Social Insurance and the Marriage Market*

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Abstract

When social insurance eligibility depends on marital status, this is a government intervention into the marriage market. I formally show that such intervention influences three behavioral margins in the marriage market, and test the theory exploiting a Swedish reform of *survivors insurance* – an annuity paid to widows, but not divorcees, upon the husband's death. First, I analyze bunching in the distribution of marriages and show that, by affecting the wedge between marriage and cohabitation, survivors insurance alters the composition of married couples up to 45 years before the annuity's expected payout. This distortion is larger in couples with higher *ex post* male mortality, holding constant the policy's value at realization and all demographics that I observe, suggesting "adverse selection" into government-provided insurance. Second, I use a regression discontinuity design to show that removal of survivors insurance from existing marriage contracts caused divorces and, in surviving unions, a renegotiation of marital surplus. Third, because survivors insurance subsidized couples with highly unequal earnings (capacities), its elimination raised the long-run assortativeness of matching. I argue that such marriage market responses to social insurance design have important implications for when it is optimal to separate social insurance from marriage in modern societies.

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1 Introduction

A major function of governments in many developed countries is to provide social insurance. The two largest social insurance programs in the United States, Social Security and Medicare, together account for more than 30 percent of federal spending. It is well recognized that the provision of social insurance to protect against adverse income or health shocks distorts markets that offer private insurance against such shocks.

I instead focus on responses in the marriage market. I do this because social insurance often represents a *twofold* intervention, both into private insurance markets and into the marriage market. This occurs whenever marital status influences eligibility for social insurance. In the United States, for example, both Social Security and Medicare fit into this category. I first ask how a link between social insurance and marriage affects the marriage market, and then examine the implications of my findings for when it is optimal to separate social insurance from marriage.

Specifically, in the context of Sweden, I study a particular type of social insurance, *survivors insurance.* Survivors insurance replaces part of the income that a household loses upon the death of one household member. Like many countries do today, Sweden used to provide survivors insurance through the marriage contract. A widow was granted a lifetime annuity of survivors benefits upon her husband's death, but cohabitating partners or divorcees were not eligible. The value of this insurance was significant; beneficiaries' average annual payout was ~\$5000, for an average duration of eight years (Section 2 provides details). But to most couples, who entered marriage in their 20s or 30s, the insurance was not likely to payout until far in the future. Marriage market responses to survivors insurance thus necessitate that couples have a long financial-planning horizon. I ask how this twofold intervention, into annuities markets and into the marriage market, affected the *volume* and *nature* of private contracting in the marriage market in Sweden. The volume of contracting is determined by entry into and exit from marriage. The nature of private contracts has a range of dimensions, of which I study two: Who contracts with — that is, marries — whom? And, which contracting party gets what share of joint realized gains?

In my empirical analysis, I exploit a 1989 reform that changed how survivors insurance was tied to the marriage contract in Sweden. The reform essentially eliminated survivors insurance, replacing the promise of a lifetime annuity with a promise of one (small) "adjustment transfer." Thus, the old marriage contract, which came with a government-provided annuity that was expected to pay out in old age, was replaced with a new marriage contract that came without it, but that otherwise was identical.

Section 3 theoretically analyzes how this affects the marriage market when individuals are forward-looking. My framework captures typical relationship dynamics. Specifically, single individuals form couples; then, in each period, each couple gets new information about (shocks to) its match-specific marital surplus ("happiness") and updates its marital decision. Formally, I introduce matching into a collective household framework and model match quality as a stochastic process. In this framework, the impact of a change in survivors insurance tied to marriage depends on whether an individual, *at the time of the reform's announcement*, is married, cohabitating, or single (to be matched in the future). For each group, the theory delivers precise, distinct, testable predictions about initial responses and about future behavior conditional on these initial responses. I test these using individual-level marital and tax records, described in Section 4. Some predictions vary depending on the properties of the stochastic match quality process. This also yields an implicit test of a largely open question about cohabitation, namely, whether cohabitation is a learning process in transition to marriage or an "end-state" substitute for marriage. My results are consistent with a match quality process that is stochastic and exhibits positive autocorrelation; this supports the interpretation of cohabitation as a learning process.

My first set of results, presented in Section 5, concerns how tying social insurance to marriage affects entry into marriage. Couples with a joint child who were not yet married at the reform's announcement in June 1988 were allowed to take up survivors insurance by marrying by December 31, 1989. By analyzing bunching in their distribution of new marriages, I can study how selection into marriage responded to a demand for survivors insurance. The distribution of new marriages displays a marriage boom in the last quarter of 1989. I estimate that a couple is on average 17 times more likely to marry in this quarter relative to the counterfactual scenario. Even couples below the age of 30 exhibit significant responses, implying financial-planning horizons as long as 45 years. Consistent with the theoretical prediction that couples in which the male is more likely to die should exhibit stronger responses, I document larger increases in take-up of marriage – and thus, of survivors insurance – among couples for which the husband's *ex post* mortality is higher. This remains true when controlling for a range of observable factors that determine the government's expected cost of extending survivors insurance to a given couple, including the husband's age at marriage. The positive correlation between couples' take-up of insurance through marriage and couples' expected cost of coverage suggests the possible presence of asymmetric information, which could pose a challenge to private markets for annuities in Sweden.

Further, I find that the reform's impact on take-up of marriage is positive but ten times *lower* in a sample of couples in which one spouse reveals a same-sex preference years later, by entering a same-sex union after they were legalized in 1995. If sexual preferences exhibit positive autocorrelation, this is consistent with the theoretical prediction that responses would be weaker in couples with weaker *ex ante* beliefs of staying together for life — the only circumstance in which the annuity would pay out.

When match quality is stochastic, the theory suggests that the reform should affect who gets married. Specifically, the marriage boom should consist both of "retimed" marriages that would have eventually taken place even in the absence of reform, and "extra" marriages that would never have occurred if the old marriage contract had remained available. Intuitively, announcing a replacement of the existing marriage contract with a less desirable one attaches option value to marrying into the desirable contract. Couples then rush to marriage even though they, in the absence of reform, would wait to see if the relationship improves. To decompose the excess mass in the bunching region into responses along the "intensive" (retiming) and "extensive" (extra marriages) margins, I develop an empirical framework that exploits the shape of the density, but also a second dimension of information: the date of birth of a couple's first child. I couple this with specific assumptions about the impact of this second dimension of information on marital behavior. Extra marriages accounted for more than a quarter of the response. Inherent in the theoretical option value explanation for this is a prediction that marriages in the boom are of a lower average match quality. Indeed, I show that such marriages are four percentage points more likely to end within 15 years than other marriages with the *same* contract. Nevertheless, a sizable portion of extra marriages survives in the long run.

I then analyze the causal impact of (losing) survivors insurance on exit from marriage. In Section 6, I study couples that were already married at the reform's announcement, but for whom grandfathering rules induced variation in insurance tied to marriage. Specifically, couples that married before January 1, 1985 were allowed to keep the contract they married into; for most couples that married thereafter, this contract was revoked and replaced with the new one. This change was announced in June 1988—three and a half years after entry into marriage, rendering impossible any manipulation in response to a demand for survivors insurance. Using a regression discontinuity design that exploits both the eligibility cutoff and the timing of the reform's announcement, I show that the removal of survivors insurance from preexisting marriage contracts raised the long-run divorce rate among these couples by six percentage points.

My second set of results concerns how tying social insurance to marriage affects the nature of contracting in the marriage market; specifically, the division of joint marital gains, and matching. I begin by studying couples' division of surplus. Among couples that married into the old contract but then lost insurance coverage, the theory predicts a renegotiation of marital surplus in favor of the wife in surviving unions. Intuitively, the wife's expected utility from marriage is the sum of her utilities in the states of the world where her husband is alive and where he is dead. The reform reduces her utility in the latter state; as compensation for this loss, her share of surplus in the former increases. I test this empirically by studying spousal labor supply, using a regression discontinuity design similar to the one described above. My results suggest that revocation of survivors insurance delayed elderly husbands' retirement – even though it affected household wealth only in a state of the world in which the husband would not be alive. This suggests that while the reform induced a statutory loss solely on the wife, intra-household bargaining resulted in the economic incidence partly being borne by the husband, who gave up utility – that is, leisure – when he was alive.

Finally, Section 7 analyzes long-term impacts on matching behavior. Because the annuity replaced household income that was lost due to the husband's death, payments were higher in couples with more spousal specialization in market and non-market work. Survivors insurance thus constituted a public subsidy on matches with highly unequal earnings (capacities). I theoretically show that under a standard supermodularity assumption on marital surplus, assortative matching arises in the absence of survivors insurance; however, when survivors benefits favor one-sided unassortative matches, assortativeness may break down. In a nutshell, removing survivors insurance should induce a larger share of skilled men to match with skilled women. To test this, I study the density of the share of highly skilled men that marry a woman of lower skill. I show that the share of newly married couples that are assortatively matched on education increases by four percentage points following the introduction of the new marriage contract; further, this increase is permanent. This suggests that the old survivors insurance program promoted unassortative matching and spousal specialization in market and nonmarket work.

These findings establish that tying social insurance to marriage has a significant impact on the volume and nature of contracting in the marriage market. In Section 8, I formally examine the implications of these findings for a social planner who wishes to alleviate old age poverty. In the presence of such marriage market responses, the social planner faces a trade-off between, on the one hand, protecting women who do not participate in the labor market against poverty at the end of life, and distorting marriage market behavior, on the other. The gender wage gap is a key determinant for whether it is optimal, in a given society, to separate social insurance from marriage.

My study builds on and contributes to several strands of the literature. First, my findings speak to the broader question of how benefits and taxes should be designed for couples. Policies differ widely across developed countries: The United States, for example, connects a wide range of (tax) benefits to the marriage contract; in contrast, in Sweden, taxation is individual and social insurance now is untied from marriage. To understand how to optimally design benefits for couples, one key step is to understand the nature of marriage market responses. I establish that tying social insurance to marriage has a significant impact on the volume of contracting in the marriage market. This finding contributes to previous literature that has shown mixed evidence on the existence of marital responses to implicit penalties or subsidies inherent in tax codes.[1] To my knowledge, this paper is the first to document responses to benefits that only pay out in the far future, and to analyze how tying benefits to the marriage contract affects matching. My discussion of the optimal design of survivors benefits in the presence of marriage market responses relates to Kleven et al. (2009), who study optimal taxation of couples' labor income. To focus on the spouses' labor supply responses, they take the marriage decision as given and assume a unitary model of the household. I begin to take a step in the opposite but complementary direction by focusing on the marriage margin and allowing for bargaining within the household.[2]

Second, my analysis of endogenous selection into marriage suggests that take-up of survivors insurance may be subject to adverse selection. This argument builds on a literature on adverse selection in private insurance markets and the premise of government intervention in response.[3] I focus not on a product offered in a private insurance market, but on a government-provided scheme that is supplied indirectly through the marriage contract.

[1] Whittington and Alm (2005), Bitler et al. (2006), and Light and Omori (2008) find that a smaller marriage penalty increases the rate of marriage relative to divorce or cohabitation, whereas Bitler et al. (2004), Fitzgerald and Ribar (2005), and Knab et al. (2008) reach the opposite conclusion. In the context of Austria, Frimmel et al. (2012) show that the removal of cash transfers upon marriage raises the marriage rate. See Alm et al. (1999) and Moffitt (1998) for surveys of earlier contributions.

[2] Casanova (2010) and Henriques (2012) also analyze the link between marriage and social security, albeit with different focuses. In the context of the U.S., Casanova (2010) studies the incentives for joint retirement and Henriques (2012) shows that husbands' claiming behaviors do not maximize total household benefits, but that they instead maximize their own benefits.

[3] Akerlof (1970) and Rothschild and Stiglitz (1976) develop the theoretical argument. Chiappori and Salanié (2000) begin to examine its key theoretical prediction, a positive correlation between demand and risk type. The empirical evidence, which I discuss in Section 5.2, is mixed.

Third, I contribute to the large literature that analyzes how married couples' decisions depend on factors that influence each spouse's relative bargaining positions,[4] by also analyzing impacts on earlier stages of the mating process: matching and cohabitation.[5] Contrary to models in which matching is modeled as the formation of married couples, here, the decision to form a couple is the decision to "start dating." This separation of couple formation from marriage is important in that it yields a marriage market equilibrium in which a strictly nonzero mass of couples chooses to cohabitate; a prerequisite for predicting for how social insurance impacts this empirically important group.[6]

Fourth, a key distinction between the extant literature and my paper is the isolation of long-run incentives. In contrast to a fiscal or legal change that affects the immediate payoffs from marriage, the reform that I study only affects transfers in the far future. The responses that I document thus reflect forward-looking behavior, with implied financial-planning horizons of up to 40 years. My findings thus relate to the literature that analyzes how individuals prepare for financial security in retirement.[7] Specifically, I show that when a publicly provided financial instrument is tied to marriage, financial-planning strategies extend beyond asset accumulation to marital decisions. A twofold intervention into marriage and financial markets could therefore create spill-overs between these markets: Indeed, one interpretation of my results is that when marriage gains in attractiveness *relative to other financial instruments*, entry into marriage contracts increases. In the presence of such spill-overs, any change in insurance markets could impact behavior in marriage markets, so long as social insurance is tied to marriage.

2 Survivors insurance in Sweden

Pre-reform survivors insurance *Eligibility.* Before the reform, which was implemented in 1990, survivors insurance was tied to the marriage contract. A divorced woman received no survivors benefits upon the death of her former husband. A widow, in contrast, could collect survivors benefits from the date of her husband's death (or her 36th birthday) given that the husband was less than 60 years at marriage and (i) they had a joint child, or (ii) they had been married for at least five years. Each married couple that satisfied one of these conditions was covered by survivors insurance during marriage. This scheme included the overwhelming majority of all married couples: Among couples that married in 1980, for example, ~86% satisfied one of the two criteria, and were thus covered. While marriage entitled a woman to survivors benefits, no other Social Security benefits were tied to marriage. Men were not eligible for survivors insurance.

[4]Such factors include gender-targeted transfers; laws governing divorce, alimony, or access to contraception; and sex ratios in the marriage market. See, e.g., Angrist (2002), Chiappori et al. (2002), Duflo (2003), Wolfers (2006), Stevenson and Wolfers (2006); Stevenson (2008), Edlund and Kopczuk (2009), Voena (2011), and Attanasio and Lechene, (forthcoming).

[5]My results for cohabiting couples relate to Chiappori et al. (2012), who study a reform of alimony laws in Canada and find that it has distinct impacts on cohabitating and married couples. My discussion of matching relates to Becker (1973), who shows that spousal specialization in market and nonmarket work is optimal in the presence of comparative advantage and learning by doing.

[6]In Sweden in 2000, 30 percent of the parents of newborn children were married, and 67 percent were cohabiting.

[7]See, e.g., Laibson (1997), Benartzi and Thaler (2004), Carroll et al. (2009), Beshears et al. (2010), Beshears et al. (2011), and Beshears et al. (2012).

Size of annuity. Survivors insurance replaced part of the husband's earned Social Security benefit, b_h. As in the U.S., earned benefits were proportional to lifetime earnings up to a ceiling; see Appendix B for details. A widow who was between 36 and 64 years old got a (monthly) survivors benefit equal to 40% of the husband's earned benefit, $s^{wife<65} = 0.4 * b_h$. For a widow who was 65 years or older, the survivors benefit also depended on her own earned benefit, b_w. Specifically, survivors insurance guaranteed that the wife got 50% of the Social Security income that the household would have received had the husband been alive. Her benefit was given by

$$s^{wife\geq 65} = 0.5^* \left(b_h + b_w\right) - b_w = 0.5^* \left(b_h - b_w\right), \tag{1}$$

so long as this did not exceed her pre-65 benefit, $s^{wife<65}$. For widows aged 65 or above, survivors benefits thus increased with the husband's earned benefit, but decreased dollar for dollar with her own earned benefit, and were equal to zero if the wife's earned benefit exceeds the husband's earned benefit. Put differently, (1) was increasing in the difference between the spouses' earned benefits. Widows who had little labor market participation thus relied heavily on survivors benefits.

Denoting the date of the husband's death by $t = 0$, the wife's 65th birthday by t_{65}, and her date of death by t_{death}, the annuity's value *upon realization* was thus given by

$$A = \sum_{t=0}^{t_{65}-1} \delta^t s^{wife<65} + \sum_{t_{65}}^{t_{death}} \delta^t min\left\{s^{wife<65}, s^{wife\geq 65}\right\}, \tag{2}$$

where δ is the wife's discount rate. Given the spouses' earned benefits, (2) illustrates that A was increasing in the annuity's (expected) duration, that is, the number of years that the wife outlived her husband. In 2002, the average realized payout to survivors insurance beneficiaries was SEK 35000 (~$ 5000), and the average duration of payments was eight years. Upon realization, the value of the average annuity, applying a zero discount rate, was SEK 280000 (~$40000).

Marital decisions are often made long before a spouse dies. The average marriage age in Sweden between 1980 and 1988 was 32.94 years for men and 29.98 years for women, and the average age of entry into widowhood was 74.7. Payout was thus, on average, expected to occur 45 years after marriage. Marital decisions taken with survivors insurance in mind therefore reflected a couple's discounted value of A.

Post-reform survivors benefits The reform eliminated the gender difference in survivors benefits in a manner that drastically reduced survivors benefits for women, while increasing them modestly (from zero) for men. In particular, a surviving spouse – regardless of gender – got a *one year* "adjustment transfer," amounting to 40% of the deceased spouse's earned benefit. Thereafter, the surviving spouse received Social Security income solely based on his or her own earned benefit, just like a divorced spouse would.

Transition The social security reform was discussed for the first time in the Parliament of Sweden on June 8, 1988. While it is unlikely that the entire population obtained knowledge of the reform upon its first mention in Parliament, I take a conservative approach and treat this date as the reform announcement.[8]

Transition rule. All couples that would have been eligible for survivors insurance if the husband had died on December 31, 1989 got pre-reform survivors insurance. All other couples get post-reform survivors insurance. Henceforth I will refer to the "old marriage contract" as the marriage contract that comes with pre-reform survivors insurance and to the "new marriage contract" as the contract that comes without pre-reform survivors insurance (but that otherwise is identical). The eligibility rules governing pre-reform survivors benefits, together with the transition rule, meant that couples that married before the husband turned 60 obtained the old marriage contract – with survivors insurance – if they (i) had a joint child together on or before December 31, 1989, and married on or before the same date; or (ii) had no joint child together on or before December 31, 1989, but married on or before December 31, 1984. This transition rule implies that couples that had a joint child at the time of the reform's announcement were given an option to enter the old marriage contract by marrying before December 31, 1989. In contrast, childless couples that were already married at the reform's announcement but that had entered marriage after December 31, 1984 faced the prospect of having survivors insurance revoked on December 31, 1989.

3 Conceptual Framework and Hypotheses

To examine how couple formation, marital decisions, and spousal welfare depend on the link between survivors insurance and marriage, I build a model of dating, marriage, and divorce. It shows that behavioral responses to a change in survivors insurance tied to marriage depend on whether an individual is married, cohabitating, or single (to be matched in the future) when this change is announced. I derive precise, distinct, testable predictions for individuals in each relationship stage. Proofs are in Section C.

3.1 Model

Consider a continuum of unmarried men of measure one and a continuum of women of measure one. The population plays a three-stage game. (Section 5.2 and Appendix D consider an infinite-horizon version.)

Stage 1 is the "matching stage," where singles match into heterosexual couples. Each man and woman is endowed with a characteristic that I call skill (e.g., educational attainment or IQ) and assume to be positively related with income. The skills of women, τ_w, are distributed according to some distribution W on $[0, 1]$, and the skills of men, τ_m, according to some distribution M on $[0, 1]$. A woman of type τ_w and a man of type τ_m derive a deterministic value $V(\tau_w, \tau_m)$ from their incomes in each period of marriage. I assume that $V(\tau_w, \tau_m)$ is continuously differentiable and supermodular, that is, spouses'

[8] Two working papers use this reform as an instrument for marriage to study the impact of marriage on child welfare (Bjorklund et al., 2007) and labor supply (Ginther and Sundström, 2010). These studies use 1989 as the reform year.

incomes are complements. This means intuitively that, all else equal, a highly skilled man places a higher value on marrying a highly skilled woman than does a lower skilled man, and vice versa. In Stage 1, couples do not make marital decisions; they only match into couples who "start dating," by proceeding to stage 2 together.

In the beginning of Stage 2, a stochastic marital "happiness" shock, $\tilde{\theta}_2$, is drawn from some distribution F on $(-\infty, +\infty)$. After observing this shock, the couple decides whether to marry or continue dating (wait, cohabit, etc.). Marriage entails a total marital surplus $S(\tau_w, \tau_m, \tilde{\theta}_2) = V(\tau_w, \tau_m) + \tilde{\theta}_2$ in Stage 2; if they instead continue dating, they obtain exogenous per-period utilities $\underline{u}_w$, and $\underline{u}_m$, which are normalized to zero. That is, $V(\tau_w, \tau_m)$ represents the value of marriage over and above cohabitation. All couples then proceed to Stage 3.

In the beginning of Stage 3, a new marital surplus shock, $\tilde{\theta}_3$, is drawn from some distribution G_{θ_2} on $(-\infty, +\infty)$ that is conditional on the realization of the marital surplus shock in Stage 2, θ_2. Upon observing this shock, married couples decide whether to divorce or stay married, and dating couples again decide whether to marry or not. Marriage entails the total marital surplus $S(\tau_w, \tau_m, \tilde{\theta}_3) = V(\tau_w, \tau_m) + \tilde{\theta}_3$ in Stage 3; if they instead divorce or remain unmarried, they obtain their exogenous utilities in Stage 3. The shocks $\tilde{\theta}_2$ and $\tilde{\theta}_3$ are non-negatively correlated: couples that are "happy" in Stage 2 are not less likely to be "happy" in Stage 3 as well. Specifically, I assume that G_{θ_2} weakly first-order stochastically dominates $G_{\theta_2'}$ if $\theta_2 > \theta_2'$: $G_{\theta_2}(x) \leq G_{\theta_2'}(x)$ for all x if $\theta_2 > \theta_2'$. Let G denote the *unconditional* (from the perspective of Stage 1) distribution of $\tilde{\theta}_3$.

So far, Stage 2 is isomorphic to Stage 3: Each couple experiences either a good or bad period, can alter its marital status in response to this, and then gets payoffs that depend on the marital decision. To analyze the impact of government-provided old-age support, I further assume that the man dies with probability p after the marital decision in Stage 3. If the couple is married at the time of his death, the government may transfer an annuity to the wife that renders her utility $U_A(\tau_w, \tau_m)$. Social insurance is thus tied to marriage. Further, the dependence of U_A on (τ_w, τ_m) captures that the value of the annuity varies depending on the match.

Utility is transferable and the spouses' interaction is efficient (Shapley and Shubik, 1971). I assume that the intertemporal allocation of utility in a match between a man and a woman is contracted upon *ex ante*, that is, in Stage 1. The man and the woman do not renegotiate this contract unless one of them credibly threatens to divorce the other. If renegotiation does occur, I assume that it results in the minimal change needed for a marriage to continue, provided that divorce is inefficient.

The model's key predictions that do not concern matching also arise if I remove the matching stage and only model (already formed) couples that obtain marital surplus shocks, and then decide, *in each period*, whether to be married and how to divide marital surplus.

3.2 Solution

I use backward induction. First, consider a couple in Stage 3 with the deterministic marital value component $V(\tau_w, \tau_m)$ and realizations θ_2 and θ_3 of the marital shocks $\tilde{\theta}_2$ and $\tilde{\theta}_3$. If the couple is married upon entry into Stage 3, the spouses remain married if and only if

$$(1-p)\,S(\tau_w,\tau_m,\theta_3)+pU_A\,(\tau_w,\tau_m)\geq 0. \tag{3}$$

Intuitively, the expected surplus from marriage in Stage 3 – a weighted sum of the joint surplus from marriage when the husband is alive, $S(\tau_w,\tau_m,\theta_3)$, and the wife's utility from the social security benefits when he is dead, $U_A\,(\tau_w,\tau_m)$ – must exceed the sum of their expected outside utilities. If the couple is unmarried at the start of Stage 3, their decision problem is identical: they choose to marry if and only if (3) is satisfied. Rearranging yields that the couple chooses marriage if and only if

$$\theta_3\geq\theta_{SB}\,(\tau_w,\tau_m)\equiv -V\,(\tau_w,\tau_m)-\frac{p}{1-p}U_A\,(\tau_w,\tau_m)\,, \tag{4}$$

where SB indicates the presence of survivors benefits. Couples that are sufficiently happy – get a sufficiently high shock – in Stage 3 choose to be married; otherwise, couples that enter the period married choose to divorce, and couples that enter unmarried choose not to marry. A couple's payoffs in Stage 3 are thus independent of the marital decision in Stage 2. The probability of being married in stage 3 conditional on the realization of the marital surplus shock in stage 2, θ_2, is $1-G_{\theta_2}[\theta_{SB}(\tau_w,\tau_m)]$. I also define the unconditional probability of being married in stage 3, (from the perspective of stage 1) as $\beta(\tau_w,\tau_m)=1-G[\theta_{SB}(\tau_w,\tau_m)]$.[9]

Consider Stage 2. The couple chooses to marry if and only if the following condition is satisfied:

$$S(\tau_w,\tau_m,\theta_2)\geq 0\;\Leftrightarrow\;\theta_2\geq\theta_{NSB}\,(\tau_w,\tau_m)\equiv -V\,(\tau_w,\tau_m)\,, \tag{5}$$

where NSB indicates "non-presence" of survivors benefits. The unconditional probability of being married in stage 2 is $\alpha\,(\tau_w,\tau_m)=1-F[\theta_{NSB}\,(\tau_w,\tau_m)]$.

Consider Stage 1, where individuals match, that is, form couples that start dating. By the law of iterated expectations, the expected value, in Stage 1, of a match with $V\,(\tau_w,\tau_m)$ can be written as $M\,(\tau_w,\tau_m)=\delta A\,(\tau_w,\tau_m)+\delta^2 B\,(\tau_w,\tau_m)$ where

$$\begin{aligned}A\,(\tau_w,\tau_m)&=\alpha\,(\tau_w,\tau_m)\,E\left[S(\tau_w,\tau_m,\tilde\theta_2)\middle|\,\tilde\theta_2\geq\theta_{NSB}\,(\tau_w,\tau_m)\right]\\ B\,(\tau_w,\tau_m)&=\beta(\tau_w,\tau_m)\left[(1-p)E\left[S(\tau_w,\tau_m,\tilde\theta_3)\middle|\,\tilde\theta_3\geq\theta_{SB}\,(\tau_w,\tau_m)\right]+pU_A\,(\tau_w,\tau_m)\right]\end{aligned}$$

and δ is the time discount factor.

Given transferable utility, the total expected surplus from a given match, $M\,(\tau_w,\tau_m)$, can be distributed between w and m. I denote the utility of the woman and man by $u\,(\tau_w)$ and $v\,(\tau_m)$, respectively. I search for a solution to the matching problem. Specifically, I search for a stable match as well as the (endogenous) utilities of all men and all women at the stable match(es). A match is stable if, for any $(\tau_w,\tau_m)\in W\times M$, $u\,(\tau_w)+v\,(\tau_m)\geq M\,(\tau_w,\tau_m)$, and $u\,(\tau_w)\geq 0$ and $v\,(\tau_m)\geq 0$. Intuitively, no two

[9] When the shocks are correlated across periods 2 and 3, the marital decisions are correlated as well. However, it remains true that the marital *decision* in period 2 per se has no causal influence on the payoffs in period 3.

matched individuals who are not matched with each other would both prefer to instead be matched, and no matched individual would be better off unmatched.

Lemma 1. *A stable match exists, at which the partners' utilities satisfy* $u(\tau_w) + v(\tau_m) = M(\tau_w, \tau_m)$.

3.3 Impact of a change in survivors insurance tied to marriage

The reform affected couples at different stages of their joint endeavors. Some couples were formed but had not yet married at the new law's announcement, and were allowed to take up survivors insurance by marrying within a limited time period. Others were already married when the reform was announced. Finally, many couples were not yet formed at the reform announcement. To derive testable predictions for the reform's impact on individuals in each relationship stage, I analyze the impact of an *unexpected* reform announcement at each of the three stages of the game.

3.3.1 Unmatched and Unmarried Individuals (Reform announced in Stage 1)

First consider individuals that were **u**nmatched and **u**nmarried (UU) at the reform's announcement.

Prediction UU1: Assortativeness of matching. Elimination of survivors insurance from the marriage contract induces a larger share of highly skilled men to match with highly skilled women.

The formal result driving this prediction is as follows:

Lemma 2. *Supermodularity of* $V(\tau_w, \tau_m)$ *implies supermodularity of* $M(\tau_w, \tau_m)$ *if* $U_A(\tau_w, \tau_m) = 0$*; but* $M(\tau_w, \tau_m)$ *may fail to be supermodular if* $U_A(\tau_w, \tau_m)$ *is decreasing in* τ_w*, given* τ_h.

In the absence of survivors benefits, the match that maximizes joint marital surplus is characterized by assortativeness: high-skilled men match with high-skilled women. In the presence of a government-provided annuity to widows that is higher for couples in which the husband earns more than the wife, however, assortative matching may fail. Intuitively, such an annuity *de facto* constitutes a subsidy to unassortatively matched couples in which the husband is of high skill. If the additional surplus from the subsidy more than outweighs the premium a skilled man puts on matching with a skilled woman, some high-skilled men prefer to match with less skilled women, and assortativeness breaks down.

3.3.2 Matched but Unmarried Couples (Reform announced in Stage 2)

Second consider couples that were **m**atched but **u**nmarried (MU) at the reform's announcement, and that could take up survivors benefits by marrying within a limited time period, that is, within Stage 2.

Prediction MU1: Retimed and extra marriages. The reform induces a marriage boom. This comprises "retimed" marriages and, given that match quality is stochastic, "extra" marriages that would never have occurred if the old marriage contract had remained available.

The mechanism driving the existence of a marriage boom is as follows. In Stage 2, unmarried couples that lose the survivors insurance if they wait to marry until Stage 3 choose to marry if and only if

$$V(\tau_m, \tau_w) + \theta_2 + \delta B_{\theta_2}(\tau_w, \tau_m) \geq \delta(1-p) A_{\theta_2}(\tau_w, \tau_m),$$

where $B_{\theta_2}(\tau_w, \tau_m)$ is isomorphic to $B(\tau_w, \tau_m)$ and $A_{\theta_2}(\tau_w, \tau_m)$ is isomorphic to $A(\tau_w, \tau_m)$, respectively, except that the expectations are formed over G_{θ_2}. The appendix establishes that this can be written as

$$\theta_2 \geq -V(\tau_m, \tau_w) - \Omega_A(\tau_m, \tau_w, \theta_2, p), \tag{6}$$

where $\Omega_A(\tau_m, \tau_w, \theta_2, p) > 0$ represents the value of the option on a claim to survivors benefits in Stage 3. A comparison with (5) shows that the threshold for marriage in Stage 2 decreases and is no longer independent of the marital decision in Stage 3. Thus, selection into marriage increases.

In Appendix C, I formally show that this increase can be decomposed into retimed and extra marriages. Intuitively, couples that marry *because* of this option have a lower $V(\tau_m, \tau_w)$ or θ_2 than couples that marry in Stage 2 also in the absence of reform. While their $V(\tau_m, \tau_w)$ or θ_2 are too low to warrant a marriage in Stage 2 without the option, they are not too low with the option, because their likelihood of marrying in the future warrants keeping the option alive. Of these couples, some would have married eventually, after observing a sufficiently high θ_3, even in the absence of reform. Others, however, would never have married because their θ_3 would have turned out too low. *Ex post*, these marriages thus turn out to be "extra."[10] Because the reform would induce only a retiming of marriages in the absence of uncertainty about θ_3, testing for existence of extra marriages offers a test of the assumption that match quality is stochastic.

Prediction MU2: Heterogeneous responses and economic incentives. It follows from (6) that the marriage boom's magnitude reflects the value of $\Omega_A(\tau_m, \tau_w, \theta_2, p)$ in the population of unmarried couples. Responses thus increase with the annuity's value at payout, $U_A(\tau_w, \tau_m)$. In Appendix C I also show that the option value Ω_A increases with p, the husband's likelihood of death, which affects the likelihood of payout.

Prediction MU3: Heterogeneous responses and expectations of lifelong commitment. Responses increase with θ_2. Intuitively, a higher happiness signal imply a higher expected happiness in the future, and thus a lower probability of divorce, which indirectly raises the likelihood of payout. Importantly, this prediction relies on the assumption that match quality shocks are positively correlated over time. Testing this prediction thus offers a test of this assumption, which is inherent in any theory of cohabitation as learning.

[10] A couple that enters marriage during the grace period does not know at that point whether their marriage would have occurred in the future in the absence of reform. This is intuitive – if a couple would know with certainty that their marriage would not have occurred in the future in the absence of the reform, the reform creates no option value, so the couple would not marry during the grace period.

Prediction MU4: Long-run divorce rate in marriage-boom marriages. Finally, couples that marry in the grace period have a higher future divorce rate. This is because couples that marry because of the reform have lower $V(\tau_m, \tau_w)$ or lower θ_2 – which imply a higher threshold $\theta_3(\tau_m, \tau_w)$ or a lower expected shock θ_3, respectively.

3.3.3 Matched and Married Couples (Reform announced in Stage 3)

Third, consider couples that were already **m**atched and **m**arried (MM) when the reform was announced, and that faced an *ex post* elimination of survivors insurance.

Prediction MM1: Marital instability. When the insurance provision is decoupled from marriage, a married couple's marital surplus falls. This induces some married couples to divorce.

Prediction MM2: Division of marital surplus. I show in Appendix C that in the marriages that survive, the wife's share of marital surplus (weakly) increases. This is because her expected utility from marriage is a weighted sum of her utility when her husband is alive and her utility when he is dead. The reform reduces her utility in the latter case. If this loss violates her participation constraint under the existing sharing rule in the household, her share of household utility, and hence her utility while the husband is alive, must increase for the marriage to continue, if that is indeed optimal. While the statutory loss induced by the reform is borne by the wife, intra-household bargaining results in the economic loss partly being borne by the husband.

4 Data

I use administrative data compiled by Statistics Sweden. From the *Historical Population Register* I obtain the universe of individuals that entered into marriage between 1968 and 2009, with exact dates of marriage and, if it occurred, divorce. I further obtain the exact death dates of all individuals who died during the same time period. For each marriage, I record whether (i) it ended in divorce, (ii) the husband died, leaving a surviving wife, (iii) the wife died, leaving a surviving husband, or (iv) the spouses remained married on December 31, 2009. For each spouse, I construct variables that describe their respective marital histories.

From the *Multi-generation Register* I obtain, for each child born in Sweden since 1971, the mother and father ID.[11] For married individuals that have children, I link spouses using the joint child ID. For each (ever or never married) couple that has joint children, I construct variables that capture the total number of joint and non-joint children, the birth dates of each of these children, and whether the children were born outside of marriage.

[11] For children born in Sweden, I can identify the mother (father) in 100 (98) percent of the cases. This contrasts with, for example, the United States, where paternity establishment outside of marriage is incomplete (Rossin-Slater, 2012). In Sweden, paternity is presumed when parents are married; for unmarried couples, the Social Services automatically conduct a paternity investigation if the child's father has not been reported within three months of the child's birth.

From the *Longitudinal Integrated Database for Labour Market Research*, I obtain labor market outcomes and demographic information for the universe of adults (aged 16 and up) for each year between 1985 and 2009. Specifically, I obtain the exact annual taxable labor income for each year in which an individual earns such income and his or her month and year of birth, education level, and immigration status. For 1985 and the years 1990 to 2009 this database identifies the spouse, which enables me to link individuals that marry but that do not have joint children.

From the *Military Enlistment Register* I obtain information on the cognitive ability of each male born between 1951 and 1988. For these men, a cognitive ability test that provides an accepted measure of intelligence (see e.g. Carlstedt, 2000) was administered during mandatory enlistment.[12]

In each part of the empirical analysis, I present summary statistics for the samples that I use.

5 Survivors insurance and selection into marriage

Couples that were both married and had a joint child together on or before December 31, 1989 were entitled to survivors insurance. In this section, I isolate the impact of (removal of) survivors insurance on the marriage decision.

Sample and descriptive statistics I start from all couples that had a joint child between January 1, 1971 and January 1, 1989. Live children born on January 1, 1989, were likely conceived before the reform announcement, which took place less than seven months earlier. For these couples, the eligibility for survivors insurance was determined by their marriage decision alone. Those who married before January 1, 1990, obtained survivors insurance; those who married thereafter did not. I study the marital behavior of these couples starting one decade before the eligibility threshold, and examine bunching in the distribution close to this threshold. Table 1 displays summary statistics.

5.1 Prediction MU1: Retimed and extra marriages

Figure 1 plots the empirical distribution of new marriages at a quarterly frequency. The figure shows that entry into marriage exhibits a seasonal pattern, with more marriages in the spring and summer. Further, there is a marriage boom at the eligibility threshold for survivors insurance, indicated by the red dashed line. After this threshold, the distribution displays missing mass.

Even though the reform was announced in the second quarter of 1988, indicated by the gray dashed line, the excess mass is concentrated at the eligibility threshold. This is consistent with the theory because waiting to marry is costly only when it entails giving up the option to get the old marriage contract, that is, in the last period. Another potential reason for this seemingly late response may be

[12] The test consists of four subtests: logical ability, verbal ability, technological comprehension and metal folding. For each part $p \in \{1, 4\}$, each individual i is given a score $s_{pi} \in \{0, 9\}$. Following Lindquist and Vestman (2011), I sum the scores from each part, $s_i = \sum_p s_{pi}$. I then percentile rank the variable s within each enlistment year and transform these scores using the inverse normal function. The resulting variable q_i is normally distributed with a mean of zero and a standard deviation of one.

my conservative choice of reform announcement date. Whereas the reform was discussed in Parliament in June 1988, Appendix Figure A1 shows that media reporting about the reform was heavily concentrated in the last three months of 1989, suggesting that it then may have been more salient (Chetty et al., 2009).

I start by estimating the size of the marriage boom. Because the distribution "vanishes" post-reform, I first use only pre-reform data, excluding marriages that were entered before the first quarter of 1990. In the spirit of Saez (2010), Chetty et al. (2011), Manoli and Weber (2011), Kleven and Waseem (2012), and Kopczuk and Munroe (2012) I start by estimating the following regression:

$$N_s = \alpha + \beta \left(\mathbf{1} \left[s = s^* \right] \right) + g(s) + \zeta_q + \varepsilon_s, \tag{7}$$

where N_s is the number of marriages in quarter s; $\mathbf{1}[s = s^*]$ is an indicator variable that takes the value of one at the eligibility threshold, $s = s^* = 1989q4$; the function $g(s)$ is a higher order polynomial; and ζ_q are (four) quarter fixed effects. Intuitively, I fit a polynomial to the counts plotted in the figure before 1990 Q1, accounting for seasonality. Because the excess mass is concentrated at s^*, β measures the size of the marriage boom, that is, the number of marriages that would not have occurred in 1989 Q4 in the absence of reform. Table 2 presents estimates of β from specifications with varying polynomial degrees. All estimates are in the range of 46000 induced marriages. The null hypothesis that there is no excess mass at the threshold relative to the counterfactual distribution (obtained by setting $\mathbf{1}[s = s^*]$ equal to zero) is rejected with t-statistics that imply p-values satisfying $p < 10^{-9}$.

This marriage boom, coupled with a missing mass after the eligibility threshold, suggests that couples that would have otherwise married after the threshold retimed their entries into marriage so as to take up survivors insurance. This can be thought of as an intensive margin response. The theoretical framework suggests that, in the presence of uncertainty about future match quality, the reform also induced couples to enter marriage that, in the absence of reform, would never have ended up marrying. Such extra marriages resemble an extensive margin response. A decomposition of the marriage boom into retimed and extra marriages requires estimation of a counterfactual density, that is, an approximation of the density in the counterfactual scenario, without any reform. Appendix Figure A2 displays attempts to use the coefficients obtained in estimation of (7) to predict a counterfactual density (out of sample); these counterfactuals are sensitive to the choice of polynomial. While the functional form assumptions made here can clearly be improved upon, the key challenge arising from a lack of "post-reform" observations is that any functional form assumption will require projection far out of sample (and be arbitrary).[13]

Empirical framework using two dimensions of information To nevertheless decompose the marriage boom into retimed and extra marriages, I further develop my empirical framework, by (i) using a second dimension of information to decompose my sample into subsamples, and (ii) coupling this with

[13] For the same reason, the (within sample) predictions of the counterfactual number of marriages in 1989 Q4 (when setting $\mathbf{1}[s = s^*]$ equal to zero) are imprecise and vary considerably between specifications, ranging from 1385 (using a polynomial of degree two) to 92 (six). This, in turn, yields estimates of the number of induced marriages *relative to* the counterfactual, ranging from 33 to 518. Alternative, and more precise, estimates of these are presented below.

specific assumptions about the impact of this second dimension of information on marital behavior.[14] Specifically, I exploit an empirical fact illustrated in Figure 2: Entry into marriage is concentrated around the birth of a couple's first child. To illustrate this further, Figure 3 replicates Figure 1 for different subsets of the sample. Panel A depicts entry into marriage in the subsample of couples whose first joint child was born in 1987 or 1988 – couples included in my sample whose children were born the latest. Panels B, C, and D, respectively, depict other "cohorts" of couples, whose firstborns were born in earlier time periods. The panels show that entry into marriage decreases within a year of this child's birth. This, in turn, suggests that the *fall* in the number of new marriages around the reform announcement can be explained by the fact that the last cohort of couples included in my sample has their first joint child before or within six months of the reform announcement. Intuitively, my estimation strategy exploits this second dimension of information – the date of birth of a couple's first child – by using early cohorts, whose marital behavior is observable for a longer period of time pre-reform, to help predict how the marital behavior of late cohorts would have evolved in the absence of reform. For the earliest cohort included in my sample, I observe 19 years of post-childbirth, pre-reform marital behavior.

Specification. Because the sample includes couples whose first joint child was born during 72 quarters, from the first quarter of 1971 until the last quarter of 1988, I divide the sample into 72 cohorts. Each cohort $c \in \{1, 72\}$ consists of couples whose first joint child was born in a given quarter, where $c = 1$ represents the first quarter of 1971, $c = 2$ the second quarter of 1971, and so on. For each of these cohorts, I observe the marital behavior at a quarterly frequency. I estimate the following regression:

$$n_{cs} = \alpha + \eta_c + \zeta_q + \beta_c \left(\mathbf{1} \left[s = s^* \right] \right) + \gamma_c \left(\mathbf{1} \left[s > s^* \right] \right) + g(s) + h(t_{pre-birth}) + j(t_{post-birth}) + \varepsilon_{cs}, \tag{8}$$

where n_{cs} is the natural logarithm of N_{cs}, the number of marriages in quarter s in cohort c, $n_{cs} = ln(N_{cs})$. I use the natural logarithm because the distribution of new marriages exhibits nonlinearities; I show the distribution of n_{cs} in Appendix Figure A3. As before, $g(s)$ is a higher order polynomial in time (quarter) and ζ_q capture seasonality. Further, η_c are cohort fixed effects. The β_c capture the cohort-specific increases in entry into marriage at the eligibility threshold, $s = s^*$, and the γ_c capture the corresponding reductions in entry after this threshold. These effects can be thought of as proportional, given that n_{cs} is the natural logarithm of the number of marriages. They are allowed to be cohort-specific because each cohort experiences the reform at different durations since birth of the first joint child, and hence – as is illustrated in Figure 3 – have different baseline levels of entry into marriage pre-reform. The functions $h(t_{pre-birth})$ and $j(t_{post-birth})$ are higher order polynomials in the number of quarters before and after the first child's birth, respectively. Inclusion of these time trends allows me to use early cohorts, whose marital behavior is observable for a longer period of time *post-childbirth* before the reform, to predict how the trend in marital behavior of late cohorts would have evolved in reform's absence. Intuitively, inclusion of these functions can be thought of as recentering the distributions around the birth of a couple's first joint child, and then exploiting the fact that different cohorts were "hit" by reform at different distances

[14] Similar in spirit, Marx (2012) develops a framework that incorporates a(nother) second dimension of information to improve upon estimation of a counterfactual density, and that allows for estimation of extensive margin responses.

in time from childbirth. Inference about the impact of the reform on take-up of marriage relies on the following assumption: In the absence of reform, couples marrying at the threshold would behave like couples marrying before the threshold at the same duration since childbirth, after allowing each cohort of couples to have a separate marriage propensity (that is, after allowing for vertical shifts of each cohort's recentered distribution). This is akin to a "common trends" assumption with respect to how the rate of marriage declines with distance from the date of childbirth.

After estimating this regression, I obtain the predicted arithmetic cohort-specific frequencies, $\hat{N}_{cs}$.[15] To predict counterfactual frequencies, $\hat{K}_{cs}$, I set $\mathbf{1}[s = 1989q4]$ and $\mathbf{1}[s > 1989q4]$ equal to zero, and use only the other estimated coefficients. I obtain seasonally adjusted frequencies, $\tilde{N}_{cs}$, and counterfactuals, $\tilde{K}_{cs}$, by also setting the quarterly dummies to 0. I rescale the seasonally adjusted frequencies so that the total number of marriages is preserved. I finally aggregate the cohort-specific counterfactual frequencies to sample-wide frequencies by calculating $\hat{K}_s = \sum_c \hat{K}_{cs}$ and $\tilde{K}_s = \sum_c \tilde{K}_{cs}$.

I use the estimated counterfactual density to decompose the marriage boom into intensive (retimed) and extensive (extra) margin responses. Specifically, I denote by A the estimated number of induced marriages at the eligibility threshold, $A \equiv \left(N_{s^*} - \hat{K}_{s^*}\right)$, and by B the estimated sum of missing marriages post reform, $B \equiv \sum_{s>s^*} \left(N_s - \hat{K}_s\right)$. If all induced marriages represent retimed marriages, then $A = B$. If $A > B$, then the difference $(A - B)$ represents extra marriages. Finally, I calculate one estimate of the change in the probability of marriage at the threshold by $\frac{\triangle p_{s^*}}{p_{s^*}} = \frac{\left(N_{s^*} - \hat{K}_{s^*}\right)}{\hat{K}_{s^*}}$; I discuss alternative estimates for the change in marriage probability in Section 5.2.

Bootstrap procedure. I use the following bootstrapping procedure to obtain standard errors. Because I use panel data (where cohort c is the panel variable and quarter s the time variable), I use a cluster bootstrapping procedure, with 10000 samples drawn with replacement. Specifically, starting from the collapsed data, I create each new sample by drawing (each of the s observations for) 72 cohorts with replacement. Each of my randomly drawn samples thus corresponds to a panel with 72 cohorts (with seasonality preserved). If an entire panel is drawn twice, the two draws are treated as different panels (samples). For each of the 10000 samples, I estimate each of the statistics above. The standard error for each of the statistics is estimated by computing the standard deviation of the 10000 estimates of this statistic.

Results Figure 4 Panel A displays the empirical distribution and the arithmetic sample-wide counterfactual (estimated using the specification that minimizes the Akaike Information Criterion (AIC)). Their seasonally adjusted counterparts are displayed in Panel B. In both panels, relative to the empirical distribution, the predicted frequency distribution entails "excess mass" in the last quarter of eligibility for survivors insurance, but displays a persistent "missing mass" post reform.

[15] I calculate $\hat{N}_{cs} = exp(\hat{n}_{cs} + \frac{\hat{\sigma}^2_{cs}}{2}) = exp(\hat{n}_{cs}) exp\left(\frac{\hat{\sigma}^2_{cs}}{2}\right)$, where $\hat{\sigma}^2_{cs}$ is the squared standard error of the regression. This is because, if N_{cs} and n_{cs} are random variables satisfying $n_{cs} \sim N\left(\mu, \sigma^2\right)$ and $N_{cs} = exp\left(n_{cs}\right)$, then $E\left(N_{cs}\right) = exp\left(\mu + \frac{\sigma^2}{2}\right)$. Because $exp\left(\frac{\hat{\sigma}^2_{cs}}{2}\right)$ is close to one, however, similar results obtain if I let $\hat{N}_{cs} = exp(\hat{n}_{cs})$.

Figure 5 replicates Panel A in Figure 4, adding indications of the areas A and B. Table 3 presents the estimated sizes of these areas, using three different higher order polynomials. The yellow area in Figure 5 depicts the marriage boom induced by the reform, which is estimated to have induced $A = 44\ 305$ marriages at the eligibility threshold (using the specification that minimizes the AIC). The white area between the red dashed and blue solid lines depicts the missing marriages post reform. These retimed (missing) marriages are estimated to sum to $B = 25\ 600$. Retiming of marriages can thus explain only parts of the extra mass in the bunching region. The difference, $(A-B) = 18\ 705$ is the estimated number of extra marriages that would not have materialized in the absence of reform.

Row 4 of Table 3 presents the estimated change in the probability of marriage at the eligibility threshold, $\frac{\triangle p_{s*}}{p_{s*}} = 18.25$. The excess mass at the threshold is 1825% of the counterfactual frequency in that quarter. This suggests that, on average, a couple is *18 times* more likely to marry in the last quarter of 1989 than it would have been in the absence of reform. I discuss the interpretation of this estimate further in Section 5.2 below.

5.2 Prediction MU2: Heterogeneous effects and economic incentives

The response in entry into marriage should be larger, the larger is the annuity's expected value. One of it's determinants is the husband's likelihood of death, p, that is, the likelihood of payout. Because mortality increases with age, Figure 7 shows the distribution of marriages in two subsamples with different husband ages. I further replicate regression (8) for different subsamples and present the results in Appendix Table A1. While the baseline rate of marriage is higher among men who are younger at marriage, the change in likelihood of marriage is higher, the older is the husband.

Two challenges arise in interpreting these results, however. First, to isolate the impact of age, I should control for other couple-specific variables, especially other variables than age that influence the annuity's expected value at payout. I thus need a regression framework.

Second, I show in Section 5.1 that the reform's impact on entry into marriage is not accounted for by retiming alone. Given this finding, an analysis of the distribution of marriages – and hence of married couples – introduces selection when analyzing impacts on the probability to marry. Specifically, in the presence of "extensive margin responses" (extra marriages), the sample at risk of entry into marriage, at any given point in time, not only includes all ever-married couples, but all couples that face the choice between taking up the marriage contract and cohabiting (waiting). This intuitively suggests that the appropriate framework is a hazard model, with the sample at risk defined as all couples whose first joint child was born before the reform announcement. Indeed, in Appendix D I show that couples' utility maximization problem when choosing between cohabitation and marriage (and survivors benefits) can be characterized by a hazard model framework. I do this by developing an infinite horizon version of the model presented in Section 3 above. To highlight the impact of the change in marriage contract on the decision whether to marry or wait (cohabitate), I there abstract from the matching decision and consider an already-formed couple that, in each time period, decides whether to enter marriage.

Empirical framework: Estimating the hazard function The reform announcement affects all couples with a joint child born before 1989. This thus constitutes the sample of couples that is "at risk" for marriage. Figure 2 shows that the probability of marriage more than 7 years before childbirth is essentially zero. I therefore define a couple whose first joint child is born in quarter s to become under risk for marriage 7 years earlier, in quarter $(s-28)$. Because I observe marital behavior from 1969, I use all couples whose first child was born between 1976 and 1989.

I estimate an extended Cox model (Pettitt et al., 1999; Fisher et al., 1999), where covariates are time dependent with fixed functions of time. For couple i, I define an "ever married indicator" δ_i; the time of marriage (measured from 28 quarters before childbirth), X_i; and a covariate path, $Z_i(t), t \in [0, X_i]$, of potentially time-varying covariates while the couple is at risk for marriage. Specifically, in keeping with the formulation above, I define two time-varying covariates, $s_i^*(t)$ and $post_i(t)$, where $s_i^*(t) = 1$ when $t_c \equiv s^* - (c-28)$, that is, in the last quarter of 1989. Each cohort of couples, defined by the quarter of birth of their first joint child, thus experiences the reform at different durations since childbirth. Similarly, I let $post_i(t) = 1$ when $t > t_c$.

I assume that conditional on a couple's covariate history, the hazard for marriage at time t depends only on the value of the covariates at that time, $h(t; \mathbf{Z}_i(t)) = h_0(t) exp(\beta \mathbf{Z}_i(t))$. The baseline hazard at time t, $h_0(t)$, is left unspecified.

I use the following regression model:

$$h(t; \mathbf{Z}_i(t)) = h_0(t) exp(\beta s_i^*(t) + \gamma post_i(t) + \delta_1 \mathbf{F}_i(t) + \delta_2 \mathbf{D}_i(t)), \tag{9}$$

where $F_i(t)$ is a vector of potentially time-varying financial characteristics that influence the annuity's expected value as defined in (1) and (2): the man's labor income and share of household income, and each partner's employment status and birth year. $\mathbf{D}_i(t)$ is a vector of demographic and other observable couple characteristics: the partners' levels of education, their marriage parities, and the couple's completed fertility. In alternative specifications, I control more flexibly for the man's labor income and birth year, by including indicator variables for eight male income ranges $l \in \{1, 2, ..., 8\}$ and eight birth year ranges, $b \in \{1, 2, ..., 8\}$. Each labor income range is SEK 25k, with the highest range $l = 8$ including incomes of $175k$ and above in 1988 (12% of the sample); each birth year range is four years. I refer to the vector which includes these flexible controls as $\tilde{F}_i(t)$.

The predicted hazard rate at t is the predicted probability that couple i in cohort c marries t quarters after $(c-28)$, given that they are unmarried until then. I calculate the ratio of these predicted probabilities for marriage in 1989 $Q4$ relative to marriage in another quarter, given by the hazard ratio of marriage in 1989 Q4, $\hat{h}_{1989\ Q4} = exp\left(\hat{\beta}\right)$. Intuitively, a hazard ratio of 10 means that a couple is 10 times more likely to marry in 1989 Q4 relative to the counterfactual scenario*, given that the couple was not yet married in the beginning of that quarter*, and holding constant couple characteristics. Standard errors are clustered on the child's quarter of birth, and standard errors of the hazard ratio are calculated using the delta method.[16]

[16]Another approach to addressing potential unobserved heterogeneity that causes correlation across observations in

Results The upper panel of Table 4 presents results from estimation of (9). The estimated hazard ratio in the full sample, is 16.88 when controlling for $F_i(t)$, the a vector of financial characteristics that influence the annuity's expected value; and 17.04 when also controlling for $\mathbf{D}_i(t)$, the vector of demographic and other observable couple characteristics that I observe. This means that a couple that is unmarried at the end of 1989 $Q3$ is, on average, 1704 percent – or 17 times – more likely to marry in the next quarter than it would have been in the absence of reform. The lower panel replicates these results including flexible controls for male income and birth year; the results remain unchanged.

I then study how this hazard ratio varies with two different measures of the economic value of the annuity, by adding interactions between $s_i^*(t)$ and each male labor income group l in $\tilde{\mathbf{F}}_i(t)$, and between $s_i^*(t)$ and each male birth year group b. I thus estimate the following regression model:

$$h(t;\mathbf{Z}_i(t)) = h_0(t)exp(\sum_l \alpha_l s_i^*(t) + \sum_b \beta_b s_i^*(t) + \gamma post_i(t) + \delta_1 \tilde{\mathbf{F}}_i(t) + \delta_2 \mathbf{D}_i(t)), \tag{10}$$

Here, the estimated hazard ratio for marriage in 1989 $Q4$ for a couple with male labor income l and birth year b is given by $e^{\alpha_l+\beta_b}$. The upper panel of Figure 8 plots the estimated hazard ratios for couples in which the male was born between 1952 and 1956, for different male income ranges. The hazard ratio increases with male income, and thus with the annuity's expected value: A man with income in the range 25k-50k is 16 times more likely to marry in 1989 $Q4$; the corresponding figure for men whose income instead is in the range 125k-150k is 21. In this sample, 150k is the 77:th percentile of labor income. In next range, the the hazard ratio decreases, which may reflect the fact that some income earners exceed the Social Security limit. (This limit is calculated based on pension rights income, which in addition to labor income includes some social insurance payments; see Appendix B for details). The last group includes couples that exceed this limit with certainty; this is indicated by the green dashed line. The hazard ratio is thus increasing in husband income in the range where a higher husband labor income raises the annuity's value. Above the Social Security limit, the annuity's value remains constant. The fact that couples in the highest bracket, above the 91th income percentile in this sample, are less responsive may suggest that the annuity is less important in the highest earning households.

The lower panel of Figure 8 plots the estimated hazard ratios for couples where the male earns income in the range 50k-75k, for different male birth year ranges. The hazard ratio is increasing with male age: A man born before 1964, and who hence is younger than 25 in 1989 $Q4$, is 9 times more likely to marry in 1989 $Q4$; the corresponding figure for men born between 1952 and 1948 is 18. For yet older cohorts, the ratio remains constant. The red dot indicates the same group of couples in both panels (and hence the same hazard ratio). The fact that responses are stronger for older men likely reflects both the fact that a higher probability of death raises the annuity's expected value, and the fact that an older husband has completed more years of pension-earning income, which raises the annuity's expected payout (see Section B for details).

cohorts is to model such correlation explicitly as shared frailty, using a random effects model. Because I do not know the correlation structure, I use clustered standard errors in my baseline specification, which assumes no particular model of correlation. Instead using a shared frailty model yields similar results, which are available upon request.

Even to the extent that husband age captures mortality risk, it is an observable characteristic that could potentially be priced into a private annuity. I therefore turn to investigating whether responses to the change in marriage contract vary with husbands' *ex post* mortality. Using death records, I identify all men in my sample that died within five years of January 1, 1990. I then reestimate (9), including an indicator variable for such couples, $1[H\ dies\ within\ 5years]_i$, as well as the interaction term $1[H\ dies\ within\ 5years]_i^* s_i^*(t)$. The upper panel in Table 5 presents the results. When controlling for financial characteristics that influence the annuity's expected value, $\mathbf{F}_i(t)$, and all other characteristics that I observe, $\mathbf{D}_i(t)$, the implied hazard contribution from the interaction term is 1.14. Thus, a couple where the husband dies within five years has a 14 percent higher hazard of marriage in the last quarter of 1989 than a couple where the husband remains alive for at least five years; this difference is significant at the 5 percent level. The lower panel presents results from separate estimations of (9) in the sample where the husband dies within five years, and the in the sample where he lives for at least five years. In the mortality sample, the estimated hazard of marriage is 21.28; in the live sample, it is 17.04.

I replicate this in subsamples with different *ex post* mortality; Appendix Table A2 presents summary statistics for these samples. I estimate (9) for each distinct *ex post* mortality sample and plot the estimated hazard rates for marriage in 1989 Q4 in Figure 9 as points along the red line (these estimates are also reported in Table A3, both with and without controlling for demographic observables, $\mathbf{D}_i(t)$). Take-up of marriage in the last quarter of 1989 Q4 is higher, the shorter is the time span until death of the husband. I thus observe a positive correlation between couples' take-up of insurance – through marriage – and couples' expected cost of coverage. This may suggest the presence of asymmetric information (Chiappori and Salanié, 2000), which may pose one challenge to private markets for annuities in Sweden. This argument builds on a literature on adverse selection in *private* insurance markets and the premise of government intervention in its response, for which evidence is mixed.[17] My innovation lies in focusing not on a product provided in a private insurance market, but on a government-provided scheme that is provided indirectly through the marriage contract. The precise conditions under which private information would generate adverse selection in private annuities markets are the following: Couples have information about husband mortality that a private insurer cannot fully price in. What my results show is that if insurance companies would have access to all the demographic information that I observe – the spouses' levels of education at marriage, the spouses' marriage parities, and the couple's completed fertility – but no more information, then adverse selection would arise in such a private market. Whether asymmetric information indeed would be a problem thus depends on whether insurance companies can or are allowed to price private annuities based on an even larger information set. If not, adverse selection

[17] Finkelstein and Poterba (2002, 2004) and Rothschild (2007) reject the null hypothesis of symmetric information in U.K. annuities markets. In the US life insurance market, Cawley and Philipson (1999) find no evidence of adverse selection in the US life insurance market, whereas He (2009) does (using a different sample). Hendren (2012) theoretically shows that sufficiently asymmetric information can explain insurance rejections, and provides empirical evidence in support of this theory in life insurance, long-term care, and disability markets in the US. Chiappori et al. (2006) theoretically analyze multidimensional private information, for which empirical evidence is presented by Heiss et al. (2007) in the market for Medicare supplemental insurance, by Fang et al. (2008) in annuities markets, and by Finkelstein and McGarry (2006) in the market for long-term care insurance. In addition to risk type, Cohen and Einav (2007) and Einav et al. (2010) analyze the import of risk preferences, and Spinnewijn (2012) the import of beliefs about risk types.

may be one reason why private markets for annuities are underdeveloped in Sweden, even to this date.

5.3 Prediction MU3: Heterogenous effects and expectations of lifelong commitment

Because the survivors insurance in Sweden only accrued to widows and never to divorcees, the *realized* value of the annuity will de facto be zero for couples that eventually end up divorcing. Couples that, at announcement, attach a smaller probability to the event that they will remain together for life thus have a weaker incentive to respond to the reform announcement by taking up marriage. This is captured in the model by the prediction that the response to reform is lower, the lower is the realized stochastic component of marital surplus in period 2, θ_2.

Testing this prediction poses a challenge, however. A precise measure of couples' ex ante beliefs in the strength of their union is not observable in my administrative data. To obtain a proxy for having a weak belief at the time of reform in remaining together for life, I turn to a 1995 law that legalized same-sex marriage in Sweden. Because my data set contains all marriages that were entered into between 1969 and 2009, including same sex marriages from 1995 onward, I can create a sample consisting of all individuals who entered into a same-sex marriage between 1995 and 2008, and hence revealed a same-sex preference. Now consider the following two assumptions: (i) Sexual preferences at two distinct points in time t and $t+s$, $s>0$, are positively correlated, and (ii) a stronger same-sex preference reduces the expected duration of a heterosexual marriage. Under assumption (i), an individual who in 1995 or later reveals a same-sex preference already had, on average, a weaker opposite-sex preference in 1988. Under assumption (ii), this translates into a weaker average belief about the longevity of heterosexual marriage. Under these two potentially restrictive assumptions, I thus predict that the marriage rate response would be smaller in this sample than the estimated response in the total population.

In Figure 6, the number of same-sex unions is depicted by a blue line, starting in 1995. At legalization, a spike in same-sex marriages occurs (potentially reflecting latent demand), after which the number of marriages first falls, and then begins to rise. The black solid line depicts the opposite-sex marriages among the same individuals, starting in 1970. While the figure displays bunching in the distribution of opposite-sex marriages in response to the reform, this spike is visibly "smaller." Indeed, in this sample, estimates of the increase in the probability of heterosexual marriage, relative to the counterfactual, are in the range of $\frac{\triangle p_{s^*}}{p_{s^*}} = 2$. This estimate is roughly ten times smaller than the estimated impact in the average population. Under the assumptions postulated above, this suggests that the response was smaller in couples with weaker *ex ante* beliefs in the union's survival.

5.4 Prediction MU4: Long-run divorce rate in marriage-boom marriages

When marital quality evolves over time, replacing the existing marriage contract with a less desirable one attaches option value to marrying into the desirable contract, which induces couples to rush to marriage even though they, in the absence of reform, would wait to see whether the relationship improves. When match quality exhibits persistence, rushed marriages should be more likely to end in divorce. Indeed, this is why, in the absence of reform, such marriages would never materialize. I thus have a prediction about *the nature of selection* into marriage in the last quarter of 1989, when rushed marriages were entered into.

To examine this, I compare the incidence of divorce among couples that marry in the last quarter of 1989 with those that marry into the *same* marriage contract earlier. Summary statistics for this subsample are presented in Columns 2 and 3 of Table 1. I first estimate the following

regression using OLS: $1\left[Divorce_x\right]_{imd} = \alpha + \beta \mathbf{1}\left[marr = s^*\right]_i + X_i'\theta + \eta_m + \zeta_d + \epsilon_{imd}$,where the variable $Divorce_x$ takes the value of one if couple i divorces within x years of marriage; the main explanatory variable of interest is a dummy taking the value of one if the couple married in the last quarter of 1989; η_m and ζ_d capture wedding month and day of week fixed effects, respectively; and X_i captures observable couple-specific characteristics: the spouses' ages, household income, and h's share of household income at marriage; the spouses' levels of education and immigration status; h's IQ level; the spouses' marriage number; and completed fertility. The key coefficient of interest is β, which measures the difference in marital stability of marriage-boom marriages relative to other marriages with the same marriage contract. Robust standard errors are clustered on the (marriage month*marriage day of week) level.

Table 6 presents the results. Consistent with the prediction, the estimates suggest that marrying in the boom is associated with a 2.28 percentage point higher probability of divorce within 5 years, and a 4.23 percentage point higher probability of divorce within 10 years, at the sample mean. The average 10-year divorce rate is 0.16. Appendix Table A4 reports average marginal effects from Probit estimation; the results are similar.[18] Because the couples that married in the last quarter of 1989 on average are older and have a higher household income – as is illustrated in Table 1 – and thus have characteristics for whom divorce rates are generally lower, it is crucial to control for these demographic and economic factors. In the Appendix, Table A5 illustrates this by omitting the couple-specific characteristics. This implies that the marginal effect should be evaluated at the mean characteristics of couples marrying in the last quarter of 1989. My Probit estimation yields $Pr\left(Div_{10}\left|X_{1989q4}, s^* = 1\right.\right) = 0.148$ and $Pr\left(Div_{10}\left|X_{1989q4}, s^* = 0\right.\right) = 0.093$; the estimated discrete increase in the probability of divorce if marrying in the last quarter of 1989 is thus 0.055. If this difference is driven by a higher divorce rate among rushed marriages, given that such marriages constitute roughly 40% of marriage-boom marriages, the implied probability of divorce in rushed marriages is 0.229. Of the 18000 rushed marriages, only around 13800 last for 10 years.

On the one hand, this implies that some long-lasting unions were prompted by the reform. On the other hand, policy-induced marriages are more likely to dissolve, underscoring that the effectiveness of policies that aim to promote lasting commitment in unions should not be evaluated solely on the policy's impact on marriage take-up because a large share of the "complier marriages" may end up in divorce.

[18] I also perform matching on age at marriage, income, and husband share of income to identify a match for each couple that marries in the last quarter of 1989, and compare the marital duration. Results are similar and available upon request.

6 Survivors insurance and pre-existing marriage contracts

I now turn to the second group of couples, those who were already married at the announcement of reform, and analyze the causal impact of survivors insurance on family well-being. For these couples I test two predictions, using the same empirical methodology.

Sample and descriptive statistics My baseline sample includes all couples that entered marriage within 180 days of the eligibility threshold, January 1, 1985, and in which the wife was born in 1945 or later. Throughout the analysis, I also present results for a narrower window around this threshold. I exclude all couples that had a joint child before the reform announcement, June 8, 1988.

All of these couples married into the same marriage contract, with survivors insurance. When the reform was announced in 1988, all couples that had married before January 1, 1985 were allowed to keep this contract. For couples that married thereafter, the survivors insurance reform revoked the old marriage contract and replaced it with the new contract without survivors insurance, unless the couple had a joint child before December 31, 1989. Table 7 presents summary statistics for the baseline sample, as well as for couples that entered marriage within 180 days of New Year's Eve one year earlier, in 1983. These groups are similar, but relative to the sample studied in Section 5, the spouses studied here are more likely be in their second marriages. This is consistent with the fact that second marriages are more likely to be childless (for three and a half years after marriage) in my data.

Empirical Framework An evaluation of the causal impact of survivors insurance on family well-being requires a comparison of couples who have such insurance with couples who do not. Section 5 illustrates, and the theory predicts, that couples strategically influence entry into marriage in order to take-up survivors insurance. This margin can thus not be exploited to identify causal effects.

Instead, the ideal experiment would be to randomly allocate survivors insurance to some couples but not to others. To mimic this, I take advantage of the relationship between survivors insurance, the date of marriage among childless, and the timing of the reform announcement. Specifically, I identify the impact of survivors insurance on family well-being exploiting variation induced by the discontinuity in time for couples that get survivors insurance if and only if they married on or before December 31, 1984. This variation allows me to use a regression discontinuity difference-in-difference estimation (RDD). The estimator identifies the average treatment effect for couples near the eligibility cut-off under specific conditions.

RDD design A regression discontinuity (difference-in-difference) design allows for identification of the impact of an endogenous regressor that is a known function of an observable assignment variable, where the assignment variable cannot be precisely manipulated (Angrist and Lavy (1999), Lee and Lemieux (2010)). In my setting, the date of marriage (*dom*) is the assignment variable and can be precisely manipulated, but the endogenous regressor – survivors insurance eligibility – is an unknown function of the observable assignment variable. Intuitively, while couples could precisely manipulate their date of marriage, it was impossible for them to manipulate the timing of marriage in response to the reform, which was announced three and a half years after these couples married. Nevertheless, the fact that the assignment variable could be precisely manipulated implies that couples on one side of the cutoff could be systematically different from those on the other.

The first panel of Figure 11 plots number of marriages in each weekly interval against distance from the survivors insurance eligibility cut-off. It shows an increase in the marriage frequency in the last week of 1984. To test for continuity in this distribution, I implement the McCrary (2008) test by collapsing the data into weekly bins and running the following regression:

$$N_b = \alpha + \beta \mathbf{1}\left[\tilde{dom}_b > 0\right] + g\left(\tilde{dom}_b\right) + \varepsilon_b, \tag{11}$$

where b indexes bins, N_b is the number of marriages in bin b, $\tilde{dom}_b$ indexes distance from the threshold in weeks, and $g\left(\tilde{dom}_b\right)$ is a polynomial. A test of $\beta = 0$ estimates whether the density is smooth. Indeed, I cannot reject the presence of a discontinuity at the eligibility threshold. This raises the concern is that "crossing" New Year's Eve has a separate effect on the outcome(s) of interest. Two features of my estimation strategy address this concern: First, to net out such an effect – provided it exists – I use an RDD design, which exploits the fact that couples that married around New Year's Eve one year earlier were unaffected by the reform. The second panel of Figure 11 shows their distribution of marriages, which is similar. Second, the timing of the announcement of the assignment rule gives me precise predictions about *when* differences in outcomes should emerge between the couples marrying close to, but on opposite sides of, January 1, 1985. Specifically, differences should emerge no earlier than three and a half years after marriage.

My estimation strategy follows that of Lalive (2008), who implements an RDD design when faced with a similar discontinuity in the density of the assignment variable. Let $Y_t = \tau_t * I_t + \sigma_t * NYE + g_t\,(dom) + U$ represent the causal relationship between the outcome of interest in time period t, Y_t, and survivors insurance status, $I_t = I_t(dom)$, where dom is the couple's date of marriage and U is a random vector of predetermined and unobservable characteristics. $NYE = NYE(dom)$ captures a potential impact of marrying after ("crossing") New Year's Eve. Given the existence of a discontinuity in insurance allocation from $t > June_{1988}$, the required identifying assumptions are that the impacts of I_t and NYE are additively separable (as above), and that, conditional on NYE, the direct marginal impact of dom on Y_t is continuous.

While all couples that marry before January 1, 1985 are certain to keep survivors insurance, couples that marry after the threshold lose survivors insurance provided that they do not have a joint child before January 1, 1990. As Figure 10 shows, only 6% of the couples that married around the threshold and did not have a joint child by the announcement date ended up having a child by January 1, 1990.[19]

[19] This is consistent with Figure 2, which shows that couples that both marry and have joint children tend to marry within a year of the first child's date of birth. Couples that married close to January 1985 and that had not conceived a joint child by June 1988 were thus likely not to ever have a joint child. Nevertheless, the reform clearly gave childless couples that married after January 1985 an incentive to have a child before January 1, 1990; this would allow them to keep their survivors insurance. I do not find any impact on fertility in this sample. One potential reason for this is that awareness

My estimation strategy involves a "fuzzy" RDD, with first stage and reduced form equations:

$$I_{it} = \alpha + \gamma 1\left[\tilde{dom_b} > 0\right]\mathbf{1}[Around85] + \delta\mathbf{1}\left[\tilde{dom_b} > 0\right] + g\left(\tilde{dom_b}\right) + h\left(\tilde{dom_b}\right)\mathbf{1}[Around85] + \epsilon_{it} \quad (12)$$

$$Y_{it} = \alpha + \beta\mathbf{1}\left[\tilde{dom_b} > 0\right]\mathbf{1}[Around85] + \eta 1\left[\tilde{dom_b} > 0\right] + i\left(\tilde{dom_b}\right) + j\left(\tilde{dom_b}\right)\mathbf{1}[Around85] + \nu_{it} \quad (13)$$

where i indexes couples, t indexes year after marriage, and $\tilde{dom_b}$ is the distance from New Year's Eve in 1985 or 1984, respectively. I include a vector of couple characteristics that is not necessary for identification but that reduces the standard errors: wedding day of week fixed effects and a spouse's educational attainment, age at marriage, age at marriage squared, and marriage parity. The RDD estimate is given by the ratio $\frac{\hat{\beta}}{\hat{\gamma}}$. The interpretation of this ratio as a causal effect requires a "monotonicity" assumption, which is satisfied here: no couple becomes eligible for survivors insurance when crossing dom_0 from above.

I test for continuity in the distributions of predetermined couple characteristics around the survivors insurance cutoff. Figure 12 plots these characteristics in each weekly interval against distance from the eligibility cut-off. I find no evidence that couples are systematically different on different sides of the cutoff.[20]

6.1 Prediction MM1: Marital instability

The first prediction for matched and married couples is that removing survivors insurance from pre-existing marriage contracts induces some couples to divorce in response to the loss of marital surplus.

Results The left panel of Figure 13 displays the empirical cumulative distribution of divorces in two subsamples: couples that entered marriage in the last three months of 1984 and in the first three months of 1985. The graph depicts the empirical estimates of the Kaplan-Meier failure (of marriage) functions, that is, the probabilities of divorce, conditional on being married, at any duration of marriage; because I include no covariates, this corresponds to the empirical distribution. Graphing the data in this way captures the average differences between couples on different sides of the cutoff at different points in time.

The figure offers graphical evidence that the removal of survivors insurance caused divorces. During the first three years of marriage, when all couples had the same marriage contract, no difference can be discerned between the functions. Upon the reform announcement in June 1988, when survivors insurance was removed for couples that married after January 1, 1985, the failure functions begin to diverge. The figure also provides an indication of the magnitude of the jump in the regression function for divorce

of the reform was not widespread immediately after the reform's announcement. Appendix Figure A1 offers suggestive evidence of this, showing that media reporting about the reform was heavily concentrated in the last three months of 1989.

[20] If I exclude the last week of 1984 and implement the McCrary (2008) test again, I find no evidence that the level of the distribution of new marriages changes discontinuously at the eligibility threshold. I have remade my analysis using a "donut regression discontinuity design," where I exploit only the 1985 cutoff but omit different parts of the last week of 1984 (and the use the cutoff around New Year's Eve 1983 as a Placebo cutoff). All results that I present in the next section remain valid. These results are available upon request.

outcomes for each of the years after the reform announcement, and suggests that the difference widens with time.

The right panel plots the same functions for couples that married within three months of another cutoff, January 1, 1984. Because both groups were unaffected by the reform, the reform should not induce a wedge between the two failure functions; indeed, the figure confirms this prediction. I find no unexpected comparable jumps in the right panel, which offers support to my interpretation of the jumps in the left panel as the causal effect of survivors insurance. Put differently, this suggests that the observed differences in the cohorts that married on opposite sides of January 1, 1985, are not driven by differences between couples that marry in different months but that they instead represent causal impacts of revoking survivors insurance.

Adding 95% confidence intervals to Figure 13, however, indicates that the difference in the left panel is statistically significant only ~20 years after marriage. Results from my RD design largely support this conjecture. Table 8 presents OLS estimates of the first stage (12), and Table 9 presents baseline 2SLS (fuzzy RD) estimates for the impact of survivors insurance on the outcome variable $divorce\_within\_t_i$, estimated at different durations of marriage t, and using polynomials $g(\tilde{dom}_b)$ of different orders and different bandwidths. One estimate suggests that removing survivors insurance raises the probability that a marriage ends in divorce within 6 years by 2.06 percentage points, but this coefficient is not significant in the specification favored by the AIC. Farther down the line, however, within 21 years of marriage, the difference in divorce rates is significant across specifications and bandwidths. The specification favored by the AIC using the smaller bandwidth suggests that the removal of survivors insurance raises the probability of divorce by 4.05 percentage points. The picture that emerges from these results is that the removal of survivors insurance pushed couples on the margin of divorce into divorce, and predominantly so closer to the annuity's expected date of realization. Indeed, with discounting, the opportunity cost of forgoing the annuity is higher, the sooner it is expected to pay out.

6.2 Prediction MM2: Division of marital surplus

The second prediction for matched and married couples is that a wife's share of household utility increases in a marriage that exists at the reform announcement and survives the reform, as compensation for the loss imposed on her in the event of her husband's death. Under the assumption that leisure is a normal good, spouses' division of market labor provides one (inverse) measure of their division of intra-household utility (Chiappori, 1992; Chiappori et al., 2002, Orrefice, 2007, 2011). Interpreting labor supply responses as an indication of a transfer of utility requires caution in this context, however. This is because labor supply may also respond to the reform for other reasons. Chief among them is that the loss of survivors insurance is a negative income or wealth shock. I exploit a non-standard feature of this wealth shock: Contrary to any shock that affects income that both spouses' *can* consume, and hence can give up when income decreases, the survivors insurance reform only affects household income in a state of the world where the husband is dead. If the spouses' labor supplies and consumption shares were to remain constant, the husband's well-being would thus be unaffected (barring altruism) but the wife's well-being would deteriorate.

This suggests that while changes in wives' labor supply in response to the reform cannot readily be interpreted as a transfer of resources within the household, an increase in husbands' labor supply can: If he works more, he gives up utility from leisure, and he does this in response to a reform that imposes a statutory loss only on the wife. Hence, this reflects a transfer of resources from husband to wife. This interoperation is consistent with, for example, the household spending resources that previously were available for joint consumption on buying a private annuity to replace the lost annuity, and the husband working more to compensate for this. Put differently, the household responds as if this were a standard wealth shock, and thus the husband bears part of its economic incidence.

My hypothesis offers an explanation for such a transfer from husband to wife, namely that he partly compensates her for the loss of survivors insurance. Note, however, that it also would be consistent with altruism (with or without bargaining); I cannot distinguish between the two.

Results *Husbands' labor supply.* Table 10 presents estimation results for the outcome variable $Y_{it} = Husb\_ls_{it}$, an indicator variable taking the value of one if the husband is working in year t, and zero otherwise. On average, I find no impacts immediately upon the reform's announcement. In the longer run, however, the estimates suggest that removing survivors insurance raises the average probability that the husband is in the labor force by circa 4 percentage points in 2004. Even though these estimates are robust across a variety of specifications, with varying polynomials and bandwidths, the magnitude of these effects suggest caution, as they imply very high responses. More careful investigation suggests that these responses need not, however, only reflect an extensive response in the traditional sense.

Specifically, these responses are largely driven by men who are very close to "retirement age." Today, there is no fixed age of retirement in Sweden, but the majority of men retires between the ages of 63 and 65. I decompose my sample of husbands into those that are older than 62 and younger than 63 in 2008 and present results for each subsample in the first row of Table 11. The left column again presents the estimated average increase in the full sample (using the baseline bandwidth and a second order polynomial), according to which the loss of survivors insurance raises the probability that a husband is working in 2008 by 3.2 percentage points. Columns two and three show that the estimated increase is smaller – and insignificant – among men who are younger than 63 in 2008; in contrast, among husbands that are older than 62, the estimated impact is larger and highly significant.

The second row of Table 11 presents analogous results for the year 2000. While the average estimated effect is insignificant, it is positive and statistically significant in the sample of husbands that are older than 62 in 2000. As the pool of men whose households lost survivors insurance ages, men who are older than 62 comprise an increasing share of the sample, which may be one reason why average responses are detectable only from 2004 and onwards. Importantly, I find no distinguishable differences in income earned before 1988 among men above and below the age of 62 (in the relevant year). The lower panel of Table 11 replicates this exercise using the age cutoff of 60. These results suggest that in response to the loss of survivors insurance, men delay their entry into retirement, which may be thought of as an intensive response along the timing-of-retirement margin.

Coming back to the hypothesis, the fact that husbands respond at all suggests that husbands behave as if the reform induced a wealth shock on the household – even though the reform only affects the household in states of the world where he is dead but leaves the surplus unchanged while he is alive. Although the statutory loss of the reform is borne by the wife alone, the economic incidence is thus partly being borne by the husband, who gives up leisure – forgoes utility – in the state where he is alive.

Wives' labor supply. To give a complete picture of household labor supply, Table 12 presents results for wives' labor supply. I find no impacts immediately upon the reform's announcement. In the longer run, however, most specifications suggest a positive impact on wives' labor market participation. The baseline specification favored by the AIC suggests that the loss of survivors insurance raises the likelihood that a wife is working in 2004 (2008) by 4.63 (4.02) percentage points. This is consistent with a number of unmodeled explanations. First, as discussed above, a wife that loses the promise of a survivors annuity experiences a wealth shock (no matter the intra-household bargaining structure). Second, as retirement looms closer – and a wife without survivors benefits soon becomes reliant on her own earned benefits – her marginal incentives to work are stronger. This is reinforced by the fact that the reform induces a marginal tax change: With survivors insurance, the wife's survivors benefits decrease dollar for dollar with her own earned benefit; without survivors insurance, no such tax applies. A third interpretation of these late effects is they reflect that women who lose the annuity stay in the labor force longer.

7 Survivors insurance and matching

While the impacts on the two groups of couples that I have hitherto analyzed represent transitory effects of the survivors insurance reform, I now turn to potential long-term impacts of this reform.

7.1 Prediction UU1: Assortativeness of matching

Because the survivors insurance tied to the old marriage contract was worth more for couples with highly unequal earnings (capacities), the old marriage contract subsidized "one-sided" unassortative matching, that is, matches between high-earning men and low-earning women. Removing the survivors insurance provision from the marriage contract is therefore predicted to have long-term impacts on matching patterns between men and women. In particular, the precise prediction concerns the *density of the share of highly skilled men that marry "down,"* that is, that marry a woman of low skill.

Sample and descriptive statistics To test this hypothesis, I begin by comparing the matching patterns of couples that choose to marry into the old and new marriage contracts. As a measure of skill, I use educational attainment at marriage, which enables me to distinguish between individuals that attended college and individuals that did not. I define college attendance as "higher education," which represents roughly 25% of all men that marry, and high school (or less) as "lower education." I use the sample of all couples with joint children that married between 1983 and 1999, a period during which the definition of educational attainment remains constant. I exclude the 6% of the observations for which I have no information about educational attainment at marriage. Table 13 displays summary statistics for this sample.

Empirical methodology I collapse the data into quarterly bins. I define the distance between a couple's quarter of marriage and the final quarter in which marriage entails take-up of the old marriage contract by $\tilde{Vigq_s} = (Vigq_s - 1989Q4)$, and estimate the following regression:

$$r_s = \alpha + \beta \mathbf{1}\left[\tilde{Vigq_s} > 0\right] + \gamma \mathbf{1}\left[\tilde{Vigq_s} > 0\right]\left(\tilde{Vigq_s}\right) + \delta \mathbf{1}\left[s = s^*\right] + g(\tilde{Vigq_s}) + \zeta_q + \varepsilon_c, \tag{14}$$

where r_s denotes the ratio of highly skilled men that marry "down,"

$$r_s = \frac{\left(\tau_h^{HIGH}, \tau_w^{LOW}\right)}{\left(\tau_h^{HIGH}, \tau_w^{LOW}\right) + \left(\tau_h^{HIGH}, \tau_w^{HIGH}\right)}.$$

The main coefficient of interest is β, which captures a discontinuous change after the threshold s^*. Further, γ captures any change in the slope at s^*, and $g(\tilde{Vigq_s})$ is a polynomial in $\tilde{Vigq_s}$.

Results Figure 14 illustrates the results. I use observations from all couples with children in which the husband has high educational attainment at marriage, collapse this data into quarterly bins, and calculate the share of marriages in which the husband married a woman of low skill ("married down"), for each quarter. I plot the relationship between this share and the date of marriage during the years 1983 to 2000, a period around the reform when my educational attainment variable definition is constant. Specifically, I plot residuals from a regression on quarter fixed effects and a dummy for the last quarter of 1989, represented by hollow circles. The black lines represent the linear fit of the share of men marrying down on the quarter of marriage, estimated separately on either side of the eligibility threshold. Finally, the dashed gray lines show the 95 percent confidence intervals. At the eligibility threshold, the figure shows a discontinuous change in the share of men marrying down. Consistent with the prediction, a larger share of highly skilled men marry down in the group that marries into the old marriage contract.

Table 14 presents the estimates from (14), using as $g(\tilde{Vigq_s})$ three different polynomials in $\tilde{Vigq_s}$. The linear model, which is favored by the AIC criterion, suggests that highly skilled men that marry into the old marriage contract are 4.47% more likely to marry a woman of low skill.

This begs the question whether the increase in assortativeness is driven by the fact that the reform induced more unassortative than assortative couples to marry. Indeed, if no unassortative matches remain unmarried, the result obtains mechanically; not by an increase in assortative matching, but by a decrease in new unassortative unions. To examine this, I plot the frequencies of new assortative and unassortative marriages in Figure 15. While the trends in assortative and unassortative matches are similar pre-reform, they diverge post-reform. Specifically, because frequency of assortative matches increases, whereas the frequency of unassortative matches slowly declines over time. Intuitively, the reform did not crowd out all new unassortative marriages because unassortatively matched couples that had a joint child before the reform's announcement constituted only a small share of all unassortative matches that considered marriage.

Pre-marital investments and the returns to education While the theory offers a precise prediction for the change in the level of assortativeness among highly skilled husbands' unions, it does not offer any prediction regarding the change in the time trend. My results suggest that after the discontinuous fall in assortative matching in 1990, the relative likelihood of an unassortative match decreases over time. One potential reason for this is that, over time, the change in marriage patterns affects the composition of high-skilled and low-skilled men and women in the marriage market. In terms of the model, this corresponds to changes in the distributions. Clearly, if I add a pre-matching investment stage when spouses invest in education, the change in the marital contract offers a greater increase in women's marginal benefit of educational investment than in men's. While a detailed analysis of this change is outside the scope of this paper, to illustrate that such impacts may be a possibility, Appendix Figure A4 shows the simple aggregate numbers of individuals entering university, by year, in Sweden in the top panel. The bottom panel splits this up by gender, with the black line depicting women, and the red line men. The upper panel indicates a large increase in overall university enrollment after 1990, which corresponded to a nationwide expansion of higher education (Bjorklund et al., 2010). The lower panel shows that the enrollment increase is greater among women. This change in enrollment is consistent with the fact that the relative returns to education were changing in the mating market.

8 Tying Survivors Insurance to Marriage

My empirical findings paint a consistent picture: In Sweden, tying survivors insurance to marriage promoted marriage, deterred divorce, and subsidized unequal matches, which in turn subsidized spousal specialization in market and domestic work. My theoretical framework provides insight into the mechanisms that drive these responses. How do such marriage market responses affect the optimal design of survivors benefits? In the U.S., Social Security played a key role in reducing absolute poverty among the elderly from 1950 to 2000 (Engelhardt and Gruber, 2004) and and in Sweden, Social Security income constituted 74% of the income of individuals above age 65 in 1994 (Palme and Svensson, 1997). In this section, I examine the implications of my findings for the social planner.

As before, I consider men and women with heterogenous income-earning abilities τ_w and τ_m. I now add a few more specifics. Outside marriage, an individual gets utility only from his or her labor income, $u(\tau_s)$, which I assume to be increasing and strictly concave, with the normalization $u(0) = 0$. Inside marriage, *gross* marital production is a function of total household labor income, $V(\tau_w + \tau_m)$. *Net* marital surplus thus is $S(\tau_w, \tau_m) = V(\tau_w + \tau_m) - \sum_{w,m} u(\tau_s)$, which I assume to be positive, as before.[21]

[21] Letting $S(\tau_w, \tau_m) = V(\tau_w + \tau_m) - u(\tau_w) - u(\tau_m)$ does not restrict $S(\tau_w, \tau_m)$ to be a function of total household income; as in Section 3, $S(\tau_w, \tau_m)$ is a function of both τ_m and τ_w. Supermodularity of $S(\cdot, \cdot)$ is preserved so long as $V(\cdot)$ is supermodular: ${}^{\partial S}/_{\partial \tau_w \partial \tau_m} = {}^{\partial V}/_{\partial \tau_w \partial \tau_m}$. Supermodularity will, however, not matter in this section, where I abstract from matching and focus on already formed couples.

The social planner's insurance objective. Following previous work on the design of social insurance against poverty (Besley and Coate (1992, 1995) and Kleven and Kopczuk (2011)), I assume that the social planner is concerned with ensuring that each individual has income above a "poverty line," denoted by z, which I normalize to 1. This *income maintenance* objective induces the social planner to transfer sufficiently much from the "rich" to the "poor" to lift the latter out of poverty. Notice that a total income of at least 4 keeps two individuals out of poverty over two periods.

Sequence of events. To focus on marriage and divorce decisions, I here abstract from matching and bargaining and consider a population of couples exogenously matched in Stage 1. In Stage 2, couples observe a stochastic marital surplus shock $\tilde{\theta}_2$ and decide whether to marry. Then, labor income is earned. In Stage 3, couples observe another stochastic marital surplus shock $\tilde{\theta}_3$, upon which married couples decide whether to divorce, and unmarried couples decide whether to marry. After this decision, men die with probability p, in which case survivors benefits may be paid to the wife. Marital surplus is generated in every period that married spouses are both alive.

Social security. The social planner uses payroll taxes to fund a social security program that may or may not include survivors benefits. For simplicity, consider a payroll tax that collects *half* of an individual's income during his or her working life (Stage 1). Without survivors insurance, each retiree has a claim in Stage 2 to the equivalent of his or her accumulated payroll taxes. Because some men die without collecting this claim, the scheme runs a surplus, and the budget is balanced by a universal lump-sum benefit b in Stage 1. With survivors insurance, I assume that a widow receives $0.5\tau_w + \max\{0.25(\tau_m - \tau_w), 0\}$ in social security benefits, whereof $\max\{0.25(\tau_m - \tau_w), 0\}$ are survivors benefits. Notice that $0.25(\tau_m - \tau_w) = 0.25(\tau_m + \tau_w) - 0.5\tau_w$ is the difference between her own social security benefits and half of the joint social security benefits that the couple would receive if the husband were alive – exactly the scheme Sweden implemented pre-reform (see equation (1) in Section 2 above). With survivors benefits, a widow's utility is thus $U_A(\tau_w, \tau_m) = u(0.5\tau_w + \max\{0.25(\tau_m - \tau_w), 0\})$. When making its marital decisions, each atomistic couple takes into account any survivors benefits the woman may receive from the social planner, but the couple does not consider the impact of its marital decisions on the universal lump-sum benefits b, which are determined by the social planner's budget constraint $b = T - E(B)$, where T is the sum of all payroll taxes and $E(B)$ is the expected social security benefits, which, by the law of large numbers, I treat as deterministic.

Traditional society. I capture a "traditional" society, where women do not participate in the labor market, by letting $\tau_w = 0$ and $\tau_m = \bar{\tau} > 4$. One can think of $\tau_m - \tau_w$ as the gender income gap. First suppose there are no survivors benefits. The couple marries in Stage 2 if $S\left({}^1\!/\!{}_2\tau_w, {}^1\!/\!{}_2\tau_m\right) + \theta_2 \geq 0$ and in Stage 3 if $pV\left(0.5\tau_w, 0.5\tau_m\right) + (1-p)u(0.5\tau_w) + \theta_2 \geq u(0.5\tau_w) + pu(0.5\tau_m)$. In the traditional society, these conditions can be written

$$\theta_t \geq \underline{\theta}^{td} \equiv u\left(0.5\bar{\tau}\right) - V\left(0.5\bar{\tau}\right) \tag{15}$$

for $t = 2, 3$. Introducing survivors benefits changes the condition for marriage in Stage 3 to

$$\theta_3 \geq \underline{\theta}_s^{td} \equiv u(0.5\bar{\tau}) - V(0.5\bar{\tau}) - \frac{(1-p)u(0.25\bar{\tau})}{p}. \tag{16}$$

Not surprisingly, survivors benefits make marriage more attractive: $\underline{\theta}_s^{td} < \underline{\theta}^{td}$. At the same time, survivors benefits eliminate old-age poverty because widows now receive social security benefits in the amount of $0.25\bar{\tau} > 1$. In regard to the social planner's budget constraint in this society, note that $T = 0.5\bar{\tau}$ and

$$E(B) = \begin{cases} (1-p)0.5\bar{\tau} & \text{without survivors benefits} \\ (1-p)0.5\bar{\tau} + p\left[1 - G_{\theta_3}(\underline{\theta}_s^{td})\right] 0.25\bar{\tau} & \text{with survivors benefits} \end{cases},$$

where $1 - G_{\theta_3}(\underline{\theta}_s^{td})$ is the share of the population that is married in Stage 3 given the existence of a survivors insurance scheme. So, the universal lump-sum benefits b are smaller in the presence of survivors benefits because benefits are being shifted to widows.

Modern society. I capture a "modern" society, in which men and women encounter more equal opportunities in the labor market but survivors benefits still matter, by letting $\tau_m > \tau_w > 2$ and $\tau_m + \tau_w = \bar{\tau}$, thus keeping total household income the same as in the traditional society. Intuitively, women earn a significant portion of the household income in this society, albeit still less than men. Without survivors benefits, the condition for marriage is

$$\theta_t \geq \underline{\theta}^{md} \equiv u(0.5\tau_w) + u(0.5\tau_m) - V(0.5\bar{\tau}) \tag{17}$$

for $t = 2, 3$. With survivors benefits, the condition in Stage 3 changes to

$$\theta_3 \geq \underline{\theta}_s^{md} \equiv u(0.5\tau_w) + u(0.5\tau_m) - V(0.5\bar{\tau}) - \frac{(1-p)u(0.25\lambda\bar{\tau})}{p} \tag{18}$$

where $\lambda\bar{\tau} = \tau_m - \tau_w$ and thus $\lambda \in (0, 1)$. Survivors benefits subsidize marriage, as before. But in this society, survivors benefits are superfluous as social insurance: the social security benefits tied to a woman's own labor income, $0.5\tau_w > 1$, are sufficient to keep her out of poverty in old age, even as a widow. The social planner's budget constraint is the same as in the traditional society, $T = \bar{\tau}$, but the expected social security benefits are now

$$E(B) = \begin{cases} (1-p)0.5\tau_m + 0.5\tau_w & \text{without surv. ben.} \\ (1-p)0.5\tau_m + 0.5\tau_w + p\left[1 - G_{\theta_3}(\underline{\theta}_s^{md})\right] 0.25\lambda\bar{\tau} & \text{with surv. ben.} \end{cases}$$

where $1 - G_{\theta_3}(\underline{\theta}_s^{md})$ is the share of the population that is married in Stage 3 given the existence of a survivors insurance scheme. As before, b must be smaller in the presence of survivors benefits.

Comparison. Survivors benefits always tend to distort marriage decisions, but their social insurance benefits are larger when the spouses' incomes are so unequal that one spouse is cast into poverty if the other dies. Survivors benefits therefore serve a greater purpose in traditional societies. A comparison of (15)-(16) with (17)-(18) flushes out these differences more clearly. First, note in (16) and (18) that, since $\lambda\bar{\tau} < \bar{\tau}$, survivors benefits increase the value of marriage on the margin more in traditional societies. But that is precisely because, in view of the large gender income gap, the social insurance is so valuable. Second, note in (15) and (17) that by Jensen's inequality, $u(0.5\tau_w) + u(0.5\tau_m) > u(0.5\bar{\tau})$ for any $\tau_w = \bar{\tau} - \tau_m > 0$, so that $\underline{\theta}^{md} > \underline{\theta}^{td}$. This means that even in the *absence* of survivors benefits, couples are much more likely to get and stay married in traditional societies than in modern societies. If most people in society marry anyway, survivors benefits have *no meaningful* impact on the marriage margin.

To see this more clearly, suppose that "happiness" shocks are bounded below so that $\tilde{\theta}_t \in (\underline{\theta}^{td}, \infty)$. In this case, all couples in the traditional society marry and none ever divorce. This provides a clean benchmark and captures a simple intuition: Women, who are not given the possibility to earn their own living in the labor market, prefer marriage – even "unhappy" ones – over living alone in poverty (on an income of $\tau_w = 0$). Nevertheless, without social insurance, a measure p of women become impoverished widows in Stage 2. This provides the rationale for survivors benefits. Moreover, as marriage is universal, survivors benefits introduce no distortion in the marriage market.

By contrast, couples in the modern society find marriage unattractive for $\tilde{\theta}_t \in (\underline{\theta}^{td}, \underline{\theta}^{md})$. Thus, not everyone marries, and some married couples divorce. This means that some couples' marital decisions are affected by survivors benefits – even though widowhood does not entail poverty. In other words, on the margin, couples that are "unhappy" choose to marry or remain married because of the financial implications of a social insurance system that no longer serves its income maintenance objective. In this modern society, survivors benefits are thus a purely distortive redistribution scheme that moves marital decisions away from the non-intervention optimum, promoting too many unions and discouraging desirable divorces. The social planner is better off abolishing survivors benefits.

Intermediate society and the social planner's trade-off. Last, I consider a society "in transition," where the population is divided into "modern" and "traditional" couples. Specifically, suppose the spouses in a fraction ϕ of couples earn "modern" wages $\tau_w^{md}, \tau_m^{md} > 2$, whereas the spouses in the remaining $1 - \phi$ couples earn "traditional" wages $\tau_w^{td} = 0$ and $\tau_m^{td} > 4$. Further assume $\tilde{\theta}_t \in (\underline{\theta}^{td}, \infty)$ such that traditional couples are always married. The social planner now faces a trade-off: weighing the social insurance benefits in one part of the population against marriage market distortions in the other.

To obtain a policy rule, I need to be more specific about the social planner's preferences. Given that income maintenance is not a strictly welfarist objective, I consider a stylized objective function that weighs "income maintenance benefits" against "marriage market distortions": Suppose the social planner places a value z on each widow "saved from poverty" and weighs these benefits against losses in "happiness" in married couples. On one hand, the social benefit of survivors insurance is then $SB = (1 - \phi)\, pz$, where $(1 - \phi)\, p$ is the measure of impoverished widows that survivors benefits would lift out of poverty. On the other hand, the social cost of survivors insurance is

$$SC = \phi \Pr[\theta_3 \in (\theta_s^{md}, \theta^{md})] \left| \left[S(\tau_w^{md}, \tau_m^{md}) - E[\theta_3 | \theta_3 \in (\theta_s^{md}, \theta^{md})\right]] \right|,$$

where $\Pr[\theta_3 \in \left(\theta_s^{md}, \theta^{md}\right)]$ is the share of modern couples whose marital decision is distorted by survivors benefits and $S(\tau_w^{md}, \tau_m^{md}) - E[\theta_3|\theta_3 \in \left(\theta_s^{md}, \theta^{md}\right)] < 0$ is the loss of "happiness" these couples suffer from this distortion *net of survivors benefits*. The social planner prefers to abolish survivors benefits when $SC > SB$, which can be written

$$\phi > \phi^* \equiv \frac{pz}{pz + \Pr[\theta_3 \in (\theta_s^{md}, \theta^{md})] \left\{ S(\tau_w^{md}, \tau_m^{md}) - E[\theta_3|\theta_3 \in (\theta_s^{md}, \theta^{md})] \right\}}.$$

The social benefit of survivors insurance is to protect $p(1-\phi)$ widows from poverty. The social cost, created by the wedge between (17) and (18), is to keep $\phi \Pr[\theta_3 \in (\theta_s^{md}, \theta^{md})]$ couples together that are "unhappy" enough to opt out of marriage in the absence survivors benefits. The higher is ϕ, the smaller are the social benefits and the larger the social costs; consequently, for high enough ϕ, survivors insurance is suboptimal. (This is only true *when* the marriage market responds; otherwise, the survivors insurance scheme entails social benefits, but no costs.)

Two main points emerge. First, whether marriage markets respond is an important factor in the optimal design of social insurance. Second, *in the presence of such marriage market responses*, the gender income gap is a key determinant for whether it is optimal, in a given society, to separate social insurance from marriage. The greater the gender income gap, the greater the share of women who do not work or earn enough in the labor market, and hence the greater the share of women who, in the absence of survivors benefits, end up as impoverished widows. If—for reasons exogenous to the design of survivors insurance— the gender income gap decreases in society, the social benefits of survivors insurance gradually vanish. Moreover, when women work outside of the household and have their own claims to retirement income, couples should no longer stay together for financial reasons when "happiness" dictates that they should part ways. As the gender income gap decreases, not only do the social benefits from survivors insurance decrease, but the costs from marriage market distortions gradually increase. At some point, as societies transition from "traditional" to "modern," it becomes optimal to decouple social insurance from marriage.

9 Conclusion

Marital status often determines social insurance eligibility. Social insurance then represents a government intervention not only into private insurance markets, but also into the marriage market. I examine how this link between social insurance and marriage affects the marriage market, exploiting a Swedish reform that eliminated *survivors insurance* – an annuity paid to widows, but not divorcees, upon the (former) husband's death. I build a model where couple formation, marital decisions, and partners' division of surplus are endogenous to the design of survivors insurance, and test it using individual-level marital and tax records. I first analyze bunching in the distribution of new marriages among unmarried couples that could take up insurance by entering marriage during a grace period. I document a significant influence on the decision between cohabitation and marriage among couples with children, with stronger responses in couples in which the male's *ex post* mortality is higher. Responses imply financial-planning horizons of up to 45 years. Among couples that entered marriage pre-reform, I then use a RDD design to show that a removal of the annuity from the marriage contract causes marital instability and a renegotiation of marital surplus in favor of the wife. Because survivors insurance replaces lost income when the husband dies, the subsidy on marital surplus is larger for unassortative matches in which the husband is of high ability and the wife is of low ability. Consistent with this, I show that the removal of survivors insurance from the marriage contract increases assortative matching of highly skilled men.

I examine the implications of my findings for a social planner and find that marriage market responses make a difference: In their presence, a social planner wishing to alleviate old age poverty faces a trade-off between, on the one hand, protecting women who do not participate in the labor market against poverty at the end of life, and distorting marriage market behavior, on the other. The gender wage gap is a key determinant for whether it is optimal, in a given society, to separate social insurance from marriage.

Many questions remain. Chief among them in my particular context is to analyze whether households' savings behavior and asset allocation responded to the survivors insurance reform. This is the natural follow-up question: I have shown that when the government ties a financial instrument to marriage that a couple otherwise would have to buy in a private insurance market, the choice between cohabitation and marriage is an integral part of a couple's long-term financial-planning strategy. Hence, the survivors insurance reform, which affected the attractiveness of marriage relative to other financial instruments, may have had an indirect effect on take-up of other insurance products. While outside of the scope of this paper, this is my next step. Another outstanding question is the converse one: Do changes in insurance markets affect behavior in the marriage market? Indeed, if linking social insurance and marriage creates spill-overs between the marriage market and private insurance markets, then any change in insurance markets could, in turn, affect the volume of marriages.

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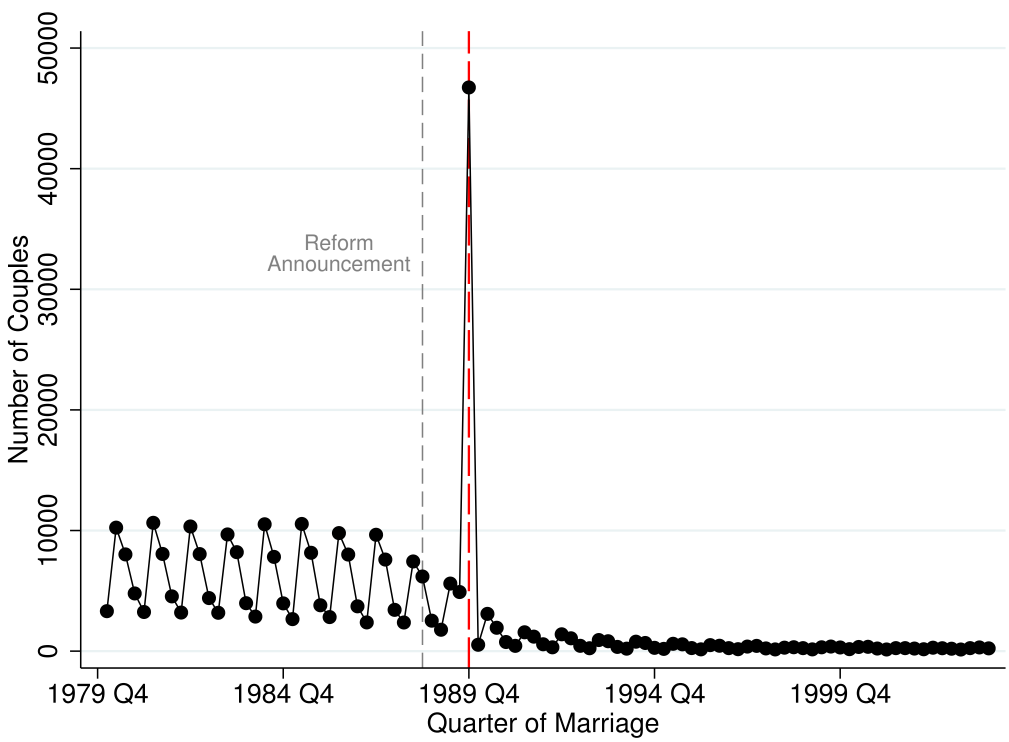
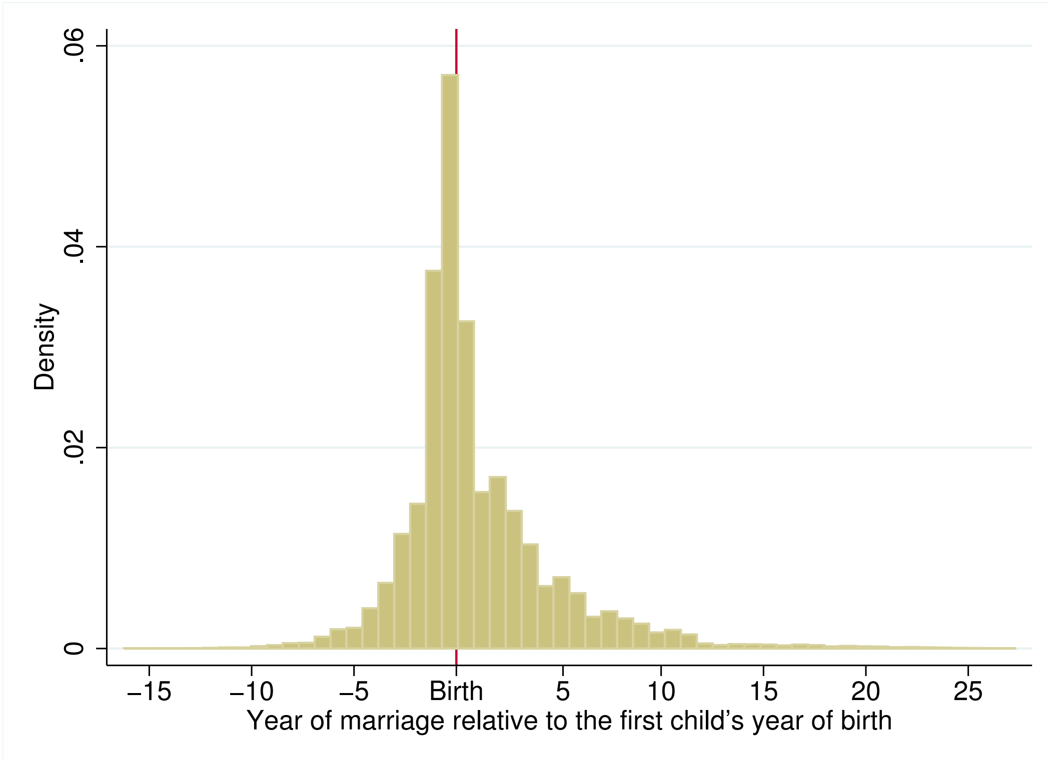
10 Figures and tables

Figure 1: Empirical Distribution of New Marriages



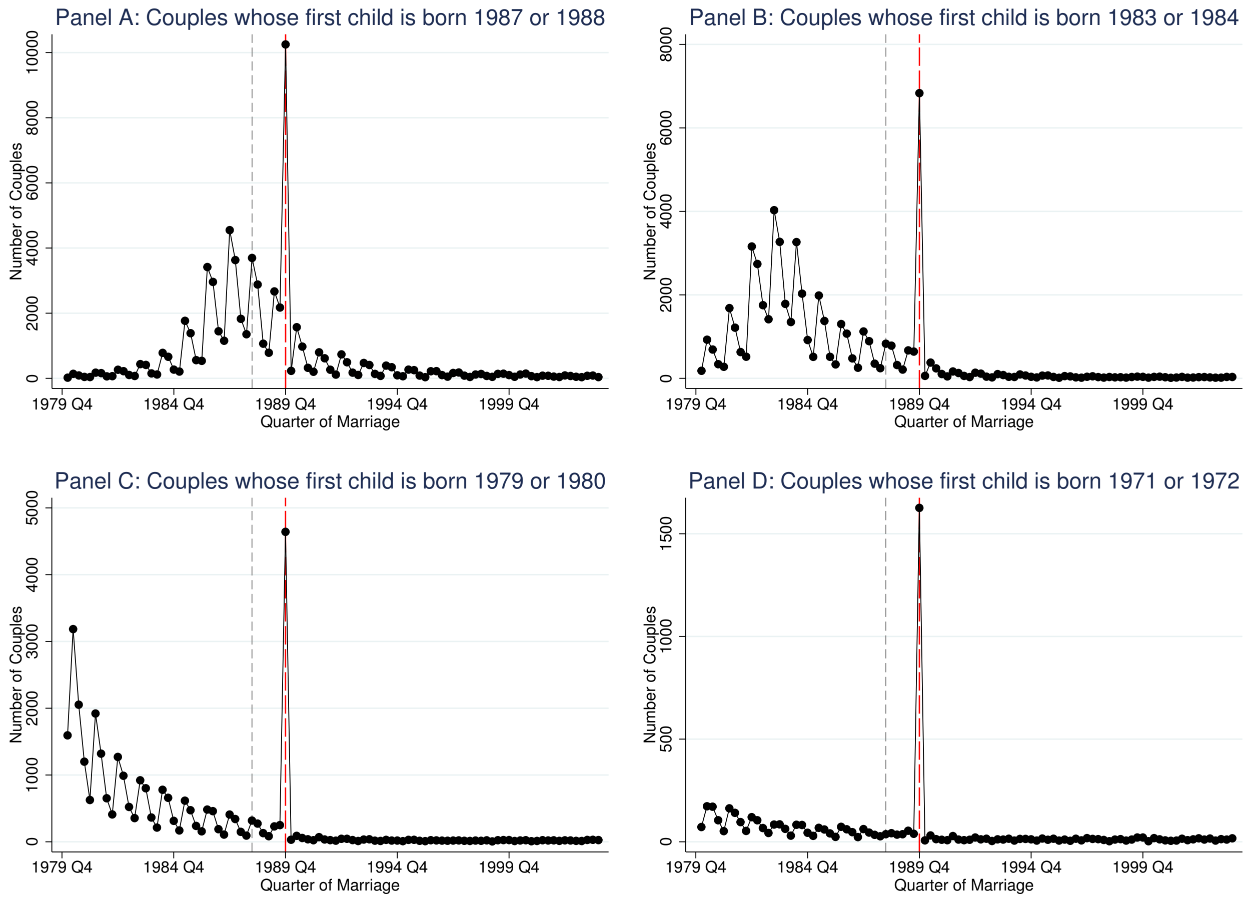
The sample includes all couples that had a joint child between January 1, 1971 and January 1, 1989. The black connected dots depict the number of marriages at a quarterly frequency, from 1980 to 2003. The green solid vertical line indicates the last quarter during which marriage entailed survivors insurance. The grey dashed vertical line indicates the quarter of reform announcement.

Figure 2: Distribution of Marriages Relative to the Birth of the First Joint Child



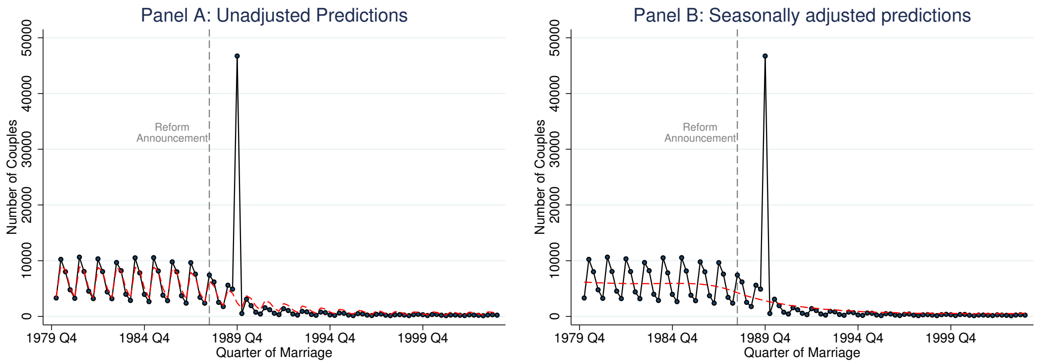
The figure plots the density of year of marriage relative to the first child's year of birth. Marriage is concentrated around the year of birth of the first joint child. The sample includes all couples that had their first joint child between January 1, 1976 and January 1, 1989. I observe marital behavior from 1969. For couples that had their first joint child in January 1976, I thus observe marital histories for seven years before childbirth. Restricting the sample to couples that had their first joint child from 1980 yields a similar graph; this is because the probability of marriage more than seven years before childbirth is close to zero.

Figure 3: Empirical Distribution of New Marriages - Specific Couple Cohorts



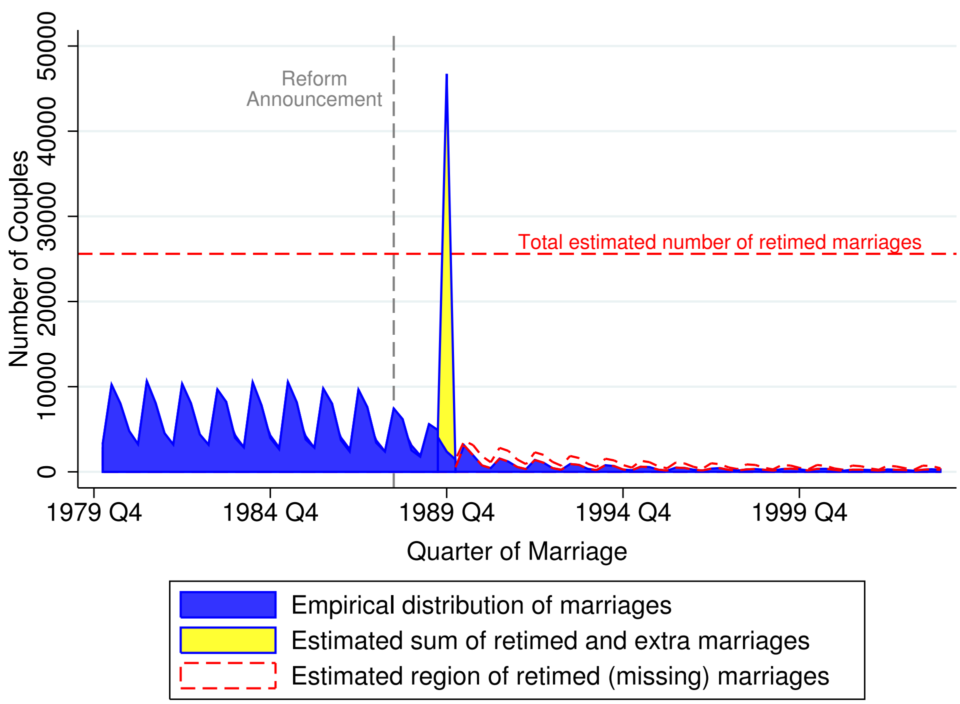
This Figure replicates Figure 1 for groups of couples. Couples are divided into cohorts depending on the date of birth of their first child. Panel A depicts new marriages among couples that had their first joint child in 1987 or 1988, the last "cohort" included in my sample. Panels B to D depicts new marriages among couples that had their first joint couples at earlier time intervals. The panels clearly illustrate that entry into marriage is concentrated around the date of birth of a couple's first joint child. Note that the scales on the four y-axes differ.

Figure 4: Predicted and Counterfactual Marriage Frequencies



The black connected line displays the empirical distribution of the marriage frequency at a quarterly frequency. The red dashed line depicts the estimated (unadjusted or seasonally adjusted) counterfactual marriage frequencies. Panel A is reproduced in Figure 5 below.

Figure 5: “Extra” and “Retimed” Marriages



I estimate the sum or retimed and extra marriages in 1989 Q4 to be 44 305 (the yellow area). Of these marriages, 25 600 are estimated to be retimed (the white area enclosed by red dashed lines; “intensive margin response”). The remainder, 18 705, are estimated to be “extra marriages” that, in the absence of reform, would not have been entered into (“extensive margin response”).

Figure 6: Heterosexual Marriage and Same-Sex Preference: Ex Ante Beliefs in Marital Decisions

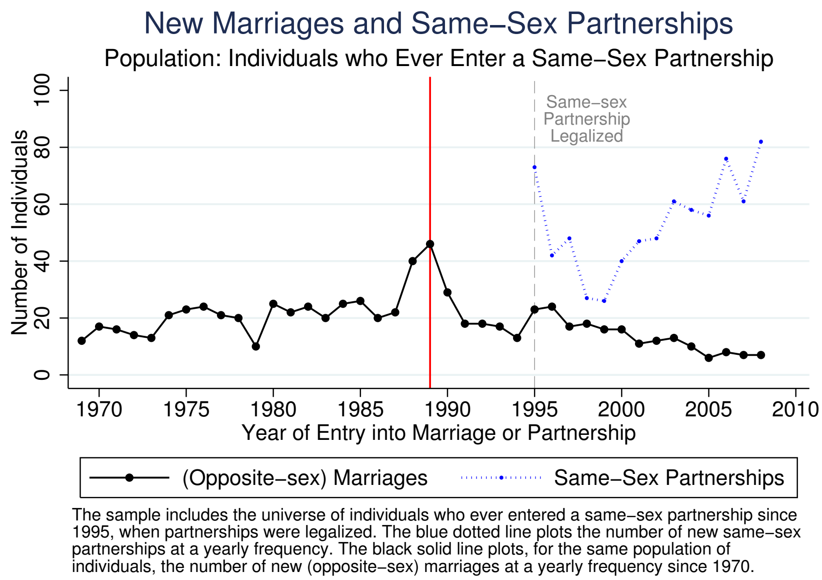
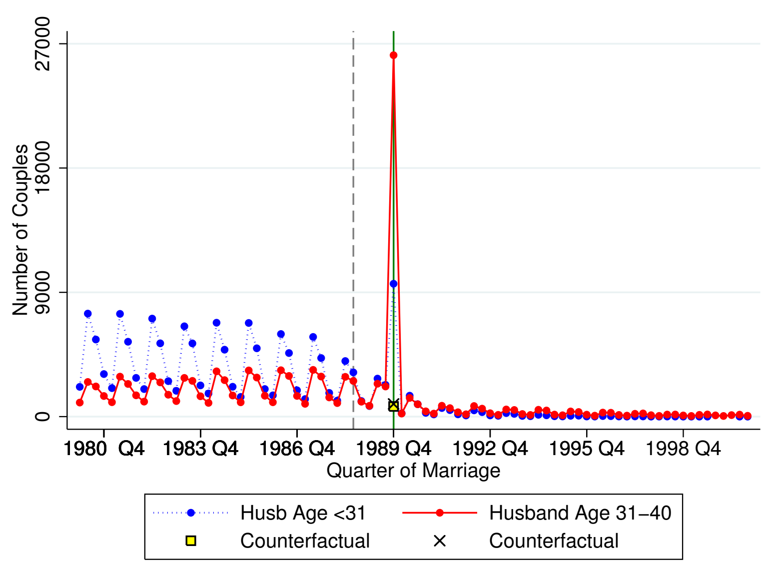
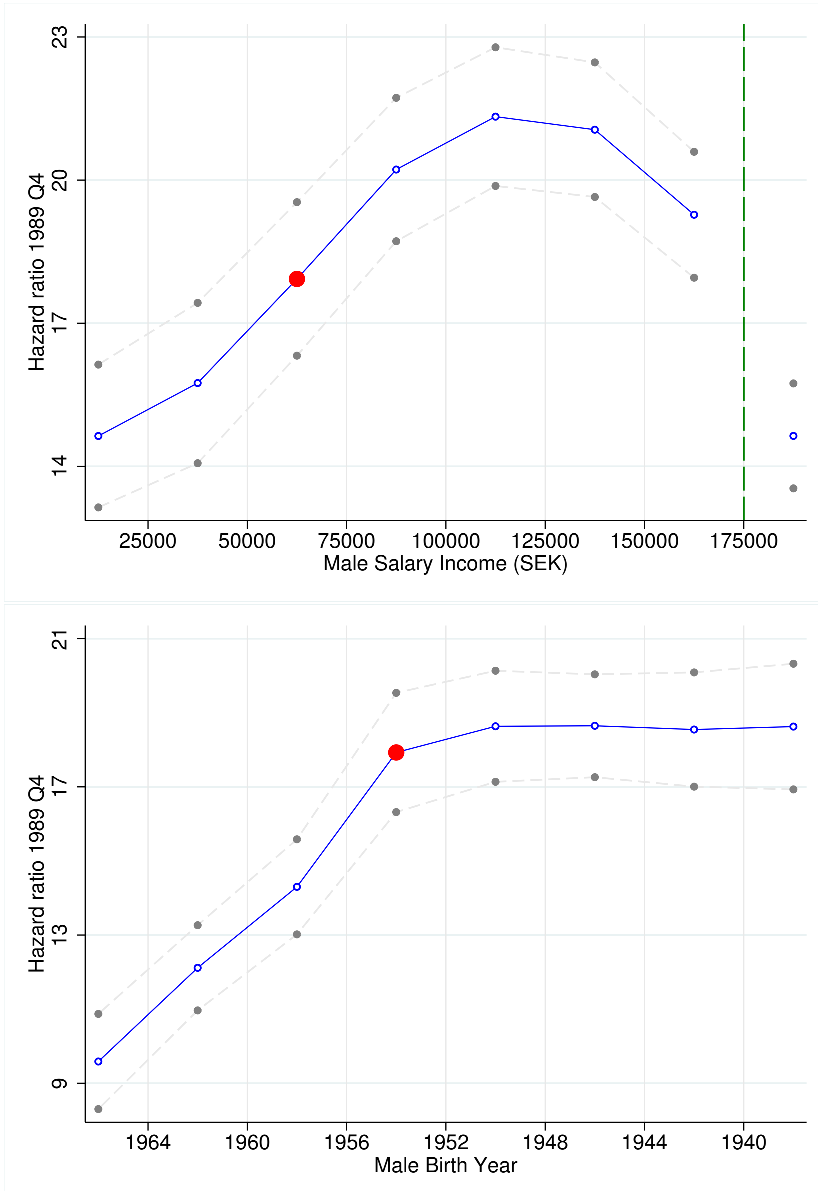


Figure 7: Heterogenous Effects w.r.t. Husband Age At Marriage



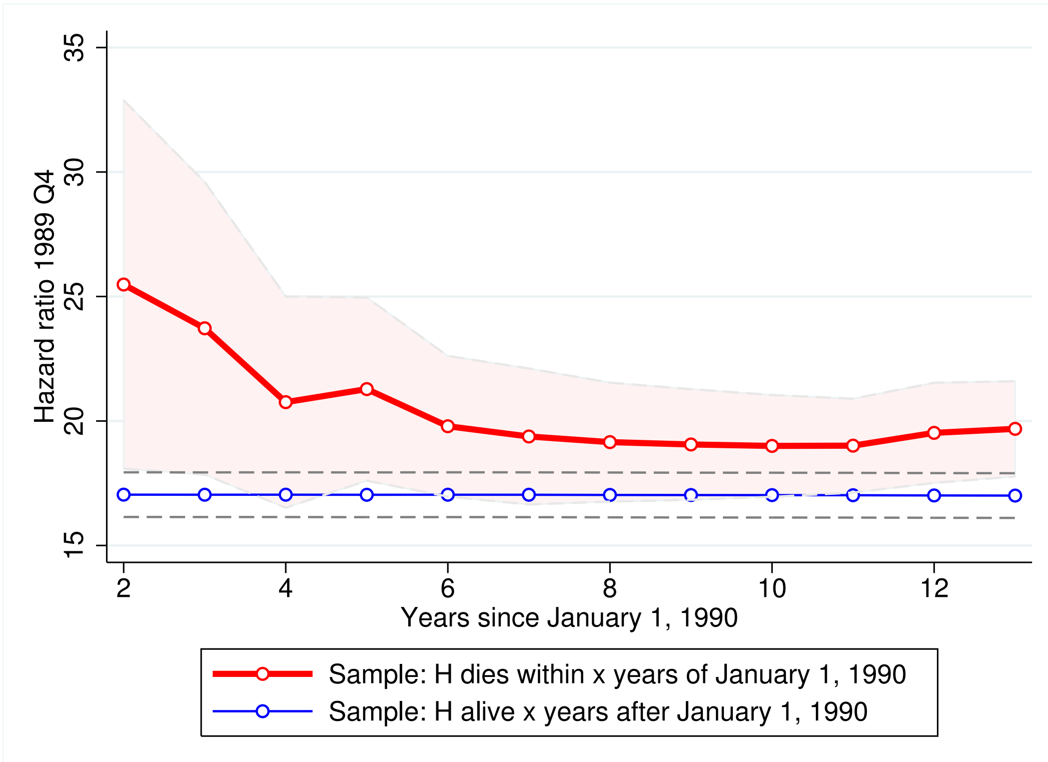
The Figure plots the distribution of new marriages at a quarterly frequency. Each point captures the number of couples that enter marriage, for each “Husband Age At Marriage” group, each quarter. The sample includes all couples who had (conceived) a joint child at the reform announcement.

Figure 8: Hazard Ratios for Couples with Different Male Income and Birth Year Ranges



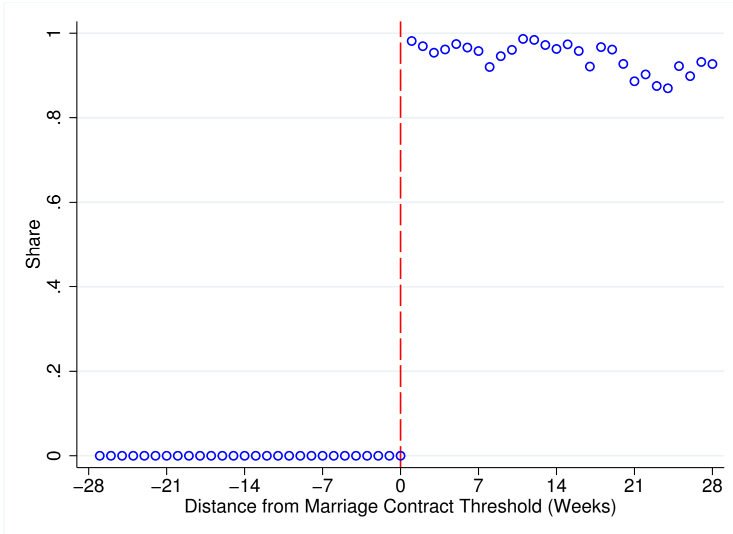
The Figure plots the estimated hazard ratios of marriage in 1989 Q4, for each male labor income group l in the upper panel and for each male birth year group b in the lower panel, obtained from estimation of (10). The specification includes two sets of interactions, between $s_i^*(t)$ and indicator variables for each group l, and between $s_i^*(t)$ and indicator variables for each group b, as well as the vectors $\mathbf{F}_i(t)$ and $\mathbf{D}_i(\mathrm{t})$. $\mathbf{F}_i(t)$ includes male income and birth year group fixed effects, the man's and woman's employment status, the woman's year of birth, and the woman's labor income. $\mathbf{D}_i(\mathrm{t})$ includes the partners' levels of education at marriage, the their marriage parities, and the couple's completed fertility. The upper panel plots the hazard ratio for couples in different male income intervals, for males born between 1952 and 1956. The last income group includes couples that exceed the Social Security limit (see Section B for details); this is indicated by the green dashed line. The lower panel plots the hazard ratio for couples in different male birth intervals, for couples with male incomes in the range SEK 50k-75k. The red dot indicates the same group, couples where the male earns labor income in the range SEK 50k-75k and was born between 1952 and 1956, captured in both panels. Gray dashed lines represent 95% confidence intervals.

Figure 9: Hazard Ratio for Different Husband *Ex Post* Mortality



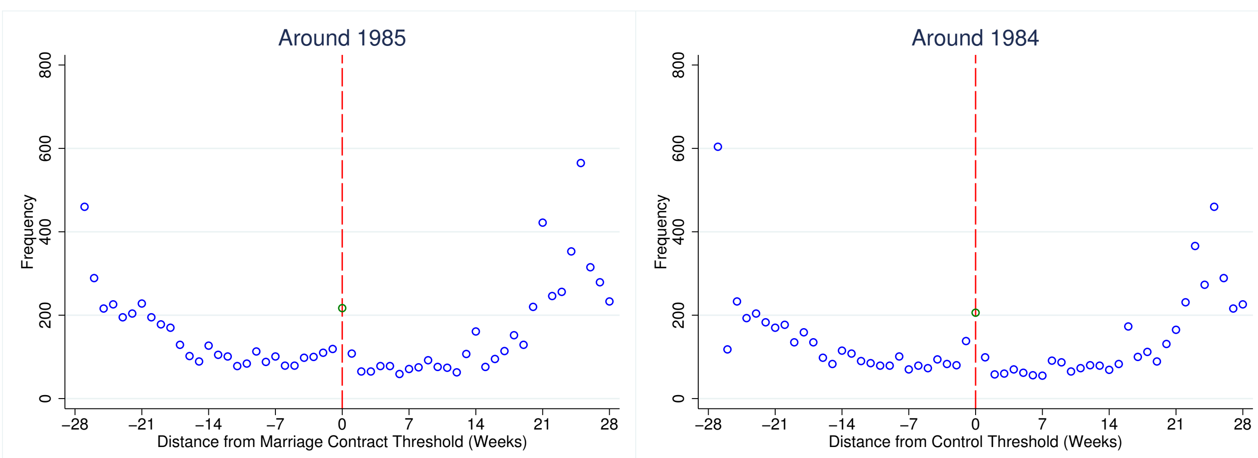
The Figure plots the estimated hazard ratio for marriage in Q4 1989 for different samples, where each sample represents couples with a specific observed husband *ex post* mortality. Each point represents a separate regression. Each point on the red line displays the hazard ratio for marriage in Q4 1989 in a sample where the husband died within a certain time frame (number of years, indicated on the x-axis) of January 1, 1990. Each point on the blue line displays the hazard ratio in a sample where the husband did not die within the same time frame (but died, or will die, later). Coefficients generating this table are reported in Appendix Table A3. Each regression includes controls for variables that, other than male age, influence the value of the insurance policy: the male's level of income, the male's share of the partners' joint income, and the female's year of birth. It further includes controls for all demographic variables that I observe – their levels of education, their marriage parities, and the couple's completed fertility. Areas represent 95% confidence intervals.

Figure 10: Distribution of treatment (=change in marriage contract) around eligibility threshold



The figure displays the share of entered marriages that experienced a switch in marriage contract due to the reform. Though my estimation strategy in practice involves a fuzzy RD, treatment is near-universal at the right side of the threshold. Among couples that married around 1985 and that had not conceived a joint child by the reform's announcement in June 1988, only 6% had a child before January 1, 1990 (and thus obtained the old marriage contract, with survivors insurance).

Figure 11: Distribution of new marriages around eligibility threshold



The figure displays the frequency of new marriages, in weekly bins, around January 1, 1985 and January 1, 1984. The green circles represents each year's last week.

Figure 12: Distribution of covariates around eligibility threshold (1985)

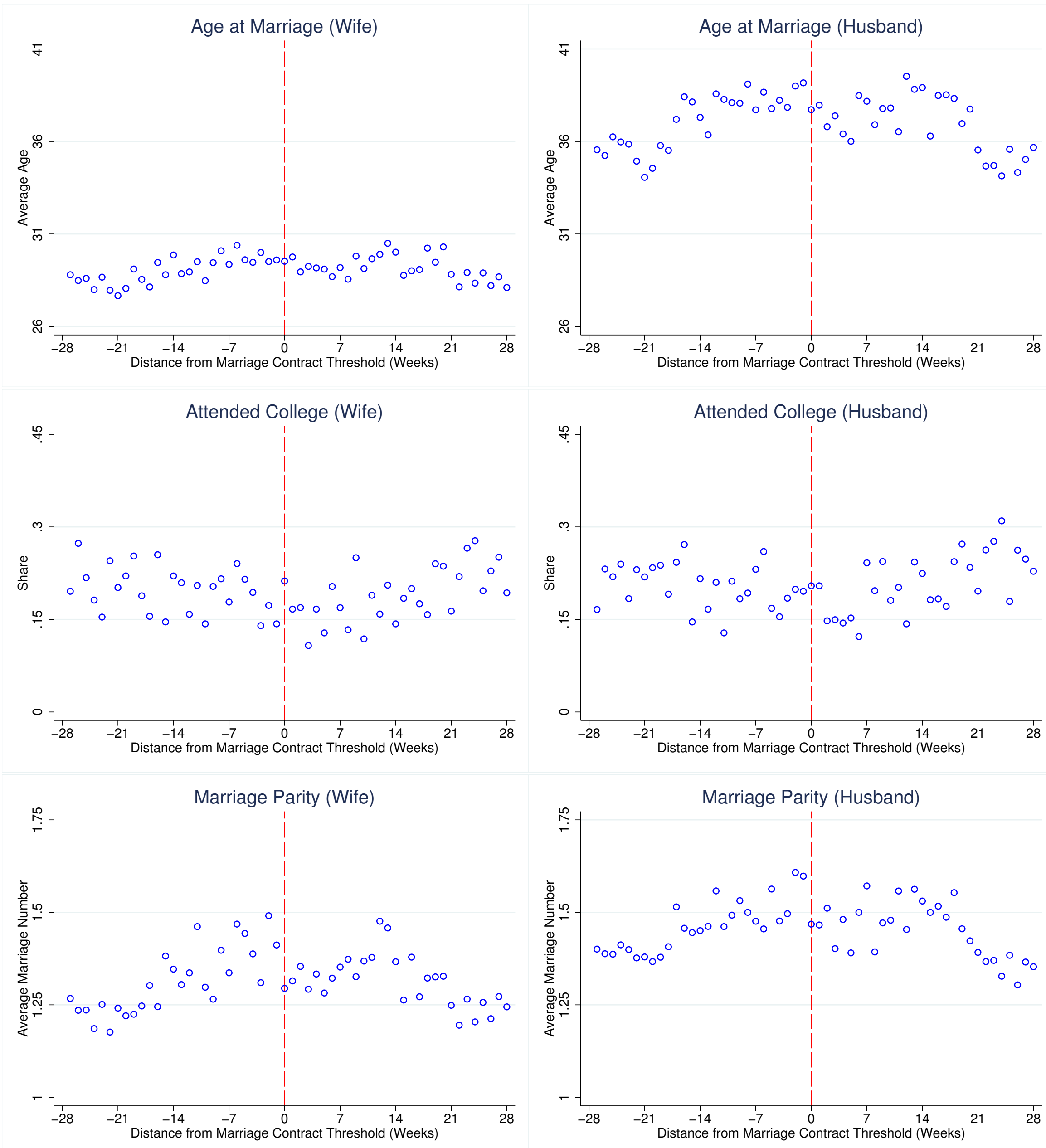
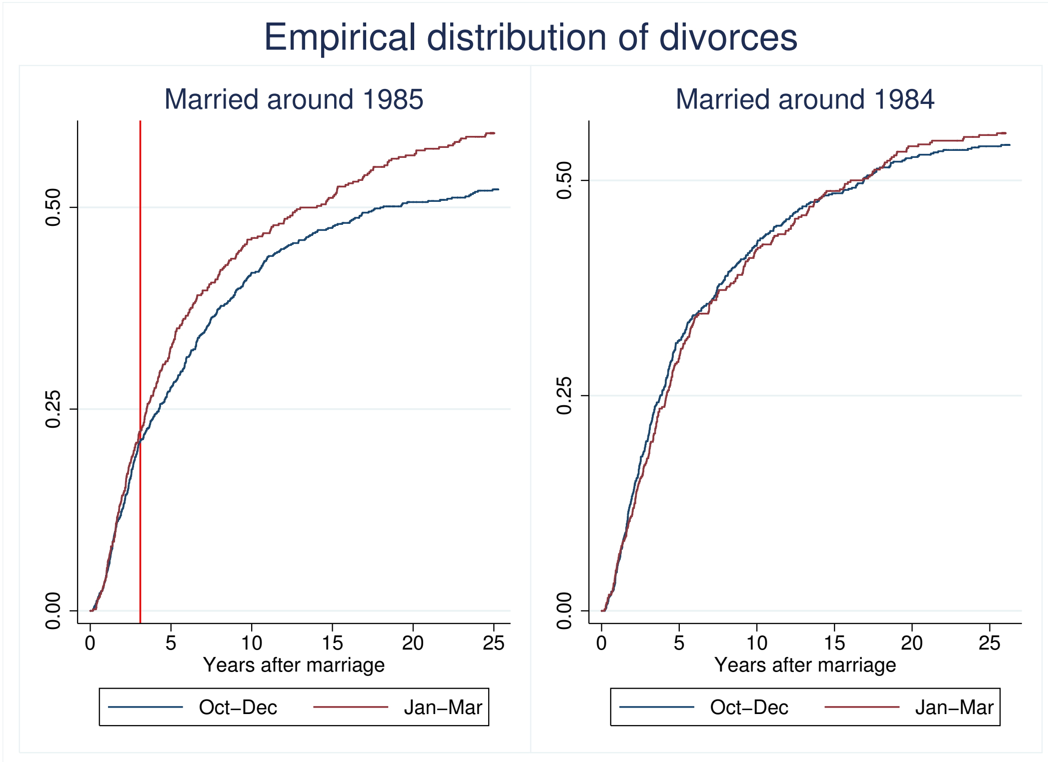
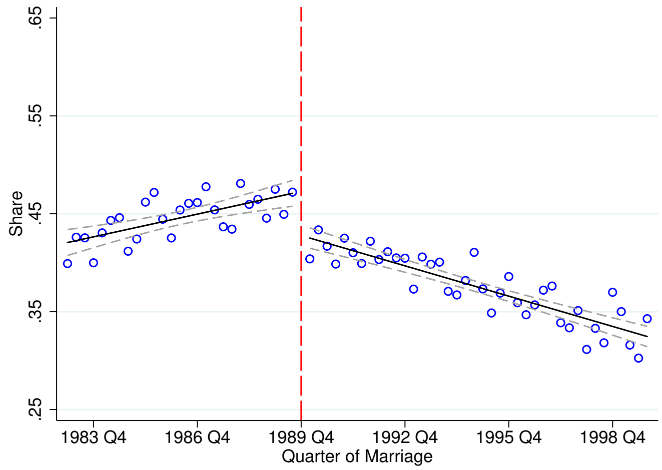


Figure 13: Empirical cumulative divorce function around eligibility threshold and placebo threshold



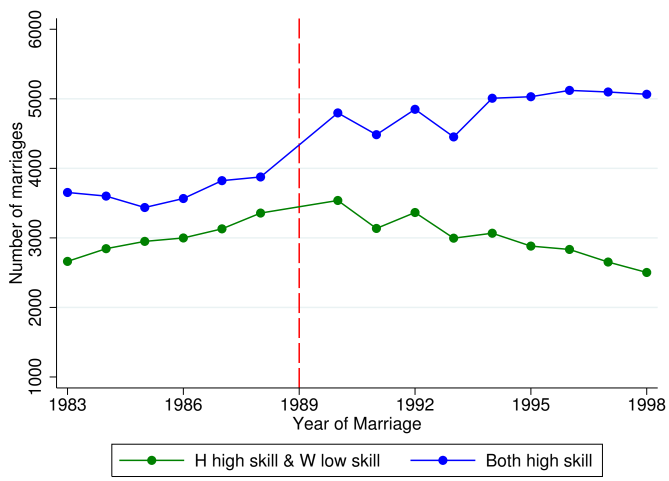
The graph depicts the empirical cumulative divorce function obtained by estimating the Kaplan-Meier survival function without including any covariates, for couples marrying around 1985 (left panel) and around 1984 (right panel), respectively. Until the announcement of the reform in June 1988, all four groups of couples had the same marriage contract. Upon the reform announcement in June 1988, the old marriage contract was replaced with the new contract – without survivors benefits – for couples that married after January 1, 1985, depicted by the red solid line in the left panel. All couples in the right panel were allowed to keep the old marriage contract that they married into.

Figure 14: Marriages: Impact on the Share of Highly Skilled Men Marrying a Low Skill Woman



Quarterly bins. The hollow circles depict the share of highly skilled men marrying a woman of low skill (seasonality adjusted). The black solid lines represent linear fits of the share of highly skilled men marrying a woman of low skill on quarter of marriage, estimated separately on each side of the cut-off. Gray dashed lines represent 95 percent confidence intervals.

Figure 15: Marriages: The Number of High-Low and High-High Matches



The sample includes all couples that had a joint child and married between 1983 and 2000 (a time period during which the definition of educational attainment at marriage is constant around the 1989 threshold), but omitting the 1989, when marital behavior responded to the reform. The table displays the number of couples where the husband has a high educational attainment at marriage and the wife has either a low educational attainment at marriage (green solid line) or a high educational attainment at marriage (blue solid line).

Table 1: Summary Statistics for Matched but not Married Sample

	Main Sample	Old Marriage Contract	
Year of marriage	1980-2003	1980-1988	1989
Demographic characteristics at marriage			
H age at marriage	31.49	29.82	35.12
W age at marriage	28.72	27.05	32.24
H education (4 lvls)	2.08	2.13	1.97
W education (4 lvls)	2.10	2.13	2.02
H cognitive capacity	0.03	0.09	-0.13
H marriage number	1.12	1.12	1.11
W marriage number	1.11	1.11	1.09
H is immigrant	0.07	0.08	0.05
W is immigrant	0.08	0.08	0.04
Economic characteristics at marriage			
Total household labor income (H and W)	10.26	10.09	12.08
H income share	0.66	0.66	0.70
Fertility behavior			
Couple's completed fertility	2.26	2.26	2.22
First child out of wedlock	0.66	0.52	1.00
Number of couples	306823	220112	59012

Note: The sample includes all couples that had a joint child between January 1, 1971 and December 31, 1988, that married between 1980 and 2003. By marrying before January 1, 1990, these couples opted into the old marriage contract, to which survivors benefits were tied.

Table 2: Marriage Boom Estimates

	Polynomial					
	2	3	4	5	6	7
Induced Marriages	45348	46195	46751	46634	46641	46745
Std error	(667)	(630)	(688)	(801)	(963)	(1192)
T-statistic	(68)	(73)	(68)	(58)	(48)	(39)
AIC	627.5	617.0	615.3	617.2	619.2	621.1
Number of obs (pre-reform quarters)	40	40	40	40	40	40
Number of couples	279124	279124	279124	279124	279124	279124

Note: Each column represents a separate regression, with a polynomial of the indicated degree. Estimates are obtained using only pre-reform data (all quarters before 1990 Q1). "Induced marriages" reports the estimated coefficient on $\mathbf{1}[s = s^{*}]$, which takes the value of one in 1989 Q4, with robust standard errors in parentheses. Each regression includes four quarter fixed effects. Instead of indicating significance with stars, I report standard errors and, for the number of induced marriages, t-statistics, which imply $p < 10^{-9}$ for all induced marriage estimates.

Table 3: Impact on Marriage

	Polynomial		
	2	3	4
Induced Marriages	44065***	44396***	44305***
	(3261)	(3238)	(3229)
Which consists of			
Retimed Marriages	23113***	21646***	25600***
	(4496)	(2748)	(3176)
Extra Marriages	20951***	22750***	18705***
	(2685)	(1931)	(1629)
Estimated Counterfactual	2668***	2337***	2428***
	(385)	(327)	(323)
$\Delta p/p$	16.51***	19.00***	18.25***
	(2.0029)	(2.1566)	(1.8862)
Number of obs (cohort quarters)	6473	6473	6473
AIC	9186	8928	8880

Note: Standard errors obtained from cluster bootstraping procedure (further described in the text) with 10 000 randomly drawn panels in parentheses. Each column represents a regression with polynomials of the indicated degree. Each regression also includes cohort fixed effects, (four) quarter fixed effects, and cohort-specific increases (decreases) in entry into marriage at (after) the eligibility threshold. AIC is obtained from the regression with the entire (non-random) sample.

* p < 0.05, ** p < 0.01, *** p < 0.001.

Table 4: Impact on Marriage: Estimated Hazard Ratios

	Financial controls		All observable controls	
	Coefficient, $\hat{\beta}$	Hazard rate, $e^{\hat{\beta}}$	Coefficient, $\hat{\beta}$	Hazard rate, $e^{\hat{\beta}}$
1989 Q4	2.83***	16.88***	2.84***	17.04***
	(0.03)	(0.52)	(0.03)	(0.55)
Flexible controls	NO	NO	NO	NO
1989 Q4	2.84***	17.05***	2.84***	17.08***
	(0.03)	(0.53)	(0.03)	(0.55)
Flexible controls	YES	YES	YES	YES
Number of couples	247194	247194	218278	218278

Note: Columns 1 and 3 report the estimated coefficient on the indicator variable for 1989 Q4, with standard errors clustered at couple cohort in parentheses. Columns 2 and 4 report the corresponding hazard ratios. Significance levels from a test of the null hypotheses that each regression coefficient is 0 or, equivalently, that each hazard ratio is 1. In the upper panel, all regressions include controls for financial characteristics $\mathbf{F}_i(t)$ that influence the annuity's expected value: Each partner's employment status, the man's total labor income and share of household labor income, and the partners' birth years. In columns 3 and 4, each regression also includes controls for other characteristics $\mathbf{D}_i(t)$ that I observe: The partners' education levels, their number of children, and the spouses' marriage parities. In the lower panel, all controls are similar except that I control flexibly for male labor income and birth year, including $\tilde{\mathbf{F}}_i(t)$ instead of $\mathbf{F}_i(t)$.

* $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$.

Table 5: Impact on Marriage: Heterogenous Effects w.r.t. Male *Ex Post* Mortality

	Financial controls		All observable controls	
Estimates from sample	Coefficient, $\hat{\beta}$	Hazard ratio, $e^{\hat{\beta}}$	Coefficient, $\hat{\beta}$	Hazard ratio, $e^{\hat{\beta}}$
1989 Q4	2.83***	16.88***	2.84***	17.04***
	(0.03)	(0.52)	(0.03)	(0.55)
Interaction	0.12**	1.13**	0.13**	1.14**
	(0.06)	(0.07)	(0.06)	(0.07)
Number of couples	247046	247046	218147	218147

	Financial controls		All observable controls	
Estimates from sample	Coefficient, $\hat{\beta}$	Hazard rate, $e^{\hat{\beta}}$	Coefficient, $\hat{\beta}$	Hazard rate, $e^{\hat{\beta}}$
Husband dies within 5 years	3.00***	20.12***	3.06***	21.28***
	(0.10)	(1.93)	(0.11)	(2.24)
Number of couples	1214	1214	1051	1051
Husband alive after 5 years	2.83***	16.88***	2.84***	17.04***
	(0.03)	(0.52)	(0.03)	(0.55)
Number of couples	245832	245832	217096	217096

Note: Columns 1 and 3 report the estimated coefficient on the indicator variable for 1989 Q4, with standard errors clustered at couple cohort in parentheses. Columns 2 and 4 report the corresponding hazard ratios. Significance levels from a test of the null hypotheses that each regression coefficient is 0 or, equivalently, that each hazard ratio is 1. The upper panel reports results from a regression including $s_i^*(t)$, an indicator variable taking the value one if the man dies within five years of January 1, 1990, and their interaction term. The lower panel reports results from separate estimations of the baseline regression equation in the two male *ex post* mortality samples. All regressions include controls for financial characteristics $\mathbf{F}_i(t)$ that influence the annuity's expected value: Each partner's employment status, the man's total labor income and share of household labor income, and the partners' birth years. In columns 3 and 4, each regression also includes controls for other characteristics $\mathbf{D}_i(t)$ that I observe, and that the price of a (private) annuity potentially could be made contingent on: The partners' education levels, their number of children, and the spouses' marriage parities.

* $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$.

Table 6: Hightened Divorce Risk in Rushed Marriages: OLS Results

	Dependent variable: Divorce within			
	3 years	5 years	10 years	15 years
Married in 1989Q4	0.0075**	0.0228***	0.0423***	0.0393***
	(0.0027)	(0.0048)	(0.0088)	(0.0091)
Demographic characteristics at marriage				
H age at marriage	-0.0007**	-0.0010**	-0.0022***	-0.0028***
	(0.0002)	(0.0003)	(0.0005)	(0.0005)
W age at marriage	-0.0031***	-0.0068***	-0.0112***	-0.0125***
	(0.0002)	(0.0003)	(0.0004)	(0.0005)
H cognitive capacity	-0.0065***	-0.0111***	-0.0177***	-0.0192***
	(0.0007)	(0.0008)	(0.0014)	(0.0014)
H second marriage	0.0206***	0.0413***	0.0909***	0.1011***
	(0.0036)	(0.0049)	(0.0074)	(0.0078)
W second marriage	0.0175***	0.0388***	0.0821***	0.1111***
	(0.0030)	(0.0049)	(0.0069)	(0.0064)
H is immigrant	0.0066	0.0115	0.0208*	0.0320**
	(0.0040)	(0.0068)	(0.0083)	(0.0105)
W is immigrant	0.0068*	0.0110*	0.0063	0.0042
	(0.0029)	(0.0049)	(0.0053)	(0.0065)
Economic characteristics at marriage				
Total household labor income (H and W)	-0.0115***	-0.0251***	-0.0368***	-0.0395***
	(0.0011)	(0.0020)	(0.0028)	(0.0030)
H income share	0.0105**	0.0403***	0.0694***	0.0717***
	(0.0037)	(0.0068)	(0.0092)	(0.0129)
W earns much more	0.0032	0.0107**	0.0240***	0.0245***
	(0.0021)	(0.0033)	(0.0050)	(0.0055)
H earns much more	0.0028	0.0062	0.0087	0.0123*
	(0.0018)	(0.0033)	(0.0048)	(0.0051)
Total joint number of children				
Couple's completed fertility	-0.0295***	-0.0622***	-0.0940***	-0.0896***
	(0.0010)	(0.0025)	(0.0048)	(0.0060)
Constant	0.2587***	0.5254***	0.9487***	1.1679***
	(0.0238)	(0.0456)	(0.1696)	(0.2124)
Number of obs	94681	94681	94681	94681

Note: Each column represents a regression with a different dependent variable. All regressions include wedding month fixed effects, wedding day of week fixed effects, and four educational level indicators for each spouse. Standard errors clustered at (wedding month*wedding day of week) in parentheses.

* $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$.

Table 7: Summary Statistics: Sample of Already Married Couples

	Samples	
	Married around 1985	Married around 1984
Demographic characteristics		
H age at marriage	36.47	36.51
W age at marriage	28.95	28.66
H birth year	1948	1947
W birth year	1956	1955
H education (4 lvls)	2.08	2.05
W education (4 lvls)	2.08	2.08
H marriage number	1.43	1.43
W marriage number	1.28	1.28
Financial characteristics		
H labor income at marriage	94262.10	94695.71
W labor income at marriage	60675.38	60469.76
Number of individuals	14845	12941

Note: The sample includes all couples that married within 180 days of January 1, 1985 and within 180 days of January 1, 1984.

Table 8: Results: First Stage

	Polynomial	
	2	3
New Marriage Contract Obtained	0.9775***	0.9773***
	(0.0062)	(0.0090)
AIC	-8897.76	-8897.76
Number of couples	13890	13890

Note: The sample includes all couples that married within 180 days of January 1, 1985 and within 180 days of January 1, 1984. Robust standard errors in parentheses.

* $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$.

Table 9: IV (RDD) estimates: Impact on divorce

	Bandwidth 150 days		Bandwidth 180 days	
	Polynomial		Polynomial	
Divorce within	2	3	2	3
3 years	0.0189	0.0067	0.0053	0.0070
	(0.0214)	(0.0232)	(0.0182)	(0.0201)
AIC	7426.27	7427.92	10650.14	10653.91
6 years	0.0256	0.0206*	0.0096	0.0019
	(0.0260)	(0.0121)	(0.0224)	(0.0248)
AIC	10975.35	10977.87	16405.37	16403.73
12 years	0.0135	0.0063	0.0010	-0.0052
	(0.0279)	(0.0303)	(0.0243)	(0.0269)
AIC	12248.80	12251.64	18647.23	18646.19
21 years	0.0405**	0.0337**	0.0236**	0.0218*
	(0.0218)	(0.0184)	(0.0125)	(0.0141)
AIC	12560.66	12564.26	19391.08	19388.23
24 years	0.0373**	0.0285**	0.0251*	0.0256**
	(0.0192)	(0.0152)	(0.0158)	(0.0142)
AIC	12571.18	12574.63	19427.76	19425.61
Number of couples	9040	9040	13896	13896

Note: Each cell represents a separate 2SLS regression. Robust standard errors in parentheses.
* $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$.

Table 10: IV (RDD) estimates: Impact on husband's labor supply (extensive margin)

	Bandwidth 150 days		Bandwidth 180 days	
	Polynomial		Polynomial	
Husband working in	2	3	2	3
1988	0.0133	0.0166	0.0158	0.0132
	(0.0171)	(0.0186)	(0.0146)	(0.0161)
AIC	10167.17	10170.21	13941.33	13943.32
1994	0.0017	0.0034	0.0167	0.0086
	(0.0209)	(0.0226)	(0.0181)	(0.0199)
AIC	15277.56	15281.15	22194.02	22196.38
2000	0.0239	0.0222	0.0319*	0.0279
	(0.0202)	(0.0219)	(0.0175)	(0.0193)
AIC	14452.73	14454.45	20976.05	20979.74
2004	0.0427**	0.0403*	0.0447**	0.0471**
	(0.0203)	(0.0219)	(0.0177)	(0.0194)
AIC	14495.58	14499.42	21244.31	21246.44
2008	0.0329*	0.0236	0.0392**	0.0351*
	(0.0195)	(0.0211)	(0.0171)	(0.0188)
AIC	13495.26	13496.24	19987.82	19991.47
Number of couples	9036	9036	13890	13890

Note: Each cell represents a separate 2SLS regression. Robust standard errors in parentheses. * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$.

Table 11: Impacts on Husband Labor Supply Driven by Men Close to Retirement

	All couples	Subsample of couples: Husband age	
		H younger than 62	H older than 62
Husband working in 2008	0.0392**	0.0222	0.0635***
	(0.0171)	(0.0242)	(0.0209)
Number of couples	13890	8309	5580
Husband working in 2000	0.0319*	0.0222	0.0719***
	(0.0175)	(0.0242)	(0.0257)
Number of couples	13890	11577	2313

	All couples	Subsample of couples: Husband age	
		H younger than 60	H older than 60
Husband working in 2008	0.0392**	0.0324	0.0546**
	(0.0171)	(0.0256)	(0.0208)
Number of couples	13890	8309	5580
Husband working in 2000	0.0319*	0.0149	0.0931***
	(0.0175)	(0.0207)	(0.0286)
Number of couples	13890	11577	2313

Note: Each cell represents a separate 2SLS regression using a second order polynomial and a bandwidth of 180. The left column presents results for the whole male sample. Columns two and three present results from estimation on two subsamples (defined separately for 2000 and 2008). Robust standard errors in parentheses.

Table 12: IV (RDD) estimates: Impact on wife's labor supply (extensive margin)

	Bandwidth 150 days		Bandwidth 180 days	
	Polynomial		Polynomial	
Wife working in	2	3	2	3
1988	-0.0137	-0.0078	-0.0109	-0.0117
	(0.0205)	(0.0223)	(0.0171)	(0.0189)
AIC	6709.15	6711.84	8879.42	8878.37
1994	0.0035	0.0221	0.0160	0.0212
	(0.0245)	(0.0266)	(0.0208)	(0.0230)
AIC	9884.96	9884.64	14366.56	14367.92
2000	0.0270	0.0297	0.0365*	0.0381
	(0.0253)	(0.0275)	(0.0215)	(0.0237)
AIC	10483.13	10487.05	15204.30	15208.20
2004	0.0537**	0.0580**	0.0463**	0.0459*
	(0.0259)	(0.0281)	(0.0220)	(0.0243)
AIC	10888.23	10892.03	15922.76	15926.19
2008	0.0383	0.0429	0.0402*	0.0508**
	(0.0264)	(0.0287)	(0.0226)	(0.0250)
AIC	11277.50	11281.30	16610.37	16613.16
Number of couples	9040	9040	13896	13896

Note: Each cell represents a separate 2SLS regression. Robust standard errors in parentheses.
* $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$.

Table 13: Summary Statistics: Marriages of Highly Skilled Men 1983 - 2000

	Old Marriage Contract		New Contract
Quarter of Marriage	1983Q1-1989Q3	1989Q4	1990Q1-1999Q4
H high skill but W low skill			
H age at marriage	30.67	36.22	31.43
W age at marriage	27.71	33.24	28.52
W education (4 lvls)	1.85	1.81	1.91
H cognitive capacity	0.75	0.55	0.67
Total household labor income (H and W)	12.04	12.28	7.82
H income share	0.70	0.73	0.68
Couple's completed fertility	2.21	2.13	2.09
First child out of wedlock	0.39	0.91	0.44
Observations	21124	5044	29752
Both H and W high skill			
H age at marriage	31.77	37.78	31.88
W age at marriage	29.46	35.29	29.61
W education (4 lvls)	3.51	3.49	3.51
H cognitive capacity	0.87	0.64	0.86
Total household labor income (H and W)	12.21	12.43	8.05
H income share	0.63	0.68	0.61
Couple's completed fertility	2.27	2.20	2.16
First child out of wedlock	0.32	0.91	0.30
Observations	25659	4936	49879

Note: The sample includes all couples that had a joint child and married between 1983 and 2000 (a time period during which the definition of educational attainment at marriage is constant around the 1989 threshold). The table displays summary statistics for couples where the husband has a high educational attainment at marriage, separately depending on whether the husband married a woman of low (upper panel) or high (lower panel) skill.

Table 14: Matching: Impact on the Share of Highly Skilled Men Marrying Low Skilled Women

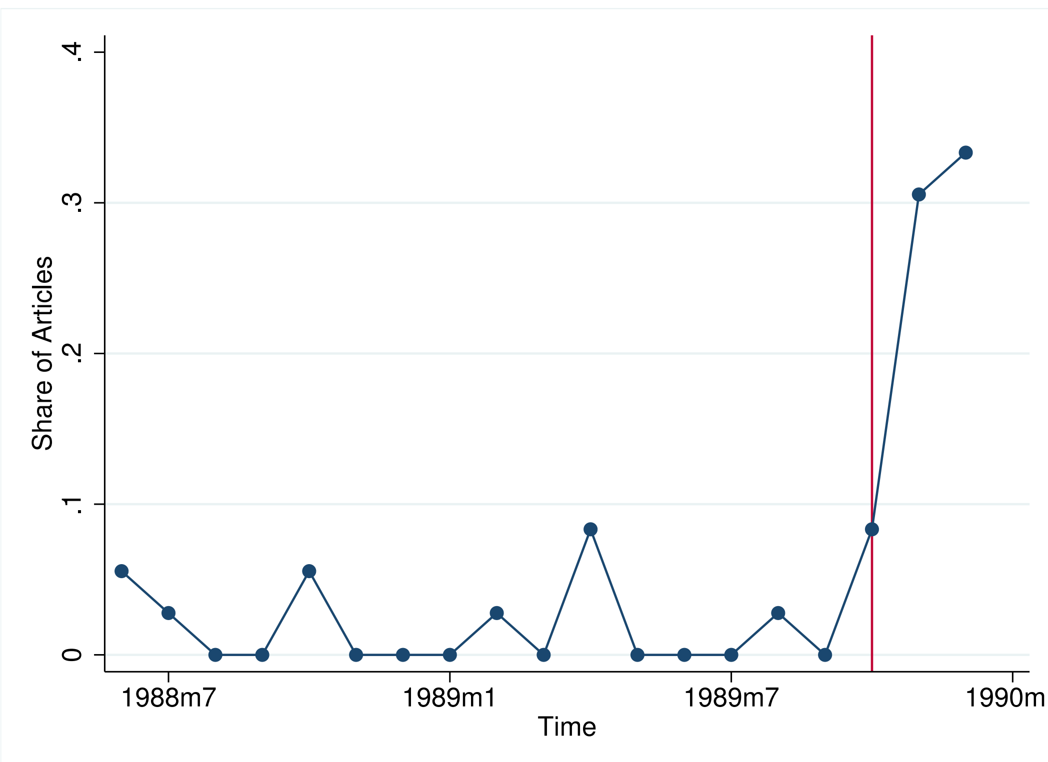
	Polynomial		
	1	2	3
New Marriage Contract	-0.0447***	-0.0364**	-0.0492*
	(0.0087)	(0.0135)	(0.0194)
Adj. R squared	0.87	0.88	0.87
AIC	-353.65	-353.06	-350.20
Number of obs	68	68	68

Note: Each column represents a separate regression with different polynomials in time. All regressions include quarter fixed effects. Robust standard errors in parentheses. The AIC criterion favors the linear model.

* $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$.

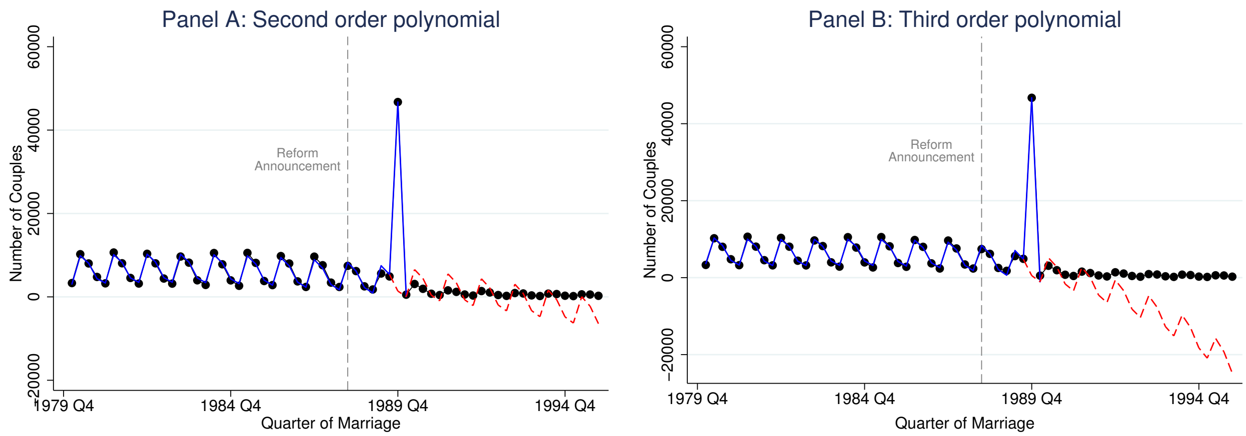
A Supplemental figures and tables

Figure A1: Media Attention and Reform Salience: The Distribution Over Time of Newspaper Articles Mentioning the Reform



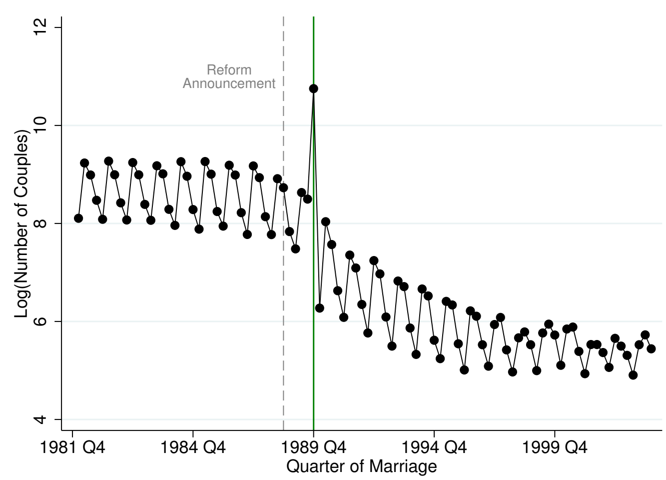
The sample includes all articles published by Tidningarnas Telegrambyrå, Västerbottenskuriren, and Dagens Industri from June 1988 until December 1998. Media coverage, and hence salience of the reform, was concentrated in the last quarter of 1989.

Figure A2: Predicted and Counterfactual Marriage Frequencies



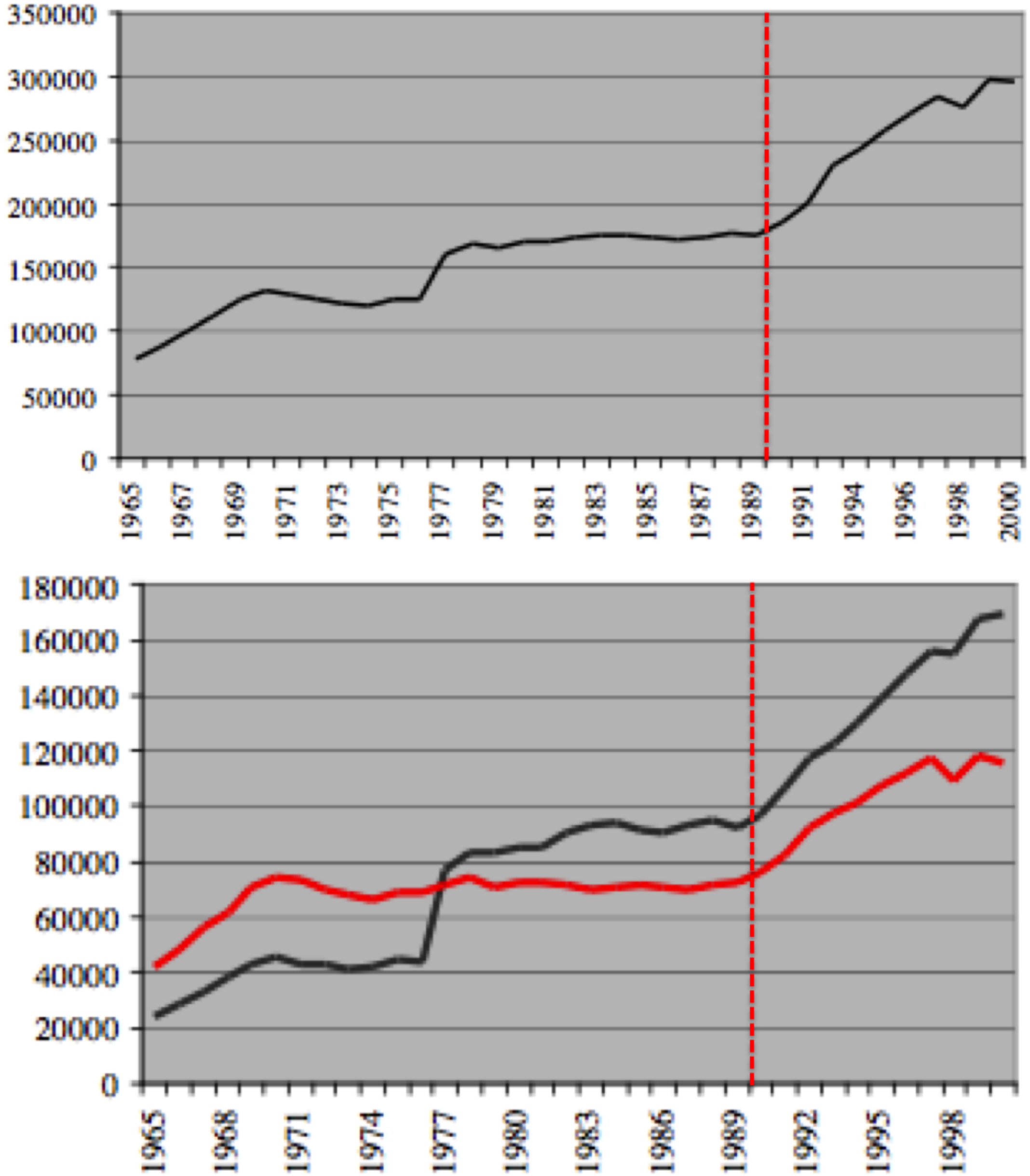
The black connected line displays the empirical distribution of the marriage frequency at a quarterly frequency. The blue solid line depicts the predicted frequencies within the estimation sample, which includes all quarters before the new marriage contract was launched, that is, all quarters before 1990 Q1. The red dotted line depicts the estimated counterfactual marriage frequency. For 1989 Q4, this counterfactual is obtained by setting the "bunching dummy," which takes the value of one in 1989 Q4, equal to zero. For quarters after 1989 Q4, the counterfactual is the out-of-sample prediction.

Figure A3: Empirical Distribution of Log (New Marriages)



The sample includes all couples that had a joint child between January 1, 1971 and January 1, 1989. The black connected dots depict the natural logarithm of the number of marriages at a quarterly frequency, from 1980 to 2003. The green solid vertical line indicates the last quarter during which marriage entailed survivors insurance. The grey dashed vertical line indicates the quarter of reform announcement.

Figure A4: The Number of University Enrollees (Total and by Gender)



The upper panel depicts the number of university enrollees by year. The lower panel depicts the gender breakdown, with the black solid line representing women and the red solid line representing men. Source: Högskoleverket. (The National Board of Higher Education.)

Table A1: Impact on Marriage Distribution: Heterogenous Effects w.r.t. Husband Age

	Husband Age At Marriage		
	≤30	31 - 40	≥40
Induced Marriages	8886***	25251***	10673***
	(712)	(3219)	(1298)
Estimated Counterfactual	735***	920***	268***
	(62)	(88)	(23)
$\Delta p/p$	12.09***	27.45***	39.2***
	(1.38)	(3.12)	(4.16)
Number of observations	110181	123993	13005

Note: Standard errors obtained from cluster bootstraping procedure (further described in the text) with 10 000 randomly drawn panels in parentheses. Each column represents a regression with second order polynomials in quarter and quarters pre and post birth. Each regression also includes cohort fixed effects, (four) quarter fixed effects, and cohort-specific increases (decreases) in entry into marriage at (after) the eligibility threshold.

* $p < 0.05$, ** $p < 0.01$, *** $p < 0.001$.

Table A2: Summary statistics for samples with varying male *ex post* mortality

	Sample: Couples where the husband died within				
	19 years	15 years	10 years	5 years	3 years
Financial characteristics					
H income share 1988	0.08	0.08	0.09	0.09	0.08
H ln labor income 1988	2.22	2.19	2.18	2.15	2.13
H is employed in 1988	1.46	1.52	1.58	1.68	1.71
W is employed in 1988	1.66	1.67	1.73	1.69	1.67
H birth year	1948	1948	1948	1948	1948
W birth year	1953	1953	1953	1953	1953.62
Observations	10719	6185	3239	1327	715
Other observable characteristics					
H education (4 lvls)	1.95	1.96	1.94	1.95	1.98
W education (4 lvls)	2.02	2.03	2.01	2.02	2.07
H age at marriage	35.03	34.77	34.33	33.77	33.81
H marriage number	1.27	1.27	1.27	1.25	1.27
W marriage number	1.18	1.17	1.17	1.17	1.18
Couple's completed fertility	1.94	1.92	1.90	1.86	1.81
Observations	10719	6185	3239	1327	715

Note: The "financial characteristics" influence the annuity's value at payout.

Table A3: Impact on Marriage: Heterogenous Effects w.r.t. Male *Ex Post* Mortality

	Financial controls		All observable controls	
Estimates from sample	Coefficient, $\hat{\beta}$	Hazard ratio, $e^{\hat{\beta}}$	Coefficient, $\hat{\beta}$	Hazard ratio, $e^{\hat{\beta}}$
Husband dies within 2 years	3.08***	21.80***	3.24***	25.48***
	(0.16)	(3.50)	(0.18)	(4.51)
Number of couples	410	410	359	359
Husband alive after 2 years	2.83***	16.88***	2.84***	17.03***
	(0.03)	(0.52)	(0.03)	(0.55)
Number of couples	246636	246636	217788	217788
Husband dies within 3 years	2.93***	18.74***	3.17***	23.72***
	(0.14)	(2.55)	(0.15)	(3.58)
Number of couples	651	651	563	563
Husband alive after 3 years	2.83***	16.88***	2.84***	17.03***
	(0.03)	(0.52)	(0.03)	(0.55)
Number of couples	246395	246395	217584	217584
Husband dies within 4 years	2.93***	18.72***	3.03***	20.76***
	(0.12)	(2.20)	(0.12)	(2.59)
Number of couples	921	921	797	797
Husband alive after 4 years	2.83***	16.88***	2.84***	17.03***
	(0.03)	(0.52)	(0.03)	(0.55)
Number of couples	246125	246125	217350	217350
Husband dies within 5 years	3.00***	20.12***	3.06***	21.28***
	(0.10)	(1.93)	(0.11)	(2.24)
Number of couples	1214	1214	1051	1051
Husband alive after 5 years	2.83***	16.88***	2.83***	17.03***
	(0.03)	(0.52)	(0.03)	(0.55)
Number of couples	245832	245832	217096	217096
Husband dies within 12 years	2.94***	18.86***	2.97***	19.52***
	(0.06)	(1.10)	(0.06)	(1.23)
Number of couples	3901	3901	3385	3385
Husband alive after 12 years	2.82***	16.86***	2.83***	17.00***
	(0.03)	(0.52)	(0.03)	(0.55)
Number of couples	243145	243145	214762	214762

Note: This table replicates the lower panel of Table 5 for different male mortality samples. See the notes to Table 5.

* $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$.

Table A4: Hightened Divorce Risk in Rushed Marriages: Probit Results

	Dependent variable: Divorce within			
	3 years	5 years	10 years	15 years
Married in 1989Q4	0.0095***	0.0250***	0.0411***	0.0367***
	(0.0021)	(0.0042)	(0.0078)	(0.0086)
Demographic characteristics at marriage				
H age at marriage	-0.0005**	-0.0008**	-0.0019***	-0.0026***
	(0.0002)	(0.0003)	(0.0004)	(0.0005)
W age at marriage	-0.0024***	-0.0056***	-0.0099***	-0.0115***
	(0.0001)	(0.0002)	(0.0003)	(0.0005)
H education (4 lvls)	-0.0019**	-0.0055***	-0.0111***	-0.0116***
	(0.0007)	(0.0011)	(0.0027)	(0.0026)
W education (4 lvls)	-0.0082***	-0.0170***	-0.0318***	-0.0394***
	(0.0008)	(0.0013)	(0.0019)	(0.0021)
H cognitive capacity	-0.0058***	-0.0104***	-0.0173***	-0.0192***
	(0.0008)	(0.0009)	(0.0013)	(0.0014)
H second marriage	0.0119***	0.0262***	0.0715***	0.0889***
	(0.0021)	(0.0031)	(0.0056)	(0.0066)
W second marriage	0.0090***	0.0237***	0.0649***	0.0984***
	(0.0019)	(0.0035)	(0.0058)	(0.0060)
H is immigrant	0.0071*	0.0123*	0.0202**	0.0306**
	(0.0032)	(0.0057)	(0.0076)	(0.0097)
W is immigrant	0.0047*	0.0079	0.0044	0.0036
	(0.0021)	(0.0042)	(0.0048)	(0.0062)
Economic characteristics at marriage				
Total household labor income (H and W)	-0.0070***	-0.0175***	-0.0304***	-0.0355***
	(0.0007)	(0.0013)	(0.0022)	(0.0025)
H income share	0.0078	0.0348***	0.0673***	0.0723***
	(0.0040)	(0.0077)	(0.0103)	(0.0138)
W earns much more	0.0033	0.0110**	0.0257***	0.0277***
	(0.0022)	(0.0035)	(0.0052)	(0.0057)
H earns much more	0.0031	0.0064*	0.0084	0.0118*
	(0.0017)	(0.0032)	(0.0046)	(0.0050)
Total joint number of children				
Couple's completed fertility	-0.0337***	-0.0692***	-0.0991***	-0.0905***
	(0.0023)	(0.0047)	(0.0070)	(0.0075)
Number of obs	94681	94681	94681	94681

Note: The table presents average marginal effects. Each column represents a Probit regression with a different dependent variable. All regressions include wedding month fixed effects and wedding day of week fixed effects. Robust standard errors clustered on the (wedding month*wedding day of week) in parentheses.

* $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$.

Table A5: Hightened Divorce Risk in Rushed Marriages: Probit Results

	Dependent variable: Divorce within			
	5 years	5 years	15 years	15 years
Married in 1989Q4	-0.0144***	0.0250***	-0.0329***	0.0367***
	(0.0041)	(0.0042)	(0.0094)	(0.0086)
Demographic characteristics at marriage				
H age at marriage		-0.0008**		-0.0026***
		(0.0003)		(0.0005)
W age at marriage		-0.0056***		-0.0115***
		(0.0002)		(0.0005)
H education (4 lvls)		-0.0055***		-0.0116***
		(0.0011)		(0.0026)
W education (4 lvls)		-0.0170***		-0.0394***
		(0.0013)		(0.0021)
H cognitive capacity		-0.0104***		-0.0192***
		(0.0009)		(0.0014)
H second marriage		0.0262***		0.0889***
		(0.0031)		(0.0066)
W second marriage		0.0237***		0.0984***
		(0.0035)		(0.0060)
H is immigrant		0.0123*		0.0306**
		(0.0057)		(0.0097)
W is immigrant		0.0079		0.0036
		(0.0042)		(0.0062)
Economic characteristics at marriage				
Total household labor income (H and W)		-0.0175***		-0.0355***
		(0.0013)		(0.0025)
H income share		0.0348***		0.0723***
		(0.0077)		(0.0138)
W earns much more		0.0110**		0.0277***
		(0.0035)		(0.0057)
H earns much more		0.0064*		0.0118*
		(0.0032)		(0.0050)
Total joint number of children				
Couple's completed fertility		-0.0692***		-0.0905***
		(0.0047)		(0.0075)
Number of obs	94681	94681	94681	94681

Note: The table presents average marginal effects. Each column represents a regression with a different dependent variable. All regressions include wedding month fixed effects and wedding day of week fixed effects. Robust standard errors clustered on the (wedding month*wedding day of week) level in parentheses.

* $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$.

B Social Security in Sweden

This information is largely obtained from Palme and Svensson (1997). In Sweden, 74% of the income of individuals above the age of 65 constitutes Social Security income (this exact figure is from 1994). Only a small share is from private pension insurance; the remainder, instead, is mostly income from employer-provided pensions, centrally negotiated by unions.

All persons living in Sweden were entitled to a *basic pension* from the Social Security system, which was not tied to earnings history (but to the time of residence in Sweden). The size of the basic pension is linked to the "basic amount" (BA), which in turn is linked to the consumer price index. In 1988, the BA was SEK 25800, and the annual salary income among men in my sample used in Section 5.2 was SEK 110968.

In addition, *supplementary pension* (ATP) payments from Social Security were linked to earned income. For an individual who had obtained labor income in Sweden for 30 years prior to retirement, the supplementary social security benefit, b_h, was given by

$$b_h = 0.6 * (Income - BA) \text{ for } Income \in (BA, 7.5BA)\,. \tag{19}$$

Here, pension rights $Income$ is earnings and self-employment income recorded in the annual tax return, which I know for each year from1985 to 2009. This is a lower bound of the pension rights income, however, which also includes income from sickness and unemployment insurance, parental leave benefits, and the partial retirement pension, neither of which I observe. The lower limit ascertains that supplementary pension is positive; the upper limit, 7.5 BA, is the social security ceiling. In 1988, the BA was SEK 25800, and the social security ceiling thus SEK 193500. In 1988, this corresponded to the 91th percentile in the distribution of salary income in the sample that I use in Section 5.2.

For individuals who worked in Sweden for $N < 30$ years prior to retirement, the supplementary social security benefit in (19) was multiplied by the function $min\left\{\frac{N}{30}, 1\right\}$. That is, the supplementary pension was reduced by $\frac{1}{30}$ for each year of work experience below 30. Finally, three years of positive pension-rights income between the ages of 16 to 65 were required to be eligible for supplementary pension.

The Social Security system is financed by employer contributions levied on wages. The level of all social contributions was 31.36 percentage points on gross earnings in 1994. The level of the contribution for the national basic pension was 5.86, for the supplementary pension (ATP) 13.00, and for the part-time pension 0.20 percentage points. General tax revenues partially finance the national basic pension. The payments from these systems (basic and supplementary) amounted to 42.4 percent and 55.3 percent of total pension payments in 1994.

Other features of the marriage contract in Sweden

The main legal distinction between cohabitation and marriage in Sweden during the time period that I analyze concerns inheritance. A surviving spouse has a right to inheritance, but a cohabiting partner does not. Taxes and benefits were largely decoupled from marriage since 1971, when joint taxation of labor income was eliminated. Naturally, some changes are implemented with respect to inheritance rules and the precise benefit structure tied to marriage during the 40 years that I analyze, 1971 to 2009. In the empirical analyses, however, I always compare groups of couples whose marriage contracts differ only with respect to whether marriage comes with survivors insurance, but where all other features of the marriage contracts are similar.

C Mathematical Proofs of Results in Section 3

Lemma 1

Proof. Under transferable utility, a matching is stable if and only if it maximizes total surplus. A solution to this maximization problem thus exists by the Bolzano Weierstrass theorem, because (i) total surplus function M_{total} is continuous in $\{\tau_w, \tau_m\}$, and (ii) the function is maximized over the set of matching allocations $\{\tau_w, \tau_m\}$, which is compact. □

Lemma 2

Proof. In Stage 1, the expected value of match is $M(\tau_w, \tau_m) = \delta A(\tau_w, \tau_m) + \delta^2 B(\tau_w, \tau_m)$ where

$$A(\tau_w, \tau_m) = \alpha(\tau_w, \tau_m)\left[V(\tau_w, \tau_m) + E[\theta_2 | \theta_2 \geq \theta_{NSB}(\tau_w, \tau_m)]\right]$$

and

$$B(\tau_w, \tau_m) = \beta(\tau_w, \tau_m)\left[(1-p)\left[V(\tau_w, \tau_m) + E[\theta_3 | \theta_3 \geq \theta_{SB}(\tau_w, \tau_m)]\right] + pU_A(\tau_w, \tau_m)\right]$$

For $U_A(\tau_w, \tau_m) = 0$, we get $\theta_{SB}(\tau_w, \tau_m) = -V(\tau_w, \tau_m) = \theta_{NSB}(\tau_w, \tau_m)$, so that $\alpha(\tau_w, \tau_m) = \beta(\tau_w, \tau_m)$ and $A(\tau_w, \tau_m)$ becomes identical to $B(\tau_w, \tau_m)/(1-p)$ except that the shocks $\tilde{\theta}_2$ and $\tilde{\theta}_3$ are drawn from different distributions. Showing that $A(\tau_w, \tau_m)$ is supermodular, or $\partial A(\tau_w, \tau_m)/\partial\tau_w\partial\tau_m \geq 0$, for an arbitrary distribution F would then imply that $B(\tau_w, \tau_m)/(1-p)$ is also supermodular. This in turn would imply that $M = \delta A(\tau_w, \tau_m) + \delta^2 B(\tau_w, \tau_m)$ is supermodular, since positively linear combinations of supermodular functions are supermodular. Now write

$$A(\tau_w, \tau_m) = \int_{-V(\tau_w, \tau_m)}^{\infty} \left[V(\tau_w, \tau_m) + \theta_2\right] f(\theta_2) d\theta_2.$$

Then,

$$\begin{aligned} \frac{\partial A(\tau_w, \tau_m)}{\partial \tau_w} &= -\left[V(\tau_w, \tau_m) - V(\tau_w, \tau_m)\right] f(-V(\tau_w, \tau_m)) \left(\frac{-\partial V(\tau_w, \tau_m)}{\partial \tau_w}\right) \\ &\quad + \int_{-V(\tau_w, \tau_m)}^{\infty} \left[\frac{\partial V(\tau_w, \tau_m)}{\partial \tau_w}\right] f(\theta_2) d\theta_2 \end{aligned}$$

Simplifying yields

$$\frac{\partial A(\tau_w, \tau_m)}{\partial \tau_w} = \int_{-V(\tau_w, \tau_m)}^{\infty} \frac{\partial V(\tau_w, \tau_m)}{\partial \tau_w} f(\theta_2) d\theta_2.$$

The second derivative is thus given by

$$\frac{\partial^2 A(\tau_w,\tau_m)}{\partial\tau_w\partial\tau_m} = \frac{\partial V(\tau_w,\tau_m)}{\partial\tau_w} f\left[-V(\tau_w,\tau_m)\right]\frac{\partial V(\tau_w,\tau_m)}{\partial\tau_m} + \int_{-V(\tau_w,\tau_m)}^{\infty}\frac{\partial V(\tau_w,\tau_m)}{\partial\tau_w\partial\tau_m} f(\theta_2)d\theta_2 > 0.$$

By contrast, for $U_A(\tau_w,\tau_m) \neq 0$, $B(\tau_w,\tau_m)/(1-p)$ is not isomorphic to $A(\tau_w,\tau_m)$ and need not be supermodular. To show this, I use the fact that, for functions on R^2, supermodularity is equivalent to *increasing differences*. Function $B(\tau_w,\tau_m)$ has increasing differences if

$$\frac{\partial B(\tau_w',\tau_m)}{\partial\tau_m} \geq \frac{\partial B(\tau_w,\tau_m)}{\partial\tau_m} \quad \forall\tau_w' > \tau_w, \forall\tau_m. \tag{20}$$

Write

$$B(\tau_w,\tau_m) = \int_{\theta_{SB}(\tau_w,\tau_m)}^{\infty} g(\theta_3)\left[(1-p)\left[V(\tau_w,\tau_m)+\theta_3\right] + pU_A(\tau_w,\tau_m)\right]d\theta_3.$$

The derivative with respect to τ_m is

$$\frac{\partial B(\tau_w,\tau_m)}{\partial\tau_m} = \int_{\theta_{SB}(\tau_w,\tau_m)}^{\infty} g(\theta_3)\left[(1-p)\frac{\partial V(\tau_w,\tau_m)}{\partial\tau_m} + p\frac{\partial U_A(\tau_w,\tau_m)}{\partial\tau_m}\right]d\theta_3.$$

I now prove by example that this derivative need not satisfy (20). Consider a large p or a V-function where $\partial V/\partial\tau_m$ is very small everywhere, such that the first term in the bracket, $(1-p)\partial V/\partial\tau_m$, is negligible. In this case, the derivative is largely determined by the second term:

$$\frac{\partial B(\tau_w,\tau_m)}{\partial\tau_m} \approx \int_{\theta_{SB}(\tau_w,\tau_m)}^{\infty} g(\theta_3)p\frac{\partial U_A(\tau_w,\tau_m)}{\partial\tau_m}d\theta_3.$$

Fix τ_m. Because $\frac{\partial U_A(\tau_w',\tau_m)}{\partial\tau_m} = 0$ for $\tau_w' > \tau_m$ and $\frac{\partial U_A(\tau_w,\tau_m)}{\partial\tau_m} > 0$ for $\tau_w < \tau_m$, (20) is violated:

$$\frac{\partial B(\tau_w',\tau_m)}{\partial\tau_m} \approx 0 < \frac{\partial B(\tau_w,\tau_m)}{\partial\tau_m} \quad \text{for } \tau_w' > \tau_w, \tau_m \in (\tau_w,\tau_w')$$

Thus, $B(\tau_w,\tau_m)$ does not have increasing differences everywhere and is hence not supermodular. □

Results in subsection 3.3.2

I formulate the results regarding couples that were matched but unmarried at the reform's announcement in the following Proposition:

Proposition 1 (Reform announced in Stage 2.)**.** *Consider all couples not yet married at the reform announcement that become eligible for survivors benefits if and only if they marry during a limited time period, that is, within Stage 2. First, selection into marriage in Stage 2 is stronger when the man is more likely to die in Stage 3 (higher p). Second, the reform induces some marriages in Stage 2 that would otherwise occur in Stage 3, but also some marriages that would never occur without the reform.*

Third, couples that the reform induces to marry are more likely to divorce.

Proof. First, I show that selection into marriage increases with p. As can be seen in (6), for this, I need only show that the option value Ω_A increases in p. Write

$$\Omega_A = \delta \left[B_{\theta_2}(\tau_w, \tau_m) - (1-p)A_{\theta_2}(\tau_w, \tau_m)\right] \tag{21}$$

where

$$\begin{aligned} A_{\theta_2}(\tau_w, \tau_m) &= \int_{\theta_{NSB}(\tau_w,\tau_m)}^{\infty} \left[V(\tau_w, \tau_m) + \theta_3\right] g_{\theta_2}(\theta_3) d\theta_3 \\ B_{\theta_2}(\tau_w, \tau_m) &= \int_{\theta_{SB}(\tau_w,\tau_m)}^{\infty} \left[(1-p)\left[V(\tau_w, \tau_m) + \theta_3\right] + pU_A(\tau_w, \tau_m)\right] g_{\theta_2}(\theta_3) d\theta_3. \end{aligned}$$

Recall that $\theta_{NSB}(\tau_w, \tau_m) > \theta_{SB}(\tau_w, \tau_m)$. Substituting $A_{\theta_2}(\tau_w, \tau_m)$ and $B_{\theta_2}(\tau_w, \tau_m)$ into (21) and rearranging yields

$$\Omega_A = \delta p \int_{\theta_{SB}(\tau_w,\tau_m)}^{\infty} U_A(\tau_w, \tau_m) g_{\theta_2}(\theta_3) d\theta_3 + \delta(1-p) \int_{\theta_{SB}(\tau_w,\tau_m)}^{\theta_{NSB}(\tau_w,\tau_m)} \left[V(\tau_w, \tau_m) + \theta_3\right] g_{\theta_2}(\theta_3) d\theta_3,$$

where the first integral is positive and the second integral is negative (by the definitions of $\theta_{NSB}(\tau_w, \tau_m)$ in (5)).

The derivative with respect to p is

$$\Omega_A = \delta \int_{\theta_{SB}(\tau_w,\tau_m)}^{\infty} U_A(\tau_w, \tau_m) g_{\theta_2}(\theta_3) d\theta_3 - \delta \int_{\theta_{SB}(\tau_w,\tau_m)}^{\theta_{NSB}(\tau_w,\tau_m)} \left[V(\tau_w, \tau_m) + \theta_3\right] g_{\theta_2}(\theta_3) d\theta_3 > 0.$$

Second, I show that the increase in marriages during the grandfathering period comprises retimed marriages and, *given uncertainty about* θ_3, extra marriages. Among those couples that the reform *causes* to marry in Stage 2, some will get shocks $\theta_3 \geq \theta_{SB}$ in Stage 3 and hence would have married in Stage 3 in the absence of the reform. However, some will get shocks $\theta_3 < \theta_{SB}$, and hence would not have married in Stage 3. Now suppose that θ_3 is known already at the beginning of Stage 2. In that case, couples with $\theta_3 < \theta_{SB}$ do not want to marry in Stage 2 only to keep the option on survivors benefits alive. This is because they will certainly not exercise the option, since they already know that they do not want to be married in Stage 3, even when eligible for survivors benefits. The aforementioned "extra marriages" during the grandfathering period would thus not happen.

Third, I show that the divorce rate increases. From inspection of (6), it is clear that the couples that the reform induces to marry in Stage 2 have lower V or lower θ_2 than those that marry even absent the reform. For a given θ_2, a couple is less likely to satisfy (3) for lower V. And for a given V, a couple is less likely to satisfy (3) for lower θ_2 because $G_{\theta_2}(\theta_3)$ first-order stochastically dominates $G_{\theta_2'}(\theta_3)$ for all $\theta_2 > \theta_2'$. □

Results in subsection 3.3.3

I formulate the results regarding couples that were unmarried at the reform's announcement in the following Proposition:

Proposition 2 (Reform announced in Stage 3.)**.** *Consider all couples already married at the reform announcement. In couples that remain married, the wife's share of household utility (weakly) increases.*

Proof. In married couples, let $u_3^i(\tau_m, \tau_w)$ denote spouse i's utility in Stage 3 if the husband is alive under the contract entered at marriage. These utilities satisfy

$$E\left[u_3^w(\tau_m, \tau_w)\right] = (1-p)\gamma\left[V\left(\tau_m, \tau_w\right) + \theta_3\right] + pU_A\left(\tau_m, \tau_w\right) \geq 0$$

and the husband wants to remain in marriage if and only if

$$E\left[u_3^m(\tau_m, \tau_w)\right] = (1-p)\left[1-\gamma\right]\left[V\left(\tau_m, \tau_w\right) + \theta_3\right] \geq 0.$$

For some shocks θ_3, the reform – the abolition of survivors benefits – causes the wife's expected utility from marriage to be smaller than her outside utility:

$$(1-p)\gamma\left[V\left(\tau_m, \tau_w\right) + \theta_3\right] + pU_A\left(\tau_m, \tau_w\right) \geq 0 > (1-p)\gamma\left[V\left(\tau_m, \tau_w\right) + \theta_3\right].$$

By contrast, the husband's expected utility is unaffected by the reform under any contract. Thus, if anyone, the wife may seek a separation after the reform. In such a state, there is scope for efficient renegotiation if

$$(1-p)\left[V\left(\tau_m, \tau_w\right) + \theta_3\right] \geq 0 > (1-p)\gamma\left[V\left(\tau_m, \tau_w\right) + \theta_3\right].$$

To keep the wife in the marriage, the husband would then have to agree to a new sharing rule $\hat{\gamma} > \gamma$ such that $(1-p)\hat{\gamma}\left[V\left(\tau_m, \tau_w\right) + \theta_3\right] \geq (1-p)\underline{u}_w$. □

D Infinite Horizon Framework of Selection into Marriage

The cohabitation versus marriage margin: Derivation of the marriage hazard To highlight the impact of the change in marriage contract on the decision whether to marry or wait (cohabitate), I now abstract from the matching decision (and therefore let $V\left(\tau_w, \tau_m\right) = 0$) and consider an already-formed couple that, in each time period, decides whether to enter marriage. For simplicity, I here consider marriage an irreversible decision. See Stock and Wise (1990) and Manoli and Weber (2011) for similar option value models.

Consider an unmarried couple that decides whether to get married or wait based on the discounted flow of lifetime income under both options. With survivors insurance, the decision in each period t, is based on the following set of state variables (in addition to time t):

$$\Lambda_t = \begin{cases} \bar{a}_t & \text{expected per-period (survivors) benefit if marrying at date } t \\ \hat{\underline{u}} & \text{per-period benefit of waiting (normalized to zero)} \\ \theta_t & \text{stochastic utility of being married in time } t \end{cases}$$

To allow for this per period formulation, I convert the value of the annuity to an annualized deterministic survivors benefit stream, where marriage in time t yields a stream of $\bar{a}_t$ in each future time period. See the next paragraph for this derivation, and and for details on the solution discussed below.

As before, a couple's optimal strategy involves marrying if and only if $\theta_t \geq \bar{\theta}_t(\Lambda_t)$. Solving for $\bar{\theta}_t(\Lambda_t)$ recursively and using a Taylor expansion of the per-period utility $u(\cdot)$ around $\hat{\underline{u}}$ yields that optimal entry into marriage is characterized by the following hazard rate at time t:

$$h_t = Pr\left\{\theta_t \geq \bar{\theta}_t(\Lambda_t)\right\} = Pr\left\{\theta_t \geq \nu(\underline{u} - \bar{a}_t) + \beta_t O_t\right\},$$

where $\nu = u'(0)$, and O_t denotes the option value of marrying at a later time,

$$O_t = E_t\left(V^W(\Lambda_{t+1})\right) - V^M(t+1, \bar{a}_t), \tag{22}$$

and $V^W$and V^M are the value functions for waiting and marrying, respectively. The distribution of marriages over time is thus determined based on the number of individuals considering marriage.

Now consider an alternative scenario where this marriage contract is only available for a limited time period, that is, the couple only receives the lifetime stream of $\bar{a}_t$ if marrying before s^*. This makes the remaining time until s^* an additional state variable; I label this expanded set of relevant state variables by $\tilde{\Lambda}_t$. Denote by A the discounted value of the annuity in s^*. After the reform's announcement, the hazard rate is given by

$$h_t = \begin{cases} Pr\left\{\theta_t \geq \nu(-\bar{a}_t) + \beta_t \tilde{O}_t\right\} & \text{if } t < s^* \\ Pr\left\{\theta_t \geq \nu(-\bar{a}_t + B) + \beta_t \tilde{O}_t\right\} & \text{if } t \geq s^* \end{cases}, \tag{23}$$

where $\tilde{O}_t$ is the option value taking into account that no survivors benefits accrue if marriage entry occurs after $s = s^*$. Even though the option value changes permanently at the reform announcement, the cost of waiting to marry only changes in one time period, s^*. In this period, there will be a discrete increase in the hazard rate, the magnitude of which is increasing in A_t . I estimate this hazard function in Section 5.2.

Annualization The discounted expected value of future survivors benefit if marrying at date t is given by

$$A_t = E_t(U_{death}) = d_t \frac{U(a_t)}{\delta} + (1 - d_t)\,\delta\left[d_{t+1}\frac{U(a_{t+1})}{\delta} + (1 - d_{t+1})\,\delta\left[d_{t+2}\frac{U(a_{t+2})}{\delta} + (1 - d_{t+2})\,\delta\,[\ldots]\right]\right],$$

where the husband's probability of death increases with time, $d_t < d_{t+1}$. Annualization requires thisNPV_t to converge. Rewriting yields

$$NPV_t = d_t \frac{a_t}{\delta} + (1-d_t)\,\delta d_{t+1}\frac{a_{t+1}}{\delta} + (1-d_t)\,(1-d_{t+1})\,\delta^2 d_{t+2}\frac{a_{t+2}}{\delta} + (1-d_t)\,(1-d_{t+1})\,(1-d_{t+2})\,\delta^3 d_{t+3}\frac{a_{t+3}}{\delta}...,$$

which can be written

$$NPV_t = d_t \frac{a_t}{\delta} + \sum_{s=1}^{\infty} t_n, \tag{24}$$

where

$$\sum_{s=1}^{\infty} t_n = \sum_{s=1}^{\infty}\left(\prod_{x=0}^{s-1}(1-d_{t+x})\right)\delta^s d_{t+s}\frac{a_{t+s}}{\delta}.$$

I now show that the sum $\sum_{s=1}^{\infty} a_n$ converges. For simplicity, let $a_{t+1} = a_t$. The ratio of two terms $\frac{t_{n+1}}{t_n}$ is given by

$$\frac{t_{n+1}}{t_n} = (1-d_{t+n})\,\tfrac{d_{t+n+1}}{d_{t+n}}\delta.$$

First, suppose that d_t were constant over time. Then,

$$lim_{n\to\infty}\left|\frac{t_{n+1}}{t_n}\right| = (1-d)\,\delta.$$

Because $(1-d)\,\delta < 1$, the series would converge, by D'Alembert's Criterion (the Ratio test). In fact, this would arise even if $d_t = 0$. Define by k_n the convergent series that arises when $d_t = 0$ for all t.

In the series t_n, however d_t is increasing over time. Because the series k_n converges, the series t_n converges if

$$|t_n| \leq |k_n|$$

for sufficiently large n. I have that

$$\left(\prod_{x=0}^{s-1}(1-d_{t+x})\right)\delta^s d_{t+s}\frac{a_{t+s}}{\delta} \leq \delta^s \frac{a_{t+s}}{\delta}$$

for all n. Hence, the series t_n converges.

NPV_t is also the discounted expected value of an—equivalent—deterministic annuity that starts paying out immediately upon marriage. Refer to this deterministic annuity U_t^{Ann} as "annualized expected survivors benefits." It is determined by

$$E_t(U_{death}) = \frac{U(\bar{a}_t)}{\delta},$$

where $\bar{a}_t$ is a deterministic stream of in every time period. It is nevertheless indexed by t to capture the fact that marriage in the next time period entails a higher deterministic payment in each time period from then and onwards.

Solving the couple's problem At each time period t the value of marriage is given by

$$V^M(t, \bar{a}_t) = u(\bar{a}_t) + \beta_t V^M(t+1, \bar{a}_t),$$

and the value of waiting by

$$V^W(\Lambda_t) = -\theta_t + \delta\,(1-d_t)\,E_t[V(\Lambda_{t+1})].$$

The series of discount factors $\delta\,(1-d_t)$ of future utility capture the couple's one-period discount factor δ; it also captures the probability of the husband's survival, and thus the probability that the couple will face the same decision problem in the next period in case they choose to wait.[22] The expected value of the optimally chosen decision in the next period is given by

$$E_t[V(\Lambda_{t+1})] = max\left\{V^M(t+1, \bar{a}_{t+1}), V^W(\Lambda_{t+1})\right\}.$$

As in the setting in Section 3, the couple's optimal strategy can be described by a reservation disutility value $\bar{\theta}_t\,(\Lambda_t)$: The couple marries if and only if the surplus from marriage (disutility from waiting) satisfies $\theta_t \geq \bar{\theta}_t\,(\Lambda_t)$, where $\bar{\theta}_t\,(\Lambda_t)$ is implicitly given by the condition

$$V^M(t, \bar{a}_t) = V^W(t, \bar{a}_t, s, \bar{\theta}_t).$$

Using a first order Taylor expansion of u around 0, I express this as

$$\bar{\theta}_t = \nu(-\bar{a}_t) + \beta_t O_t, \tag{25}$$

where $\nu = u'(0)$ and O_t denotes the option value of marrying at a later time,

$$OV_t = E_t\left(V^W(\Lambda_{t+1})\right) - V^M(t+1, \bar{a}_t), \tag{26}$$

Here, O_t includes both future survivors benefits and expectations of future (potentially negative) realizations of θ_t. Intuitively, an individual may get a negative theta of considerable magnitude in the future, in which case she would be very unhappy in marriage, or equivalently she would get a high utility from being unmarried tomorrow. Equation 25 defines $\bar{\theta}_t$ dynamically. Solving the equation recursively, starting from a fixed maximum T, yields that the hazard rate of marriage at time t can be written:

$$h_t = Pr\left\{\theta_t \geq \bar{\theta}_t\,(\Lambda_t)\right\} = Pr\left\{\theta_t \geq \nu(-\bar{a}_t) + \beta_t O_t\right\}.$$

The distribution of marriages over time thus depends on the population of couples at risk, that is, the size of the population that considers entry into marriage.

[22] Note that the value of waiting also "includes" the term $\delta d_t \frac{u(0)}{\delta} = 0$.

Impact of a change in the law I now turn to the alternative scenario where a reform is announced, stating that this marriage contract only will remain available for a limited time period. Specifically, the couple only receives the lifetime stream of $\bar{a}_t$ if and only if marrying before s^*. This makes the remaining time until s^* an additional state variable; the expanded set of state variables is given by $\tilde{\Lambda}_t$. In this scenario, a similar procedure and Taylor expansion as above yields that equation (25) instead is given by

$$\tilde{\theta}_t = \begin{cases} \theta_t \geq \nu(-\bar{a}_t) + \beta_t \tilde{O}_t \text{ if } s < s^* \\ \theta_t \geq \nu(-\bar{a}_t + A) + \beta_t \tilde{O}_t \text{if } s \geq s^* \end{cases}.$$

After the reform's announcement, the hazard rate is thus given by

$$h_t = \begin{cases} Pr\left\{\theta_t \geq \nu(-\bar{a}_t) + \beta_t \tilde{O}_t\right\} & \text{if } t < s^* \\ Pr\left\{\theta_t \geq \nu(-\bar{a}_t + A) + \beta_t \tilde{O}_t\right\} & \text{if } t \geq s^* \end{cases}, \tag{27}$$

where $\tilde{O}_t$ is the option value taking into account that no survivors benefits accrue if marriage entry occurs after $s = s^*$. Even though the option value changes permanently at the reform announcement, the cost of waiting to marry only changes in one time period, s^*. In this period, there the hazard rate increases. This increase increases with A_t .