if the wife has a pension on her own or is expecting a pension, her husband is less likely to provide survivor benefits through his pension. This holds for both datasets and is statistically significant in both datasets.

[21] One potential confounding factor is that that REA could have increased the take-up of lump sum from pension plans and that the estimates here could be biased because of that. To address this concern, the analysis of CPS 89 data was also performed where everyone who ever had taken a lump sum of more than US$3000 and who did not have survivor benefits in a current pension was included in the no-survivor-benefits category. The estimated program effect was approximately 1% smaller and the statistical significance of the effect was unchanged. A direct test of whether the tendency to have taken any lump sum payments over US$3000 increased after 1985 lead to an insignificant 1% effect of REA. Given these results, the increased lump sum take-up cannot be the factor driving the results in the HRS and AHEAD portion of the empirical analysis where the data does not allow controlling for lump sum take-up.

Table 4
Results on life insurance holding probability (HRS–AHEAD)

	Across 1993 and 1998				1998 only			
	D	DD	DD	DD	D	DD	DD	DD
Husband young	0.081	0.111	n.a.	n.a.	0.036	0.057	n.a	n.a
	(0.032)*	(0.041)**			(0.035)	(0.047)		
Husband pension		0.156	0.129	0.126		0.147	0.128	0.116
		(0.047)**	(0.046)**	(0.047)**		(0.054)**	(0.054)*	(0.056)*
Young×Pension		−0.030	−0.018	−0.020		−0.021	−0.014	−0.007
		(0.052)	(0.051)	(0.052)		(0.059)	(0.058)	(0.060)
Wife pension				−0.018				−0.004
				(0.026)				(0.025)
N	1039	1952	1951	1937	951	1775	1775	1761
Demographics and education	No	No	Yes	Yes	No	No	Yes	Yes
Income	No	No	No	Yes	No	No	No	Yes

Life insurance holding probabilities

	1998	1993	1998
% Insured	Young (%)	Old (%)	Old (%)
Pension	84.23	76.15	80.63
No pension	71.64	60.58	65.98

Columns denoted by D refer to the first-difference model (Eq. (4)), where the model is estimated only for households where the husband has pension. DD refers to the difference-in-differences model (Eq. (5)), where the non-pension households are used as a control group.
** and * indicate significance at 1% and 5% confidence levels, respectively. Standard errors are in parentheses.

These results on increased joint annuitization are consistent with the bargaining model and with the "almost dictatorial"model. They constitute a rejection of the standard single-utility-function model of the household. However, the magnitude of estimated effect is not very large. This is natural given that already before the legislation a majority of married husband were choosing survivor benefits (approximately 70%).[22]

4.3.2. Probability of having life insurance protection for the wife

From HRS–AHEAD, first-difference and difference-in-differences models were estimated for the probability that the husband has life insurance policy in which the wife is listed among the beneficiaries. With the exception of the first-difference estimator across HRS and AHEAD data (in which the old group is from AHEAD Wave 1), no significant effect on this margin is found. The point estimates from difference-in-difference models are slightly negative, but given their standard errors, they are consistent also with large positive effects. The results are presented in Table 4.

[22] In the HRS–AHEAD data, the proportion selecting survivor benefits is on the average approximately 5 percentage points lower than in CPS 1989. One possible interpretation for at least some of this difference is that HRS–AHEAD data, having been collected later than CPS 1989, suffers more from the mortality bias.

Table 5
Life insurance amount results using 1993 and 1998 cross-sections (HRS–AHEAD)

Across 1993 and 1998

	OLS				Median regression				Robust regression			
	D	DD	DD	DD	D	DD	DD	DD	D	DD	DD	DD
Husband young	37,185	24,956	n.a.	n.a.	15,660	10,773	n.a.	n.a.	12,031	3739	n.a.	n.a.
	(5906)**	(9661)**			(842)**	(959)**			(1582)**	(1148)**		
Husband pension		−9896	−14,727	−12,816		2000	−39	−661		1648	269	−154
		(8524)	(7989)	(8825)		(1118)	(2130)	(1107)		(1350)	(1570)	(1638)
Young×Pension		12,229	14,155	13,478		4887	6237	4846		3684	4422	4383
		(11,323)	(10,305)	(11,130)		(1283)**	(2444)*	(1273)**		(1563)*	(1813)*	(1890)*
Wife pension				−12,774				84				−224
				(5716)*				(715)				(1097)
N	912	1714	1713	1700	912	1714	1713	1700	912	1714	1713	1700

1998 cross-section

	OLS				Median regression				Robust regression			
	D	DD	DD	DD	D	DD	DD	DD	D	DD	DD	DD
Husband young	28,203	−585	n.a.	n.a.	12,396	6887	n.a.	n.a.	6874	2196	n.a.	n.a.
	(6760)**	(17,466)			(2954)**	(1416)**			(2345)**	(1823)		
Husband pension		−26,455	−35,323	−33,757		1377	−0.00	−746		3715	2543	1878
		(17,180)	(18,879)	(185,581)		(1797)	(1925)	(3210)		(2295)	(2574)	(2619)
Young×Pension		28,788	34,128	33,959		5509	6405	4589		2461	2738	2585
		(18,728)	(19,326)	(19,137)		(1977)**	(2120)**	(3521)		(2523)	(2825)	(2860)
Wife pension				−13,950				1570				−194
				(7185)				(1753)				(1428)
N	814	1545	1545	1532	814	1545	1545	1532	814	1545	1545	1532
Demographics and education	No	No	Yes	Yes	No	No	Yes	Yes	No	No	Yes	Yes
Income variables	No	No	No	Yes	No	No	No	Yes	No	No	No	Yes

Insurance holdings	Average			Median		
	Young	Old (1993)	Old (1998)	Young	Old (1993)	Old (1998)
Pension	52,374	15,189	24,172	20,660	5000	8264
No pension	50,041	25,085	50,626	13,773	3000	6887

** and * indicate significance at 1% and 5% confidence levels, respectively. Standard errors are in parentheses.

4.3.3. The amount of life insurance protection

From HRS–AHEAD, both first-difference and difference-in-differences models were estimated for the value of life insurance protection of life insurance plans where the wife is listed as a beneficiary (note that zero values were included since that is a valid amount of life insurance protection).

Results from three different statistical models were estimated: standard linear regression by OLS, median regression and a version of robust regression that uses biweight-weighting scheme to downweight outliers (Hamilton, 1991). The latter two models are estimated to ensure that the results are not driven by small number of outliers. These results are presented in Table 5.

The results for life insurance protection imply that the law change increased it by approximately US$5000 (median regression estimate), which would correspond roughly to 10 times the husband's median monthly pension. The magnitude and significance of this estimate varies across different models, the OLS estimates being significantly larger and the robust regression results being smaller than those of the median regression. The estimated effect of REA is statistically significant for all specifications and samples with one exception for median regression, for robust regression they were significant when the old group data is from AHEAD 93 and never for OLS. The overall pattern of statistical significance across provides evidence that REA increased the amount of life insurance holdings.[23]

The required identifying assumption here (and in the life insurance probability regressions) is that any difference in the time patterns across the two groups is due to REA and not existing differential trends across groups. Using AHEAD, the comparison between the old in our dataset and even older cohorts (those 70–73 year olds in 1985) found some evidence on existing differential trends in life insurance holding probability and statistically significant evidence on differential trends in life insurance holdings across the pension and no-pension households. However, both of these differential trends were in the opposite direction to the now estimated effects of REA on life insurance holdings, so in this sense this makes the effects of REA even stronger.

This result of increased life insurance protection is consistent with the Nash-bargaining model of household behavior and inconsistent with both the classical model and the "almost dictatorial model."

[23] Tobit models were also estimated to explicitly account for the censoring of life insurance at zero. The estimated Tobit treatment effect were always between the OLS and median regression and they were statistically significant when the data from both groups was from HRS 98. Regular regression estimates for log of insurance holdings (where the zero values were omitted from sample) imply approximately 4% increase in life insurance holdings and this effect was almost always statistically significant across specifications and samples. A two-step selection model for log insurance (where the binary selection equation had same covariates as the main equation) implied similarly a 4% increase (the coefficient on the main equation) and this effect was statistically significant in most specifications.

5. Discussion and conclusion

The US spousal consent signature requirement studied here is not the only one of its kind that has been proposed and or implemented. Similar requirements have been implemented in Canada and have been proposed in the UK. In Canada, several provinces and the federal government (which oversees some private sector pension plans) passed a requirement very similar to REA to apply to private sector pension plans between 1987 and 1992 (Federal and Provincial Laws of Canada).[24] In the UK, the direction of reform is more uncertain. Currently, occupational defined benefit pension plans provide survivor benefits as a condition of contracting-out of the State Earnings Related Pension Scheme. An independent government commissioned report called for elimination of this mandate ("the Pickering Report"; Pickering, 2002). On the other hand, the Equal Opportunities Commission has called for a mandate similar to REA to apply even to annuities purchased from other types of pension plans (Equal Opportunities Commission, 2004). Under current law, the primary annuitant can freely choose a single life annuity from these plans; only 19% married individuals choose survivor annuities (Association of British Insurers, 2002). In the US, extensions of REA to require spousal consent signature on all defined contribution pension plan payments have also been suggested (United States Senate, 2004), but no legislation has been passed.

To summarize the paper, a tractable Nash-bargaining model for household decision making over survivor protection was presented. The model made two specific predictions regarding the effects of the Retirement Equity Act on the choices that households make: the selection of joint annuities would increase, and, for most households, life insurance holdings would increase. These predictions are in stark contrast with the predictions of the classical single-utility-function maximization model. Using several microdata sources, it is shown that the predictions of the Nash-bargaining model are confirmed. This constitutes a rejection of the single-utility maximization model of household behavior in this decision-making realm.

These results imply that the change in the selection of survivor annuities were not the only effect of the Retirement Equity Act. The increase in life insurance holdings through the household bargaining mechanism, while increasing income security for widows, was neither foreseen nor intended by the legislation. In this, there is an important lesson for policy making that targets the resource allocation within a family. Because we do not yet fully understand the decision-making dynamics in the family, policies can have unanticipated effects due to the household decision-making process. The model and the empirical results presented here take the literature one step closer to understanding this process and its implications.

[24] Studying the issues presented here also with Canadian data could be fruitful, given that the different timing of these reforms could yield an empirically clean identification strategy.

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