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WOMEN'S INCOME AND MARRIAGE MARKETS IN THE UNITED STATES: EVIDENCE FROM THE CIVIL WAR PENSION

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ABSTRACT

Under the Civil War pension act of 1862, the widow of a Union Army soldier was entitled to a pension if her husband died as a direct result of his military service; however, she lost her right to the pension if she remarried. I analyze the effect this had on the rate of remarriage among these widows. This study fits into a modern literature on the behavioral effects of marriage penalties. In addition, it offers a unique perspective on 19th century marriage markets, which are little understood. Using a new database compiled from widows' pension files, I estimate the effect of the pension on the hazard rate of remarriage using variation in pension processing times. Taking steps to account for the potential endogeneity of processing times to marital outcomes, I find that receiving a pension lowered the hazard rate of remarriage by 25 percent, which implies an increase in the median time to remarriage of 3.5 years. Among older women and women with children, this effect is substantially greater. This indicates that women were willing to substitute away from marriage if the alternatives were favorable enough, suggesting that changes in the desirability of marriage to women may account for some of the aggregate patterns of first marriage documented for this period.

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1 Introduction

Certain social assistance programs tie benefits to marital status, typically by reducing benefits if a recipient marries. For example, the Aid to Families with Dependent Children (AFDC) program is targeted at single mothers. Similarly, survivors' pension benefits in the U.S. and Canada were, until recently, reduced if the beneficiary remarried.[1] The behavioral consequences of these marriage penalties have frequently been studied by economists, with a particular focus on marriage market outcomes. Such analyses are of interest in part because they provide insight into the unintended welfare consequences of specific policy interventions; however, they are more broadly interesting because of the light they shed on the economics of family formation. In this paper, I analyze the effects of the marriage penalty built into the Union Army pension. Established in 1862, this was America's first large-scale social assistance program. In addition to supporting sick and wounded veterans, this pension provided compensation for the widows of Union Army soldiers who died as a direct consequence of their military service. However, these widows lost their right to this pension if they remarried. Using a new database compiled from Civil War pension records, I measure the extent to which the Union Army pension caused widows to delay remarriage.

Studying the effect of marriage penalties on women's behavior is useful because it offers insight into women's marriage-market responses to independent income sources. Becker (1973, 1991) argues that marriage generates utility by allowing couples to exploit increasing returns through division of labor; a marriage will occur if marital output exceeds the sum of the output that both partners produce while single. If women have access to an independent income source, this will lower the net gains they experience from marriage, which should discourage them from marrying. Search models of the marriage market predict that, if a woman's income functions as an alternative to marriage, it should raise the value of being single relative to the value of being married, thus raising the woman's reservation match quality. Under random matching, this will lower the probability that any given proposal of marriage will be deemed suitable, which will tend to result in delayed marriage.[2] In general, these theoretical predictions are difficult to test because of the interrelatedness of decisions regarding career and family. For instance, a woman's labor income depends on her human capital investments, which may be endogenous to preferences for marriage. Because a social assistance program with a marriage penalty generates an income stream that varies only with marital status,

[1] See Rosensweig (1999); Baker, Hanna and Kantarevic (2004); Brien, Dickert-Conlin and Weaver (2004).

[2] See Rogerson, Shimer and Wright (2005) for a survey of basic search models. See Weiss (1997) for a review of search models applied to marriage markets. Gould and Paserman (2003) and Loughran (2002) use a search framework to investigate the effect of wage inequality on marriage rates.

analyzing its effects circumvents many of these endogeneity issues. Studying the marriage penalty built into the Civil War pension is in some ways preferable to studying modern social assistance programs because it was neither age-based nor means tested, so its effects may be considered more general. Moreover, because cohabitation outside marriage was relatively uncommon in the 19th century,[3] the effect of the Civil War pension can be presumed to reflect factors like increased selectivity in the search for mates rather than substitution of cohabitation for marriage.[4]

The effects of the Civil War pension on widows' choices about marriage also provides insight into 19th century marriage markets, which is something we know relatively little about. Certain aggregate trends in marital outcomes are well documented. For instance, the female age at first marriage rose steadily over the course of this century, increasing from roughly 20 during the colonial period to a peak of 23.6 in 1890 (Haines 1996). Explanations for these patterns include a decline in land availability, which increased the the cost of establishing new households, and falling male-to-female ratios, most notably in the aftermath of the Civil War.[5] Much less attention has been paid to the role of women's economic opportunities in altering the desirability of marriage to women.[6] If Civil War pension income has a causal effect on marriage behavior, this suggests that this latter avenue may be important for understanding developments during the 19th century.

I compile a novel database containing information on widows' pension applications and subsequent marriages from the Civil War pension files at the National Archives in Washington, DC. To assess the extent to which pensions caused widows to delay remarriage, I make use of variation in the timing of pension decisions, or pension processing times. Because pension amounts were standardized, processing times provide the most plausibly exogenous variation in pension income within my sample. I estimate a proportional hazards model of remarriage in which the rate of remarriage is allowed to shift at the moment a pension is granted. As such, I estimate a treatment effect of transitioning from having a pending claim to having an accepted claim.

One concern with this approach is the possible endogeneity of pension processing times to marital outcomes. This is largely due to sample selection, which is generated by the decision to apply for a pension. Women whose pensions take a long time to process tend to be those with ambiguous

[3]In any case, openly cohabiting couples were considered "married" for the purposes of pension eligibility.

[4]See Lundberg and Pollak (2013) for a discussion of the increased incidence and importance of cohabitation as a type of marital relation during the later 20th century.

[5]See Easterlin (1971; 1976), Haines (1996) and Hacker (2008)

[6]One recent study (Hacker 2008) includes this in a set of possible correlates of first marriage and documents a correlation between the age at first marriage and labor force participation among unmarried women in the 1860 census. However, this paper does not address the potential endogeneity of female labor force participation to norms of marriage or marriage market conditions.

claims, and those who choose to incur the cost of applying for a pension even though their claims are ambiguous may be systematically different from those who apply with straightforward claims. To address this, I exploit the fact that my treatment variable is a duration variable, which provides more information than is available in a standard cross-sectional setting. Variation in observables and the relationship between the hazard rates of pension receipt and remarriage provide sufficient information to correct for correlated unobserved heterogeneity in these two risks (Abbring and Van den Berg 2003; 2005). As an additional test, I estimate a linear version of this model using two stage least squares. My instrument for pension processing time is a measure of surname spelling homogeneity, calculated as the dispersion of unique spellings within phonetic surname groups in the censuses of 1860, 1870 and 1880. This generates variation in the difficulty of proving a soldier's identity, which altered the amount of time it took for a claim to be granted.

I find that receiving a pension caused the rate of remarriage to drop by 25 percent, implying an increase in the median time to remarriage of 3.5 years. Moreover, I find that this effect is heterogeneous: the effect of the pension increases in magnitude with the widow's age and number of children. These effects are especially striking because of the small size of the pension. At eight dollars per month, the pension was less than half the monthly income of a typical farm laborer in 1870, so it was hardly enough to comfortably support a family. While the effect of the pension is small for young, childless women, these findings demonstrate that women responded to outside income sources in the marriage market during this period. This lends credence to the idea that the incremental changes in female labor market opportunities seen in the 19th century may have contributed to the aggregate changes in marriage patterns that occurred during this period; at the very least, my results suggest that this channel should be further investigated.

2 Related Literature

A number of recent social assistance programs contain explicit marriage penalties, and there is an economic literature that evaluates the effect these have on behavior. Rosenzweig (1999) studies the effect of the AFDC program on marriage and out-of-wedlock childbearing for young women, and he finds that AFDC benefits tend to encourage fertility outside marriage. Baker, Hanna and Kantarevic (2004) find a significant negative effect of marriage penalties on remarriage, which they identify through the removal of marriage penalties from the public pension system in Canada during the 1980s. Brien, Dickert-Conlin and Weaver (2004) find that American widows and widowers delayed remarriage until after the age of 60 in response to the marriage penalty built into Social Security before 1979.

There is a broader literature on the effect of women's labor income on their marriage-market behavior. The main challenge associated with measuring this is that a woman's income is not exogenous to her marital outcomes. As such, most of the empirical literature on the effect of female income on marriage rates is descriptive, largely demonstrating a negative correlation between income or career opportunities for women and marriage rates.[7] A paper that deals explicitly with causality is Blau, Kahn and Waldfogel (2000), who look at the effect of city-wide marriage and labor market conditions on marriage rates. They find that better female labor markets tend to decrease marriage rates, while better male labor markets tend to increase them. Still, it is not clear from this analysis that female labor market opportunities cause women's choices about marriage to change: areas in which these opportunities are greater may have different social norms surrounding marriage.[8]

The evaluation of a 19th century program with a marriage penalty contributes to the literature on marriage patterns specific to this period. While the increase in the female age at first marriage that occurred during the 19th century is well documented, its causes are not well understood.[9] Most explanations focus on opportunities rather than preferences for marriage. In contrast to Western Europe, where "couples often delayed marriage until the prospective bridegroom inherited the family farm" (Fitch and Ruggles 2000, p. 62), land in the United States was cheap and abundant and did not pose an impediment to early marriage. However, as land became increasingly settled, marriage patterns started to more closely resemble those in Europe. As farmland grew scarcer and more expensive, "men were forced either to postpone marriage, working as farmhands or manual laborers until they had saved up enough money to set up their own farms, or to migrate to the western frontier" (Hacker 2008, p. 312). Easterlin (1976) also links the closing of the frontier to fertility control within marriage. As international and internal migration patterns changed over the course

[7]Keeley (1977) finds that women with high wages tend to marry later, although men with high wages tend to marry earlier. Ruggles (1997) argues that increasing female labor market opportunities contributed to the rise in divorce rates during the twentieth century. Weiss and Willis (1997) find that women with high earnings are more likely to divorce, while the opposite is true of men with high earnings. Price-Bonham and Balswick (1980) argue that widows are less likely to remarry than divorced women, as are older and more educated women with fewer children. Bahr (1979) finds that more affluent women are less likely to remarry after divorce. See also Waite and Spitze (1981) for an investigation into determinants of female age at first marriage.

[8]A different approach is due to Choo and Siow (2006), who propose a statistic to directly measure the net gain from marriage for a given pair of male and female "types." They attempt to quantify the net benefit from marriage for men and women using data from the 1970 U.S. Census and Vital Statistics. They find that the net benefit of marriage declined between 1970 and 1980 for both men and women, but more so for women. This is suggestive, as opportunities in the labor market for women grew significantly during this decade.

[9]Haines (1996) shows an increase in the female age at first marriage up to about 1890. Fitch and Ruggles (2000) also find an increase in the female age at first marriage between 1850 and 1880; however, this increase is quite small, and seems to be concentrated in the years following the Civil War.

of the 19th century, declining male-to-female ratios likely contributed to the rising age at first marriage among women (Haines 1996; Hacker 2008). This would have been especially true in the years immediately following the Civil War.[10]

A small number of studies link women's economic opportunities to delayed marriage before the 20th century. Hacker (2008) offers evidence from the 1860 census that women tended to marry later in areas in which economic opportunities for women were greater; this is measured by local unmarried female labor force participation. Wanamaker (2012) links industrialization to declining fertility in the 19th century, with a focus on fertility within marriage. Goldin (1997) indirectly links economic opportunities to delayed marriage by noting a tendency for women's education and marriage to be mutually exclusive. She describes a "stark set of alternatives between career and family" (p. 1) for women born at the end of the 19th century, noting that approximately half of college-educated women graduating in 1910 were childless. While this references a somewhat later period, women's colleges in the late 19th century were similarly labeled "spinster factories" (Monahan 1951, p. 242). This paper's findings will provide further support for the idea that this mechanism contributed to the patterns observed during this period.

3 Institutional Background: Widows and the Civil War Pension Law

The original Civil War pension law, called the General Law, was passed on July 14, 1862. This act provided compensation for soldiers and the dependents of soldiers who had fought honorably for the Union and who had been wounded in such a way that they were unable to work. Over time, this pension system expanded into a form of old-age security for Union Army veterans and their families. Pension expenditures grew from $29 million in 1870 to $160 million by 1910, covering almost one million veterans and their dependents (Linares 2001). It is generally considered America's first large-scale social assistance program (Skocpol 1993; 1995).

Eligibility for a widow's pension under the General Law depended on three main criteria. A widow was entitled to a pension if she did not remarry, and if her husband had served honorably in the Union army and died of a disease or injury sustained in the service. The qualifying widow of a private in the Union Army was entitled to eight dollars per month plus two dollars per minor

[10] See Abramitzky, Delavande and Vasconcelos (2011) for an analysis of the effect of sex ratios on assortative matching in post-WWI France.

child (under the age of 16) beginning on July 25, 1866.[11] To give a sense of the size of this income, a typical daily wage for a common laborer in the north was approximately one dollar in 1860 and two dollars in 1870; including room and board, a farm worker would typically make 11to 15 dollars per month in 1860 and 18 to 20 dollars per month in 1870 (Margo 2000). If a widow remarried, she lost her right to a pension. Entitlement to the pension then passed to the soldier's minor children, who were allowed to receive it until the youngest turned sixteen.

The pension law was amended at various times. The most significant amendment was the act of June 27, 1890, which changed the eligibility requirements for both veterans and widows. Under this law, a widow could claim a pension if her husband had served honorably for at least 90 days in the Union Army, regardless of how he died. However, she had to demonstrate that she was "dependent upon her daily labor for support" (Linares 2001). Under the act of July 14, 1862, widows permanently lost their right to a pension if they remarried. However, later changes to the General Law altered this somewhat. As of June 7, 1888, a widow who had remarried could apply for a General Law pension in arrears, commencing on the date of her first husband's death and terminating on the date of her remarriage. On March 3, 1901, a widow who was eligible under the General Law but had remarried was allowed to be restored to the pension rolls after her new husband died, provided she had never divorced this second husband, and she was needy. It became progressively easier for remarried widows to be restored to the rolls through the 1920s (Glasson 1900).

3.1 Procedures for Pension Applications

The process of applying for pensions was costly and time consuming. In contrast to soldiers who filed pension claims, widows did not need to be examined by a surgeon; however, they were required to provide a great deal of evidence in support their claims. A widow had to appear before a court of record. If she lived more than 25 miles from a court of record, she could appear before a pension notary stationed in her locality (Oliver 1917). Here, she would make her declaration, which involved filling out a form in the presence of witnesses. The instructions attached to this form outline the information and documents she was required to furnish:

> She must prove the legality of her marriage, the death of her husband, and that she is still a widow. She must also furnish the names and ages of her children under sixteen years of age, at her husband's decease, and the place of their residence... The legality of the marriage may be

[11] Glasson (1900; 1918); Song (2000). Officers' widows were entitled to a larger pension than widows, but the UA data contains only privates.

> ascertained by the certificate of the clergyman who joined them in wedlock, or by the testimony of respectable persons having knowledge of the fact, in default of Record evidence. (Widow's Certificate No. 8,336).

This evidence was mailed to the pension bureau in Washington, DC, where claims were adjudicated. This adjudication process involved obtaining the soldier's military record from the war department. If a widow could not prove that she was legally married to the soldier or that his death was a direct result of his military service, her claim would be rejected.

In many instances, claimants hired attorneys to prosecute their claims. The quality of the attorney could have a dramatic effect on the speed with which a claim was processed; there are ample instances of claims pending for years because of attorney neglect, a problem well known to the pension board. The 1883 annual report of the pension commissioner condemns the behavior of these pension lawyers:

> There are certain ignorant, unscrupulous, and useless persons, whose only object seems to be, first, to procure applications from soldiers, regardless of merit, to be filed through them, and then, while acting simply as transmitters of the papers, assiduously dun the claimant until the ten-dollar fee is secured, and thereafter practically abandon the case (United States Pension Bureau 1883, p. 16).

3.2 Fraud

A challenge associated with using information about marital status from pension records is accuracy. Widows had a clear incentive to hide remarriages from the pension board, since disclosing this information would result in loss of pension. The incentive to fabricate marriages to veterans also existed. As the 1872 annual report of the pension commissioner remarks, "So long as pensions are to be granted upon evidence which (except record evidence) is purely *ex parte*, so long frauds will continue to exist" (United States Pension Bureau 1872, p. 13). The pension bureau was especially concerned about widows' claims: "The evidence to sustain a widow's or dependent's case is purely *ex parte*. As a result of this, a very considerable percentage of those cases are wrongfully established" (United States Pension Bureau 1872, p. 13).

If the pension authorities suspected a fraud, they would send a special examiner to the widow's place of residence to conduct an investigation. If found guilty of fraud, the widow lost her pension. Fraud was usually reported by either the postmaster who oversaw the delivery of pension vouchers and checks, or by members of the pensioner's community. There are a handful of examples in my sample of both sources reporting frauds[12]. However, notwithstanding the pension bureau's concerns

[12] A letter of instruction to a special examiner in the case of Catherine Matthews describes allegations of remarriage

about fraud, there is little evidence that hidden remarriages were a frequent occurrence. Women receiving pensions regularly interacted with the pension board throughout their lives; yet, in only 20 out of the almost 800 cases analyzed in this study is there any indication of investigation into pension fraud. Moreover, only a few of these cases resulted in the widow being stripped of her pension. Still, to address concerns about fraud, I check marital status using links to the federal censuses of 1870 and 1880. Unless a large number of women were engaged in an elaborate fraud involving hiding second husbands from census enumerators, hidden remarriages or cohabitation do not appear to pose a significant problem.

4 Data

4.1 Pension and Military Records

Data used in this paper comes from three main sources, two of which are newly collected from primary sources. The first data source is the Union Army (UA) database created by the Center for Population Economics (CPE) at the University of Chicago.[13] I have chosen a random sample of approximately 800 women who were married to soldiers in the UA database. Useful for this study, this database provides information about soldiers' families, including when, where, and to whom they were married, as well as the birth dates and names of their children. I use this information to identify women that meet two important conditions. First, I restrict attention to women widowed by 1880. This is because I expect such women to be most representative of the unmarried female population; they will be relatively young and thus more plausible marriage candidates. I choose

by the postmaster of Malone, New York. The examiner is instructed to ascertain "whether the pensioner, by regular ceremony, by cohabitation, or by any other manner has performed such an act as will constitute marriage (re-marriage) under the laws of New York" (Widow's Certificate No. 6,916). Another example of fraud is the case of Maria van Buren, whose remarriage to Frank Stoffer is reported to the pension board by a close acquaintance. An excerpt from the examiner's report reads, "Stoffer had in his possession several letters, written in the same chirography, with the one hereto attached, none having a signature, all about equally dirty, but differing vastly in tone and purpose. The first a threatening message, demanding that she return to him by 7 o'clock and at least bid him farewell 'like a lady,' or he would have her in the penitentiary immediately. The next, breathing undying attachment of enormous dimensions, and asking her forgiveness for having 'told on her'. The third a sarcastic letter to Stoffer, and the fourth a letter of farewell and filled with threats of vengeance for her rejection of his 'ardent heart.' Mrs Van Buren acknowledged that she was living wtih Stoffer, and had done so 'off and on when she felt like it', but denied that she had married him, denied that he is Van Buren, who is now, she remarked, if not in heaven, certainly not on earth; denied that she intended to run away and professed several times an unusually strong desire to be arrested. I was, of course, satisfied that the case was not one which I was authorized to further investigate without direct instruction" (Widow's Certificate No. 23,529). She was ultimately removed from the pension rolls because of remarriage, demonstrated by "cohabitation and recognition" (Widow's Certificate No. 23,529).

[13]These data were collected as part of the project Early Indicators of Later Work Levels, Disease, and Death, sponsored by the National Institutes of Health and the National Science Foundation (Federal grant number P01 AG10120; see Fogel 2000). See data appendix for further details.

1880 as a cutoff because it facilitates the linking of my observations to the 1880 census.[14]

The second restriction is that the widow had to apply for a pension within five years of her first husband's death. This restriction is intended to minimize sample selection bias due to limited data availability. Ideally, one would observe the widows of all soldiers in the UA database. However, because of the nature of this data source, the availability of spousal information depends on actions taken by subjects. For soldiers who died before 1880, all such information comes from dependents' pension applications, the vast majority of which are widows' applications. As such, it is extremely rare to observe a widow who never filed for a pension.[15] Women who first applied for a pension, say, ten years after widowhood will be those who had not applied earlier *and* had not remarried during those ten years. This will be a highly selected sample of *all* widows who did not file for a pension before ten years had elapsed. Given that my sample is necessarily restricted to applicants, there is a certain amount of selection that is unavoidable; however, I expect including late applicants to exacerbate this problem.

The majority of the information I use in this paper comes from data that I have collected from the Civil War pension files at the National Archives in Washington, DC. The CPE project focuses on soldiers' outcomes, so the UA database does not follow widows and children after the soldier died. After drawing my sample, I collect information about widows' pensions and marriage histories from their pension files. See appendix B for details of the data collection process. Because these data are compiled from historical records and not from surveys designed to avoid selection bias, the source of every piece of information is important. With this in mind, I will explain in detail where my most important variables come from.

The pension information is largely straightforward to collect, as any action a widow took with respect to pensions is recorded in her correspondence with the pension bureau. The case files contain all materials in a widow's pension application, which includes her application form and supporting evidence. If the widow was granted a pension, her file will contain both a pension brief and a pension certificate, indicating the amount of the pension, the effective start date, the date at which the pension was granted, the agency she was to be paid from, and the name of her attorney.[16] If the

[14] I cannot link widows to the 1890 census, because these manuscripts were lost in a fire. Linking to the 1900 census is less useful, as most Civil War widows were well past the age at which they could reasonably expect to remarry by 1900. The importance of census links is described later in this section.

[15] Soldiers on the pension in 1898 were required to inform the pension bureau of the name of their spouse and children. Before 1898, it is possible to have spousal information about a soldier if his widow never filed a claim but his mother or children did; however, this is quite rare.

[16] This information can be independently verified using the index to the pension files, which indicates the number attached to the widow's application and pension certificate. As these numbers are issued chronologically, the approximate date of application and issuance of the certificate can be inferred from these numbers.

widow did not receive a pension, it can be difficult to determine why. In later years, rejected claims contain a brief indicating the date of and reason for rejection; however, during the years immediately following the Civil War, information about rejection merely consists of a stamp somewhere in the file that reads "rejected." In such cases, it is impossible to determine the reason for or date of rejection. Similarly, if a widow abandoned her claim, I cannot be certain why or when.

Information about a widow's remarriage is slightly more complicated. Figure 1 illustrates the possible pension and marital outcomes for women in my sample. The first thing that occurs is the widow's pension application. After applying, the widow may remarry or die before her claim is adjudicated.[17] Otherwise, she will receive a decision from the pension board, which may be favorable or not. After receiving this decision, the widow may or may not remarry. The outcome of a pension application is always certain; however, in 20 percent of cases it is impossible to determine whether or not the widow ever remarried.[18]

A widow's remarriage is observable if her children filed a minors' a pension claim or she applied to be restored to the pension rolls under the act of March 3, 1901.[19] A widow's failure to remarry is observable if her death date is known, and there is no indication of remarriage. If she was receiving a pension when she died, her file will often contain a card indicating that she was dropped from the pension rolls due to death. If not, this information may come from minors' pension applications or other correspondence with the pension board. Marital status is not observable if the widow stopped communicating with the pension board some time before her death. The fact that knowledge of marital status is contingent on potentially endogenous actions taken with respect to pensions is of obvious concern and will be important to the sensitivity analysis I do later on.

Table 1 presents summary statistics from the pension file data I have collected (791 records in total). All women in this sample applied for a pension within five years of widowhood and had not remarried before doing so. The average age when widowed is 32; however, this ranges from 15 to 73. There are 625 women for whom remarriage status is certain, meaning that I observe them either remarrying or dying while single. There is no evidence that the other 166 women either remarried or died. Of these 625 women, 55 percent remarried at some point in their lives, which implies that the true fraction of women who ever remarried is between 43 and 64 percent. Of the 672 women for

[17] If a widow remarried with a pending claim, she was still entitled to be paid from the date of her widowhood to the date of her remarriage, provided she had applied for the pension before remarrying.

[18] After around 1880, the pension bureau started including records of pensioners being dropped from the rolls for any reason. Women whose marital status is unknown are missing these records; thus, if they were on the pension, it is likely that they died, remarried, or stopped collecting their pensions some time before 1880.

[19] In some cases, a widow may have filed a claim for a pension she was not entitled to, or there may have been some other correspondence with the pension board indicating that she had remarried.

whom this information is available, 16.5 percent remarried before receiving a pension.[20] On average, a woman who remarried did so 4.3 years after her first husband's death. This average is much lower among women who remarried before getting a pension (2.5 years), which is unsurprising. It is, however, suggestive that the average time that elapsed between receiving a pension and remarriage is 3.7 years, which is much greater than 2.5 years.

The average amount of time that elapsed between the soldier's death and his widow filing for a pension was eight months, and the median was less than four months. The probability of ever having a General Law claim accepted was 0.88; however, fewer than 80 percent of women were receiving a General Law pension within five years of applying. The average processing time for a pension was more than two years, although this is highly skewed: the median processing time is slightly less than one year. Most women in my sample were first married during the 1850s and were widowed during the war. These women tended to come from the Mid Atlantic region (31 percent) or the East North Central region (42 percent). Very few come from Southern or Western regions.

Finally, I use information from the pension file data to link my observations to the federal censuses of 1870 and 1880. The primary reason for linking widows to the census is that it provides a check on the marriage information available in the pension data. For one thing, these links allow independent verification of widows' marital status, which alleviates concerns about inaccuracies due to fraud. These links also allow me to observe other characteristics of the widows in my sample, such as birthplace. Links are performed manually using the genealogical website ancestry.com, and 70% of the women in my sample are successfully matched to the 1870 census, the 1880 census, or both.

5 Empirical Framework

5.1 Source of Variation in Pension Income: Theoretical Justification

The major challenge associated with measuring the effect of widows' pensions on the timing of remarriage is locating an appropriate source of variation in pension income to exploit. Because pension amounts are standardized, there is no variation in pension income among pensioners. Moreover, it is not straightforward to compare women who had pensions to those who did not, as I do not observe women who never make pension applications. There are two possible sources of variation in pension income: the pension board's decision and the timing of this decision.

[20] Even if I do not know whether or not a widow *ever* remarried, I may know that she did not remarry with a pending claim if she communicated with the pension board subsequent to her claim being granted.

The pension board's decision is not an ideal source of variation for a few reasons. First, this variable is only defined for women who complete their claims. Recall from figure 1 that at least twelve percent of my sample remarried while their claims were pending. A simple comparison between women with accepted and rejected claims will discard this potentially valuable information. Another issue is that rejections take significantly longer to process than acceptances. It takes approximately five years longer to reach the "rejected" node in figure 1 than the "accepted" node. Thus, my sample of rejected widows ought to look very different from the universe of potentially rejected widows, as many of these are likely to have remarried before the board's decision was rendered. A final technical issue has to do with accuracy: it is often unclear when or why a claim was rejected.

Because of these issues, I use variation in the timing of the pension board's decision, rather than the outcome, to estimate the effect of pensions on the timing of remarriage. Specifically, I look for a treatment effect of having a pension claim granted, or of transitioning from having a pending claim to an accepted claim. So long as there is uncertainty about if and when a pension claim will be approved, and discounting of future income, having a claim accepted will represent a real positive utility shock relative to having a pending claim. In appendix A, I develop a simple search model of marriage and pensions, in which I show that, in the presence of this type of uncertainty, widows with accepted claims will have higher reservation match qualities and will spend less effort searching for mates than widows with pending claims. As such, the rate of remarriage should shift discontinuously downward at the moment a pension is granted. This effect should be augmented in the presence of liquidity constraints.

5.2 Empirical Approach: Details

To evaluate the effect of the pension on the rate of remarriage, I estimate a proportional hazard model of both pensions and marriage, allowing the rate of remarriage to shift at the moment a pension is granted. Variation in processing times allows me to observe women with and without pensions at every point in time, which allows me to estimate a hazard rate of remarriage that differs by pension status. Some of this variation is plausibly exogenous. For example, idiosyncrasies in the postal service, clerical errors, or unexpectedly capricious behavior on the part of pension attorneys certainly affected processing times in a random fashion. However, a portion of the variation in processing times is likely endogenous to marital outcomes. For example, women with poor marriage prospects may have been more invested in getting a pension because they knew their alternatives were poor. So, those who got pensions quickly may have tended to remarry slowly because of poor marriage prospects, not because of a causal effect of the pension. Another concern is that processing times are highly correlated with the quality of a pension claim: rejections take significantly longer to process than acceptances.

Why is this a threat to identification? If dying in the war is random, the pension legibility should random too; as such, the ambiguity of a claim should be similarly exogenous. However, bias may be introduced by the decision to apply. Applying for a pension is costly: a widow will choose to incur this cost if the benefit is great enough. The expected benefit from applying is lower for a widow with an ambiguous claim, as the probability of ever receiving a pension is low. Thus, women who apply with ambiguous claims may be systematically different from women who apply with straightforward claims. In particular, they may have worse alternatives, either financially or in the marriage market. The direction of this bias on the timing of remarriage is unclear: women with poor alternatives might receive fewer proposals per unit of search effort; however, they might also be less selective.

To overcome these endogeneity problems, I use a method developed by Abbring and Van den Berg (2003a; 2005). This is a novel approach to identifying treatment effects in the presence of an endogenous treatment when both the treatment and outcome are duration variables. The approach involves jointly estimating the hazard rates of pensions and remarriage, allowing for correlation between the unobserved heterogeneity in these two risks. The hazard rate at time t refers to the probability of realizing an outcome (pension or marriage) at t, conditional on not having realized it earlier. The hazard rate of pension income is given by

$$\theta_p(t|X, v_p) = \lambda_p(t)\exp(X\beta_p + v_p) \tag{1}$$

and the hazard rate of marriage is given by

$$\theta_m(t|X, v_m, t_p) = \begin{cases} \lambda_m(t)\exp(X\beta_m + v_m) & \text{if} \quad t \leq t_p \\ \lambda_m(t)\exp(X\beta_m + \delta + v_m) & \text{if} \quad t > t_p \end{cases} \tag{2}$$

For each $i \in \{p, m\}$, λ_i is the baseline hazard function, which characterizes duration dependence, and X is a matrix of explanatory variables that may shift the hazard rate. The term t_p represents the time at which a pension is granted, and v_i reflects unobserved heterogeneity.

Allowing for duration dependence ($\lambda_i(t)$) and the effect of covariates ($X\beta_i$) is crucial to the identification of δ. Duration dependence refers to the way the hazard rate changes over time, i.e.

whether marriage or pension receipt becomes more or less likely as time passes. Failing to account for duration dependence will bias the estimate of δ. For example, suppose there is negative duration dependence in the rate of remarriage, so the probability of remarrying declines with time in the marriage market. Then, women will appear to remarry at a slower rate upon receiving a pension, simply because these women will have been in the marriage market longer. This will lead to an overestimation of δ. Failure to account for observables will bias the estimate of δ to the extent that these are correlated with pension status. For example, suppose the hazard rate of pension receipt increases with age, and the hazard rate of marriage declines with age. If I do not control for age when estimating δ, the estimate will be biased away from zero, as women who receive pensions quickly will tend to be older, and these women will tend to remarry slowly.

These concerns apply to any standard proportional hazards model. An additional issue that arises in this particular setting is the possibility that v_m and v_p are correlated. For example, if v_m and v_p are negatively correlated, the estimate of δ may be negative even if the true δ is zero. Correlated unobserved heterogeneity generates bias in a similar fashion to omitted observable controls. If women who get pensions quickly tend to have large values of v_p, they will also tend to have small values of v_m, which means they are likely to take longer to remarry even if the pension itself has no causal effect.

Abbring and Van den Berg (2003a; 2003b) show that this model is identified even if v_m and v_p are correlated. Moreover, it is identified without exclusion restrictions or assumptions about the functional form of either the baseline hazard or the joint distribution of the unobserved heterogeneity terms; the necessary assumption is simply proportional hazards. The unobserved heterogeneity directly affects the *rate* of treatment but not the precise timing of treatment. Put another way, a high v_p raises the probability of receiving a pension at time t; however, there remains a stochastic element to which event, pension or no pension, actually occurs at time t. The problem is disentangling this random assignment from the non-random assignment.

To understand how this is possible, first notice that, in a simple proportional hazards setting, the distribution of unobserved heterogeneity is identified from variation in observables. To see this, consider the rate of pension receipt. Suppose one woman has a very good pension attorney (high $X\beta_p$), and a second woman has a poor pension attorney (low $X\beta_p$). Now, suppose these two women both take a long time to receive a pension (large t_p). We can infer from this that the probability that the first woman has an ambiguous pension claim (low v_p) is higher than it is for the second woman. In general, the distribution of v_p, conditional on t, depends on observables, which allows its distribution to be pinned down.

How does this help to identify correlated unobserved heterogeneity in the rates of remarriage and pension receipt? Using the same example, suppose that the quality of pension attorney has no direct effect on the rate of remarriage, so women with good and bad pension attorneys have the same $X\beta_m$.[21] This means that we should not expect to see systematically different marital outcomes by the quality of pension lawyer. However, recall that, conditional on t, the distribution of v_p is not independent of the quality of pension lawyer. So, if v_m and v_p are correlated, the distribution of v_m will similarly be dependent on pension lawyer quality. Say v_m and v_p are negatively correlated, and recall that, fixing t, $E(v_p)$ is higher for women with bad lawyers than it is for women with good lawyers. This means that, among women who are in the sample at time t, those with good lawyers will tend to remarry fastest, because these women tend to have higher v_m. Similarly, if v_m and v_p are positively correlated, women with bad lawyers will tend to remarry more quickly. In other words, different joint distributions of v_m and v_p will be observationally distinct. Once the correlation between v_m and v_p has been corrected for, the remaining difference between the marriage rate before and after a pension is granted can be interpreted as a causal effect of the pension.

I estimate this model by maximum likelihood. The survival function, or the probability of remaining a widow (m) or not having a pension (p) at time t, is denoted $S_i(t)$, and it has the following form:[22]

$$S_i(t) = \exp\left(-\int_{t_0}^{t} \theta_i(s)ds\right), \qquad i \in \{m, p\}$$

If t is a random variables denoting time an event occurs, its density is given by

$$f_i(t) = \theta_i(t)S_i(t)$$

So, the likelihood of an event occurring at t depends on both the hazard function and the survival

[21] This example is used for clarity and does not imply the necessity of an exclusion restriction for identification. In general, as long as $\beta_m \neq \beta_p$ and there is sufficient variation in the data, there exists some X, X' such that $X\beta_m = X'\beta_m$ but $X\beta_p \neq X'\beta_p$ (Abbring and van den Berg 2003a, 2003b, 2005). This is all that is required. Also notice that the values of β_m, β_p are identified using "early" parts of the sample, when v_m and v_p are independent of observables. This dependency arises "later" in sample, due to selective sample attrition. Note that identification relies heavily on the proportional hazards assumption.

[22] See Lancaster (1990).

function. For pensions, the survival function is straightforward to define:[23]

$$S_p(t|X, v_p) = \exp\Bigg(- \int_{t_0}^{t} \lambda_p(t) \exp(X\beta_p + v_p) \Bigg)$$

The survival function for marriage is somewhat more complicated, because it shifts at a point in time. The survival function before and after receiving a pension are given by the following two equations, respectively:

$$S_{m,1}(t|X, v_m) = \exp\Bigg(- \int_{t_0}^{t} \lambda_m(t) \exp(X\beta_m + v_m) \Bigg)$$
$$S_{m,2}(t|X, v_m, t_p) = S_{m,1}(t_p|X, v_m) \times \exp\Bigg(- \int_{t_p}^{t} \lambda_m(t) \exp(X\beta_m + \delta + v_m) \Bigg)$$

To understand the definition of $S_{m,2}$, consider the meaning of its two parts separately. Suppressing X and v_m, the first term reflects $Pr(t_m \geq t_p)$, and the second term reflects $Pr(t_m \geq t | t_m \geq t_p)$.

There are four possible outcomes for women in the sample, which I index below by $k \in \{1, 2, 3, 4\}$. A woman can remarry before she gets her pension ($k = 1$); she can remarry after her claim is granted ($k = 2$); she can be censored before her claim is granted, meaning that she dies or disappears from the sample ($k = 3$); or she can be censored after her claim is granted ($k = 4$). Each of these events is associated with a different likelihood. Conditional on her unobserved heterogeneity terms, the likelihood contribution of woman i can be written as

$$L_i(t) = \begin{cases} \theta_m(t|X, v_m, t_p) S_{m,1}(t|X, v_m) S_p(t|X, v_p) & \text{if} \quad k = 1 \\ \theta_m(t|X, v_m, t_p) S_{m,2}(t|X, v_m, t_p) \theta_p(t_p|X, v_p) S_p(t|X, v_p) & \text{if} \quad k = 2 \\ S_{m,1}(t|X, v_m) S_p(t|X, v_p) & \text{if} \quad k = 3 \\ S_{m,2}(t|X, v_m, t_p) \theta_p(t_p|X, v_p) S_p(t_p|X, v_p) & \text{if} \quad k = 4 \end{cases}$$

To estimate this model, I make certain parametric assumptions about the baseline hazard rate and the joint distribution of the unobserved heterogeneity terms, v_m and v_p. I attempt to make the least restrictive parametric assumptions possible. For the baseline hazard, I use a piecewise constant function, where time is divided into discrete "bins," and $\lambda(t) = \lambda_t$ takes on some unrestricted value in each of these bins. I use bins of one year, with a single bin for the tail of the time distribution, extending from $t = 8$ until the last observation leaves the sample. Following eight

[23]This construction follows Abbring and van den Berg (2005), who apply this model to evaluating the effect of unemployment insurance sanctions on the rate of transition to employment.

years after widowhood, first marriages and pensions occur with insufficient frequency to identify hazard rates at finer intervals. Following Abbring and Van den Berg (2005), I assume that the unobserved heterogeneity terms both obey a discrete distribution with two unrestricted mass points: $v_m \in \{v_m^{low}, v_m^{high}\}$ and $v_p \in \{v_p^{low}, v_p^{high}\}$. Thus, there are four possible combinations of v_m and v_p, each of which is associated with a certain probability. The location of each of these mass points and the probability of each combination of the two are estimated in the model. A discrete distribution is considered the most flexible parametric assumption that can be made about the joint distribution of unobserved heterogeneity terms, as it allows any correlation between the two variables to be achieved.[24] I estimate the model parameters using the EM algorithm (Heckman and Singer 1984).

6 Results

Before presenting estimates of the model described in section 5, it is useful to get a sense of what the hazard rates of remarriage and pension receipt look like. Figure 2 plots the empirical hazard rate of both pensions and remarriage, estimated non-parametrically using a kernel method.[25] The top panel illustrates the rate of remarriage measured before and after a pension is granted; the bottom panel illustrates the hazard rate of pension decisions. Time is measured in years since widowhood; however, individuals do not enter the sample until they apply for a pension. Notice that, for the first four years, the rate of remarriage for women who have not yet received a pension lies uniformly above that of women who have pensions. After four years, the two lines are close to one another, with the rate of remarriage slightly lower for women with pending claims. This may indicate that the pension only lowers the rate of remarriage in the short run; however, it may also reflect differences in the characteristics of pensioned and unpensioned women in later years. It is important to note that these empirical hazard rates are calculated without controlling for observable or unobservable characteristics.

Table 2 contains parameter estimates for the model described above, with the estimated effect of covariates on the rate of pension receipt listed next to their estimated effect on the rate of remarriage. In column (1), I estimate the model with no covariates or correction for correlated unobserved heterogeneity. In this specification, the estimated effect of the pension is barely negative and not significantly different from zero. In column (2), I add covariates to the hazard rate of both risks, which significantly increases the magnitude of the estimate, to -0.269 (0.154), which is significant at

[24] Heckman and Singer (1984); Abbring and Van den Berg (2005); Van den Berg (1996).

[25] This is done using the STS package in STATA.

the 10% level. This suggests that selection on observables biases this effect toward zero: observable characteristics of women with ambiguous claims tend to slow the rate of remarriage, leading to an underestimate of the effect of the pension when these controls are omitted.

In column (3), I introduce the possibility of correlated unobserved heterogeneity in the rates of pension receipt and remarriage. At -0.283 (0.159), the estimated effect of the pension changes little from the previous specification, suggesting that much of the selection problem is captured by the controls for covariates. The estimate from the full model can be interpreted to mean that receiving a pension lowered the hazard rate of remarriage by 25%.[26] This estimate implies that, for a woman with median characteristics, immediately granting her a pension would raise her median time to remarriage from 5.7 to 9.2 years, an increase of 3.5 years.[27] This timing increase is consistent with the summary statistics from table 1, although the implied medians are substantially higher than they are in this table, as they should be. These summary statistics are calculated using women who actually remarry. The medians implied by the model estimates incorporate information from women who never remarry, which will tend to raise them substantially.

Other variables affect the rate of remarriage in plausible ways. Older women tend to remarry more slowly, as do women with more children. The year of widowhood has a negative effect on the rate of remarriage, which may reflect sample selection, as claims become more ambiguous the further removed the soldier's death is from the war. The county male to female ratio speeds up remarriage quite significantly. The only variables that significantly affect the hazard rate of pension income are year of widowhood and time to pension application, which presumably reflects the fact that claims become more ambiguous with distance from the war and distance from the soldier's death. There are also regional differences: claims from the New England were processed significantly faster than claims from the Mid-Atlantic, the Midwest or the South.

The parameters of $\lambda_m(t)$ and $\lambda_p(t)$ are also listed in table 2, with λ_m and λ_p on the interval [0,1) both normalized to 1. These estimates suggest non-monotonic duration dependence in both risks. In both cases, the hazard rate initially increases and then falls. One can imagine plausible explanations for this pattern in the hazard rate of marriage. The rate of remarriage may rise in the short run if women lower their reservation match qualities as time passes, either due to revised expectations

[26] This comes from the fact that $\theta^{PEN}/\theta^{NOPEN} = \exp(-0.283) = 0.75$, so $\frac{\theta^{PEN} - \theta^{NOPEN}}{\theta^{NOPEN}} = -0.25$.

[27] For women with pensions, this calculation is done by solving the following for t_{med}:

$$0.5 = Pr(t \geq t_{med}) = S_2(t_{med}|X, v_m)$$

For women without pensions, I do the same calculation, replacing S_2 with S_1. For X, I use median characteristics and mean regions; I integrate over v_m and v_p using estimates from the model.

or changing preferences for matching. However, this rate is likely to fall eventually if part of what makes women desirable in the marriage market is fertility. In the case of pensions, this pattern may reflect changes in the composition of claims as time passes. Among very straightforward claims, the probability of receiving a pension is likely to increase with processing time. However, at some point, all straightforward claims will have been processed, leaving only ambiguous ones. The probability of ever getting a pension with an ambiguous claim is low.

The unobserved heterogeneity terms are very imprecisely estimated. In fact, the two mass points in the distribution of v_p converge to indistinguishably similar values, which means that it is impossible to calculate standard errors for the probability of observing each of these values. Because of this, I have restricted both mass points to take on the same value, which follows Abbring and Van den Berg (2005). This likely indicates that unobserved heterogeneity in the rate of pension receipt is well controlled for by covariates and the duration dependence function, leaving little systematic unobserved heterogeneity.

In table 3, I estimate heterogeneous effects of receiving a pension on the rate of remarriage by interacting the effect of the pension with different observable variables: age, number of children, county male-to-female ratio, population density, a measure of the widow's wealth, and region of residence. Continuous variables are demeaned, so the estimated δ should be interpreted as the effect at the mean value of the interaction variable. The widow's wealth measure is not taken from the pension file data, as the pension bureau did not ask about a pension applicant's means. However, the Union Army database links soldiers to the 1860 census, and these links contain information about the soldier's wealth (or the wealth of the head of the soldier's household in 1860). I use this as a rough measure of the widow's wealth after the soldier dies. Only two thirds of my sample has been linked to the 1860 census, so I do not use this variable in the baseline specification.

There is some evidence of heterogeneity in the effect of the pensions across regions, although there is no evidence of heterogeneity by geographic characteristics related to marriage market conditions, i.e. sex ratios or population density. The interaction with first husband's wealth is insignificant, but the point estimate is positive, which suggests that the rate of remarriage among wealthier widows may have been less influenced by pension receipt. This may reflect the fact that the pension represented less of a shock to the utility of wealthier women, or that wealthy women were less bound by liquidity constraints. The strongest result comes from interacting the effect of the pension with the widow's age and number of children. Receiving a pension has a significantly larger effect on older women and women with more children; while this is not shown, the inclusion of interactions with age and number of children together reveals that the interaction with number of children largely works through age (as older women tend to have more children).

The results in first column of table (3) indicated that, for 32 year old widows, receiving a pension causes the hazard rate of remarriage to fall by 0.566 (0.176), and that this effect grows in magnitude by 0.063 (0.017) with every additional year of age. The difference in effect by age is quite striking: for a 25 year old woman, receiving a pension causes the median time to remarriage to increase very little, from 4.1 to 4.6 years. However, for a 35 year old woman, the effect of the pension is to increase the median time to remarriage from 7.6 to 71.3 years. This can be interpreted to mean that the median 35 year old woman who receives a pension is predicted not to remarry: receiving a pension lowers the probability that a 35 woman has remarried within 10 years of widowhood from 0.52 to 0.30.

7 Sensitivity Analysis

7.1 Instrumental Variables Analysis

The hazard model described in section 6 is the most exact representation of the relationship between the receipt of pensions and the rate of remarriage. However, it is possible that the estimates are sensitive to some of the assumptions required for identification, namely the proportional hazards assumption. So, as a complement to the analysis in section 6, I include a linear analysis of the relationship between pensions and the timing of remarriage.

Using a series of time frames ranging from one to five years ($\tau \in \{1, 2, 3, 4, 5\}$), I create an indicator variable equal to one if a widow had received a pension within the time frame ($I(t_p \leq \tau)$) and an indicator equal to one if she had remarried within the time frame ($I(t_m \leq \tau)$). I estimate the following by OLS:

$$I(t_m \leq \tau) = \alpha + \beta I(t_p \leq \tau) + X\gamma + u$$

The matrix X includes all controls used in section 6. If pensions discourage remarriage, I should find $\beta < 0$. Here, the endogeneity problem is quite severe: many women who were not receiving pensions within, say, three years of applying had been denied pensions *because* they had remarried. I use instrumental variables to circumvent this problem.

Details of the application and review process provide potentially valid instruments for pension income.[28] The instrument that I use is based on the spelling of last names. To receive a pension,

[28] This approach is similar in spirit to Maestas, Mullen and Strand (2013) who use spending allowances of the

a widow had to prove that she was married to a soldier, that he served honorably in the military, and that his death was connected to the service. This involved locating military service records, hospital records, and marriage certificates. If there were discrepancies in the spelling of his surname in these records, additional steps were required to demonstrate that the records referred to the same individual. In the pension files, there are examples of secondary affidavits explaining name spelling discrepancies.

I construct an indicator of name spelling homogeneity from the one percent IPUMS samples from 1860, 1870, and 1880. I compile a list of all household heads in each of these years, and I group last names by codes generated using the New York State Identification and Intelligence System (NYSIIS) algorithm (Atack and Batemen 1992). Frequently used to create linked census samples, this algorithm collects names into phonetically similar groups.[29] I construct a Herfindahl index of the dispersion of unique name spellings within these phonetic groups. Greater values indicate that there is little variation in name spelling; smaller values indicate that names in this group are spelled in many different ways. I perform two tests of the validity of this measure. First, I check whether or not a low name homogeneity index predicts multiple spellings of the veteran's last name in the pension data. I find that a one standard deviation increase in this index raises the probability of observing multiple surname spellings in the pension data by 8.5 percentage points; this is highly significant. Second, I check whether or not a name with a high homogeneity index is more likely to exactly match the most common spelling in its phonetic group in the census. Again, I find that a one standard deviation increase in the index raises the probability of such a match by 25 percentage points, which is also highly significant.

A concern is that this measure may not be exogenous to marital outcomes. Names that belong largely to immigrants may be spelled in multiple ways, and immigrant status is likely endogenous to marital outcomes. Names that belong to lower socioeconomic status families may be frequently misspelled if the literacy rate is low among these families. Because there is no information on nativity or literacy in the pension data, I cannot control for these variables without restricting my sample to individuals linked to the census. However, I can control for average literacy, immigrant status and socioeconomic status, measured as the occupational income of the household head,[30] by phonetic name group in the IPUMS data. I include these controls to preserve the validity of the instrument.

examiners assigned to individual cases as an instrument for disability insurance to identify a causal effect of disability insurance on labor supply.

[29] Ferrie 1996; Abramitzky, Boustan and Eriksson 2012.

[30] See data appendix for an explanation of this variable.

Table 4 contains both OLS and 2SLS results. The OLS estimate is negative for all values of τ, but only significant at the five percent level when $\tau \geq 2$. The 2SLS estimates are also everywhere negative, but they are implausibly large in magnitude, and they are estimated very imprecisely: the estimates are only significantly different from zero when $\tau \geq 4$. The first stage F statistics are typically below 10, which suggests that the instrument might be weak. Thus, I present 90 percent Anderson-Rubin confidence intervals for the effect of the pension, which are robust to weak instruments.[31] In most cases, these confidence regions do not include zero. Given their imprecision, it is difficult to attach significance to the size of the 2SLS estimates. However, this analysis provides some corroborating evidence that the causal effect of pensions on the timing of remarriage is negative.

7.2 Alternative Sample Restrictions

An additional concern is that the results may be sensitive to the source of information on remarriage. Recall that knowledge of a widow's remarriage is contingent on her communicating in some way with the pension board. Specifically, I observe a widow's remarriage if her children file a minors' pension claim, or if she files a new claim under the act of March 3, 1901. If the source of information is distributed differently among women who remarry before and after obtaining a pension, and if the source of this information is correlated with marital outcomes, this might bias my results. For example, minors' pension applications are the source of evidence of remarriage in 71 percent of cases that occur before a pension is granted and 85 percent of cases that occur after a pension is granted. This means that my sample of women who remarry before receiving a pension may be disproportionately composed of childless women who lived to 1901. These women may be younger and healthier by construction, and thus better marriage prospects.

I use two alternative sample restrictions to address this concern. First, I restrict the sample to women who have children under the age of 16 when they are widowed. So, these women's remarriages are potentially observable through a child's pension application. Second, I restrict the sample to women who are observed to live at least to 1901. Death dates are only observable for women who are on the pension at the time of their death. As such, all remarried women in this second sample had to have applied to be restored to the pension roles under the act of March 3, 1901; all unmarried women in the sample would have to have remained on the pension roles for the entire sample period. This restriction forces information on marital status to come from the same place for all women; moreover, it eliminates any systematic health differences between women who remarry before or

[31] To calculate this confidence region, I use the condivreg command in Stata.

after receiving the pension.

To mitigate concerns about pension fraud, I restrict the sample to women who are successfully linked to the census of 1870 and/or 1880, and whose marital status is corroborated by these links. Women have an incentive to lie to the pension board about marital status; however, there should be no such incentive to lie to census enumerators. Another advantage of using a sample linked to the census is that it allows me to control for immigrant status. Finally, I restrict the sample to women whose husbands actually died during the war. Dying during the war is arguably more random than failing to recover from a non-life-threatening injury or disease contracted during the war, so it is worth verifying that the results are robust to this sample restriction.

I estimate the proportional hazards model described earlier under these sample restrictions, and the results appear in table 6. The baseline results, with and without a correction for correlated unobserved heterogeneity, are repeated in panel A. Results that include interactions between pension status and the widow's age and number of children are also included. The remaining panels contain results under the sample restrictions outlined above. The average effect of the pension is not particularly sensitive to these sample restrictions; however, the estimate often fails to achieve statistical significance, even at the 10 percent level. The truly robust result is the interaction effect with age. The effect of the pension is significantly negative for women in their early 30s, and age significantly increases this effect, in every specification except panel C, in which the sample is restricted to women who live until 1901. In this specification, the estimates are negative and similar in magnitude to the others, but the standard errors are very large; this may be due to the small sample size of 253. In general, these results indicate that the negative effect of the pension on marriage rates is very robust for older women with more children.

8 Implications and Discussion

The results presented above show a clear effect of the marriage penalty built into the Civil War pension on the marital outcomes of Union Army widows. Having a claim granted lowered the rate of remarriage by 25 percent overall, implying an increase in the median time to remarriage of 3.5 years. For women over 32, the pension lowered the rate of remarriage by more than 40 percent. My estimates imply that, if a typical 32 year old widow immediately received a pension, this would reduce the probability that she remarries within 10 years by 20 percentage points relative to an identical widow with no pension. These are striking results, for which I provide context and interpretation in this section.

The apparent heterogeneity in the effect of the pension is consistent with the hypothesis that the pension lowered marriage rates by making women more selective in the marriage market, or by reducing the effort they spent searching for husbands. Older women with more children may have had less favorable marriage prospects, which could mean that they needed to expend more effort at the margin to procure an acceptable match. This should generate a larger response, in terms of search effort, to the increase in utility the pension afforded these women. Similarly, if younger women with fewer children faced a more favorable distribution of match qualities, their reservation match qualities may have been closer to the lower tail of this distribution. This should cause pensions-induced changes in the probability of encountering a suitable match to be smaller for younger women.

These results offer a dim view of marriage for 19th century women. In particular, they suggest that many women during this period entered into marriages that were not preferable to an income stream barely above subsistence level. This may be surprising if we believe that social pressures to marry were greater in past eras. Moreover, opportunities for women outside the home were typically unappealing: domestic service and factory work were the most common forms of employment for women at this time. However, the results become less surprising in light of the potential costs associated with marriage. For example, the risk of death during childbirth was high. In 1900, childbirth was the second largest cause of death among women aged 15-44, with 1 in 118 mothers dying while giving birth (Albanesi and Olivetti 2013). Stringent divorce laws made it difficult for women to escape bad marriages or to remarry if their husbands deserted them (see Cvercek 2009 and Schwartzberg 2004).

What do these results tell us about aggregate marriage patterns during this period? Given the small effect of the pension for young, childless women, my findings suggest that, in order for economic opportunities for women to have had a meaningful impact on patterns of first marriage, these opportunities would have to have been more valuable than the Union Army pension. At the same time, my estimates are generated by a comparison between women who have been granted a pension and women who are *still waiting for a pension.* The rationale behind this approach is that there is uncertainty about if and when the pension claim will be granted, so discounting and the possibility of rejection should generate differences in behavior. However, if the data allowed a comparison between women with a pension and women *with no possibility* of a pension, these differences may be starker. Another interesting point to note is that the probability of rejection, at about 14 percent, is quite low. So, the results likely reflect a high discount rate, which suggests liquidity constraints.

9 Conclusion

This paper documents the effect of pension income on the marital outcomes of Union Army widows during the mid to late 19th century. Accounting for potential endogeneity of pension processing times to marital outcomes, I find that having a pension claim granted significantly lowered the hazard rate of remarriage, particularly for older women. These results are consistent with other research that finds a negative effect of marriage penalties on marriage rates. Moreover, the findings further our understanding of the profound demographic effects the Union Army pension had in the northern United States during this period.

More broadly, the results of this paper indicate that women's economic incentives mattered for marriage market outcomes in the 19th century. This is potentially informative about changes in first marriage that occurred over the course of this century. While the 19th century did not see as radical an increase in opportunities for women as the 20th, industrialization in the later part of the century facilitated women's work (Wanamaker 2012), as did the rise of the clerical sector beginning around 1890 (Goldin 1984). This is rarely cited in accounts of 19th century marriage patterns; however, it is quite possible that these opportunities contributed to the rising age at first marriage observed during this period. My results suggest that changes in women's opportunities would need to have been larger than the Civil War pension to have a discernible effect on the behavior of young, childless women; still, they strongly suggest that this channel merits further investigation.

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Tables and Figures

Table 1: Summary Statistics from Pension File Data

Variable:	Mean	Median	SD	Min	Max	N
Pension Variables						
Applied within 1 year	0.818	1.000	0.386	0.000	1.000	791
Time to first application	0.683	0.293	0.973	0.005	5.767	791
General law claim accepted	0.876	1.000	0.330	0.000	1.000	791
Processing time of accepted gen law claim	2.249	0.935	4.193	0.074	50.500	692
Age/Marriage Variables						
Age widowed	32.121	31.000	9.400	15.000	73.000	769
Age at first marriage	20.897	20.000	5.233	9.000	48.000	750
Age at remarriage	32.232	31.000	7.503	18.000	65.000	332
Number of children (first marriage)	2.564	2.000	2.148	0.000	13.000	791
Husband died during war years	0.716	1.000	0.451	0.000	1.000	791
Remarried	0.550	1.000	0.498	0.000	1.000	625
Remarried without pension	0.165	0.000	0.372	0.000	1.000	672
Time to Remarriage:						
All	4.338	3.375	3.542	0.230	26.036	340
Remarried with pending claim	2.460	1.838	1.910	0.230	8.778	110
Remarried after pension	5.236	4.351	3.786	0.860	26.036	230
Time to remarriage following pension	3.737	2.605	3.682	0.000	25.463	225
Calendar Years						
First marriage	1854.4	1856	7.953	1822	1879	778
Widowhood	1865.5	1864	4.492	1861	1879	790
Remarriage	1868.8	1867	4.851	1863	1889	340
Pension application	1866.2	1865	4.934	1862	1883	791
Pension certificate	1869.2	1866	9.250	1862	1928	724
Region of Residence						
New England	0.114	0.000	0.318	0.000	1.000	778
Mid Atlantic	0.314	0.000	0.464	0.000	1.000	778
East North Central	0.419	0.000	0.494	0.000	1.000	778
West North Central	0.093	0.000	0.290	0.000	1.000	778
South Atlantic	0.024	0.000	0.154	0.000	1.000	778
East South Central	0.033	0.000	0.180	0.000	1.000	778
West South Central	0.001	0.000	0.036	0.000	1.000	778
Mountain	0.000	0.000	0.000	0.000	0.000	778
Pacific	0.001	0.000	0.036	0.000	1.000	778

Sample includes women who were widowed before 1880 and who applied for a pension within five years of widowhood. Sample drawn from Union Army Database (Fogel et al 2000). Data collected from Civil War pension files at the National Archives in Washington, DC.

Table 2: Determinants of the Hazard Rate of Remarriage and Pension Receipt

	(1)		(2)		(3)	
Outcome:	Remarriage	Pension	Remarriage	Pension	Remarriage	Pension
Effect of pension	-0.036		-0.269*		-0.283*	
	(0.130)		(0.154)		(0.159)	
Age at widowhood			-0.094***	0.004	-0.096***	0.004
			(0.012)	(0.008)	(0.014)	(0.009)
Number of Children			-0.067	-0.021	-0.072	-0.021
			(0.047)	(0.029)	(0.052)	(0.030)
Year of widowhood			-0.056***	-0.071***	-0.057***	-0.071***
			(0.020)	(0.013)	(0.021)	(0.013)
Time to pension application			0.048	-0.164**	0.034	-0.164**
			(0.086)	(0.076)	(0.094)	(0.076)
Potential minor pension at widowhood			0.056	0.136	0.061	0.136
			(0.130)	(0.096)	(0.141)	(0.096)
No pension attorney			0.228	0.269	0.237	0.269
			(0.205)	(0.164)	(0.214)	(0.164)
Washington pension attorney			0.142	-0.084	0.145	-0.084
			(0.162)	(0.132)	(0.169)	(0.133)
First husband: age at death			0.017	-0.009	0.018	-0.009
			(0.012)	(0.009)	(0.013)	(0.010)
First husband: log occupational wage			0.193	-0.244	0.220	-0.244
			(0.353)	(0.238)	(0.427)	(0.243)
First husband: height (feet)			-0.439	-0.214	-0.448	-0.215
			(0.281)	(0.219)	(0.300)	(0.223)
County male-to-female ratio			2.428***	-0.254	2.539***	-0.254
			(0.875)	(1.004)	(0.903)	(1.077)
County percent urban			0.396	0.273	0.401	0.273
			(0.293)	(0.225)	(0.304)	(0.227)
County population density			-0.035*	-0.018**	-0.036*	-0.018*
			(0.020)	(0.009)	(0.021)	(0.009)
Mid Atlantic			0.202	-0.674***	0.211	-0.674***
			(0.217)	(0.165)	(0.227)	(0.166)
East North Central			0.092	-0.553***	0.089	-0.553***
			(0.221)	(0.178)	(0.230)	(0.182)
West North Central			0.442	-0.482**	0.445	-0.482**
			(0.289)	(0.239)	(0.298)	(0.246)
South			-0.555	-0.697***	-0.580	-0.697***
			(0.353)	(0.248)	(0.372)	(0.249)
λ for years:						
[1,2)	1.610***	1.000***	1.833***	1.191***	1.965***	1.191***
	(0.363)	(0.106)	(0.456)	(0.137)	(0.534)	(0.137)
[2,3)	1.494***	0.764***	2.129***	1.026***	2.386***	1.026***
	(0.354)	(0.105)	(0.553)	(0.158)	(0.764)	(0.159)
[3,4)	1.423***	0.465***	2.143***	0.741***	2.461***	0.741***
	(0.350)	(0.090)	(0.587)	(0.160)	(0.880)	(0.160)
[4,5)	1.127***	0.292***	1.872***	0.590***	2.181**	0.590***
	(0.296)	(0.076)	(0.551)	(0.170)	(0.847)	(0.170)
[5,6)	1.160***	0.167***	2.052***	0.301**	2.413**	0.301**
	(0.309)	(0.060)	(0.623)	(0.129)	(0.973)	(0.130)
[6,7)	0.668***	0.251***	1.329***	0.517***	1.572**	0.517***
	(0.210)	(0.081)	(0.462)	(0.196)	(0.697)	(0.196)
[7,8)	0.626***	0.117**	1.188***	0.328*	1.409**	0.327*
	(0.205)	(0.059)	(0.437)	(0.170)	(0.651)	(0.170)
[8,∞)	0.078***	0.249***	0.169***	0.601***	0.201**	0.601***
	(0.020)	(0.041)	(0.050)	(0.131)	(0.084)	(0.132)
v_{low} *(constant in columns 1-2)*	-2.543***	-0.473***	-1.729	3.316	-2.066	3.317
	(0.184)	(0.071)	(2.785)	(2.181)	(3.464)	(2.281)
v_{high}					-0.347	-
					(3.289)	-
π_1					0.968***	
					(0.057)	
π_2					0.032	
					(0.045)	
Log Likelihood	-2220.771		-1854.559		-1855.270	
Observations	770		688		688	

Hazard coefficients are reported. Sample: women widowed before 1880 who applied for a pension within five years of husband's death. Column (3) includes a correction for correlated unobserved heterogeneity, and does not include a constant as this is not identified separately from one of the mass points in the distribution of the unobserved heterogeneity terms; columns (1) and (2) make no such adjustment, and include a constant. Age at widowhood and all widows' pension variables (including county of residence) are taken from the pension file data collected by the author. First husband characteristics come from the UA data and are based on enlistment variables; occupational wages measured using 1900 occupational wage distribution assigned to 1950 occupational codes, with an imputed wage for farmers (Preston and Haines 1992; Abramitzky Boustan and Eriksson 2012; Olivetti and Paserman 2013; Salisbury 2014). County-level variables are taken at the time of pension application; they are the weighted average of these variables at the decadal censuses preceding and following the date of pension application (Haines and ICPSR 2010). On the time interval [0,1), the hazard rate of both risks is normalized to one (this is necessary because I include a constant in the model). The variables v_{low} and v_{high} are the two mass points in the distributions of v_m and v_p. Note that both mass points in the distributino of v_p converged to the same value. The variables π1-π2 are the estimated probability of each unobserved heterogeneity event.

Table 3: Effect of Pension on Hazard Rate of Remarriage: Interaction Effects

Interaction Variable:	Age	Number of Children	Male to Female Ratio	Population Density	Soldier 1860 Wealth
			Simple Model		
Pension	-0.545***	-0.299*	-0.273*	-0.282*	-0.239
	(0.170)	(0.155)	(0.158)	(0.158)	(0.201)
Pension X Variable	-0.061***	-0.166**	0.153	-0.012	0.122
	(0.017)	(0.067)	(1.621)	(0.035)	(0.089)
Log likelihood	-1847.944	-1851.558	-1854.555	-1854.508	-1247.970
Observations	688	688	688	688	457
			Full Model		
Pension	-0.566***	-0.307*	-0.287*	-0.293*	-0.240
	(0.176)	(0.160)	(0.165)	(0.163)	(0.208)
Pension X Variable	-0.063***	-0.170**	0.293	-0.013	0.124
	(0.017)	(0.069)	(1.765)	(0.036)	(0.092)
Log likelihood	-1848.673	-1852.214	-1855.345	-1855.280	-1248.348
Observations	688	688	688	688	457

Interaction Variable:	New England	Mid Atlantic	East North Central	West North Central	South
			Simple Model		
Pension	-0.231	-0.373**	-0.077	-0.291*	-0.330**
	(0.157)	(0.173)	(0.201)	(0.160)	(0.156)
Pension X Variable	-0.662	0.355	-0.385	0.205	1.423*
	(0.440)	(0.279)	(0.248)	(0.404)	(0.788)
Log likelihood	-1853.528	-1853.726	-1853.354	-1854.429	-1852.461
Observations	688	688	688	688	688
			Full Model		
Pension	-0.236	-0.369**	-0.079	-0.292*	-0.324**
	(0.161)	(0.175)	(0.203)	(0.165)	(0.162)
Pension X Variable	-0.685	0.347	-0.379	0.217	1.510*
	(0.448)	(0.275)	(0.253)	(0.418)	(0.839)
Log likelihood	-1854.258	-1854.373	-1854.077	-1855.227	-1853.104
Observations	688	688	688	688	688

All specifications include the full set of controls from table 2; see notes to this table for explanation. The full model includes a correction for correlated unobserved heterogeneity; the simplel model does not. Soldier's 1860 wealth is derived from links to the 1860 census in the CPE database. Only observations that have been successfully linked to this census are included in this specification.

Table 4: OLS and 2SLS Estimates of Relationship between Pension and Remarriage

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)
	Dependent variable: married w/in time frame (OLS)					*Dependent variable: married w/in time frame (2SLS)*				
Time Frame	1 year	2 years	3 years	4 years	5 years	1 year	2 years	3 years	4 years	5 years
Pension granted w/in time frame	-0.030	-0.083**	-0.194***	-0.267***	-0.231***	-0.512	-0.901	-0.664	-0.775*	-1.046**
	(0.024)	(0.037)	(0.046)	(0.050)	(0.052)	(0.653)	(0.670)	(0.426)	(0.432)	(0.452)
Age at widowhood	-0.002	-0.011***	-0.016***	-0.020***	-0.023***	-0.006	-0.011***	-0.015***	-0.021***	-0.021***
	(0.002)	(0.003)	(0.003)	(0.004)	(0.004)	(0.006)	(0.004)	(0.004)	(0.004)	(0.005)
Number of children	-0.004	-0.005	-0.010	-0.012	-0.014	-0.002	-0.009	-0.013	-0.017	-0.025
	(0.007)	(0.010)	(0.011)	(0.012)	(0.012)	(0.009)	(0.014)	(0.013)	(0.013)	(0.015)
Year of widowhood	0.001	-0.002	-0.011**	-0.014**	-0.014**	-0.012	-0.030	-0.030*	-0.035**	-0.046**
	(0.003)	(0.005)	(0.005)	(0.005)	(0.005)	(0.019)	(0.024)	(0.017)	(0.017)	(0.018)
Time to pension application	-0.014	-0.016	-0.008	-0.017	-0.024	-0.056	-0.105	-0.054	-0.056	-0.094*
	(0.014)	(0.020)	(0.023)	(0.023)	(0.024)	(0.059)	(0.076)	(0.052)	(0.045)	(0.050)
Potential minor pension	-0.000	-0.000	-0.000	-0.000	-0.000	-0.000	0.000	0.000	-0.000	0.000
	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)
No pension attorney	0.050	0.028	0.004	-0.013	0.027	0.112	0.080	-0.000	-0.027	0.006
	(0.040)	(0.059)	(0.065)	(0.067)	(0.069)	(0.096)	(0.090)	(0.071)	(0.073)	(0.084)
Washignton pension attorney	0.021	0.025	-0.004	-0.001	-0.009	-0.042	-0.072	-0.061	-0.055	-0.065
	(0.031)	(0.045)	(0.051)	(0.053)	(0.054)	(0.090)	(0.097)	(0.065)	(0.062)	(0.067)
First husband: height	-0.052	-0.032	-0.079	-0.134	-0.137	-0.087	-0.034	-0.090	-0.153	-0.190*
	(0.052)	(0.077)	(0.088)	(0.093)	(0.095)	(0.085)	(0.103)	(0.094)	(0.099)	(0.115)
First husband: log occupational wage	0.055	0.014	0.019	-0.142	-0.023	0.063	0.144	0.086	-0.155	-0.061
	(0.062)	(0.091)	(0.103)	(0.107)	(0.109)	(0.080)	(0.161)	(0.123)	(0.115)	(0.130)
First husband: age at death	0.000	0.003	0.004	0.005	0.006*	0.003	0.005	0.005	0.008*	0.009**
	(0.002)	(0.003)	(0.004)	(0.004)	(0.004)	(0.005)	(0.005)	(0.004)	(0.004)	(0.005)
County male-to-female ratio	0.715***	0.387	0.490	0.528	0.979***	0.889***	0.410	0.516	0.495	0.999**
	(0.218)	(0.317)	(0.353)	(0.362)	(0.369)	(0.336)	(0.432)	(0.383)	(0.392)	(0.440)
County percent urban	0.017	0.004	0.081	0.078	0.183*	0.066	0.117	0.113	0.062	0.163
	(0.055)	(0.083)	(0.092)	(0.096)	(0.098)	(0.099)	(0.147)	(0.103)	(0.102)	(0.116)
County population density	-0.000	-0.000	-0.000	-0.000	-0.000*	-0.000	-0.000	-0.000	-0.000	-0.000
	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)
Last name: mean occupational income	0.400**	0.300	0.194	0.427	0.554**	0.842*	1.082*	0.587	0.920**	1.181***
	(0.164)	(0.239)	(0.265)	(0.273)	(0.279)	(0.474)	(0.629)	(0.365)	(0.362)	(0.405)
Last mean: mean immigrant status	0.052	-0.040	-0.092	-0.105	-0.089	0.042	-0.088	-0.089	-0.059	-0.088
	(0.061)	(0.090)	(0.100)	(0.105)	(0.107)	(0.090)	(0.139)	(0.117)	(0.118)	(0.133)
Last name: mean literacy	-0.045	0.093	0.237	0.162	-0.093	-0.164	-0.084	0.058	-0.021	-0.072
	(0.131)	(0.191)	(0.214)	(0.219)	(0.223)	(0.232)	(0.342)	(0.301)	(0.306)	(0.345)
Constant	-5.095	1.185	19.336**	25.017**	23.698**	16.683	49.311	51.937*	61.974*	81.489**
	(5.837)	(8.629)	(9.715)	(10.191)	(10.418)	(33.045)	(42.328)	(30.131)	(32.172)	(34.018)
Observations	597	585	574	561	557	586	574	564	551	547
First stage F statistic						1.281	3.159	7.348	8.335	10.07
AR 90% Confidence Region for pension effect						(-∞, +∞)	(-11.67, -0.070)	(-1.88, 0 .007)	(-1.93, -0.096)	(-2.29, -0.407)
R-squared	0.056	0.078	0.146	0.200	0.221					

Instrument used in 2SLS specification is name homogeneity index. "Last name" variables are means by phonetic name group among household heads in IPUMS 1 percent sample from 1860-1880. AR 90% confidence region for pension effect is 90 percent confidence interval for the effect of the pension based on the Anderson-Rubin statistic, which is robust to weak instruments. See notes to table 3 for description of sample and other variables.

Table 5: Sensitivity of Estimates to Sample Restrictions

Model:	Simple			Full		
			Panel A. Baseline			
Pension	-0.269*	-0.545***	-0.299*	-0.283*	-0.566***	-0.307*
	(0.154)	(0.170)	(0.155)	(0.159)	(0.176)	(0.160)
Pension X Age at widowhood		-0.061***			-0.063***	
		(0.017)			(0.017)	
Pension X Number of children			-0.166**			-0.170**
			(0.067)			(0.069)
Log-Likelihood	-1854.559	-1847.944	-1851.558	-1855.270	-1848.673	-1852.214
Observations	688	688	688	688	688	688
			Panel B. Women with Children under 16			
Effect of Pension on Marriage Rate	-0.225	-0.483***	-0.290*	-0.223	-0.491***	-0.288*
	(0.166)	(0.181)	(0.168)	(0.171)	(0.185)	(0.173)
Pension X Age at widowhood		-0.061***			-0.063***	
		(0.018)			(0.019)	
Pension X Number of children			-0.223***			-0.231***
			(0.075)			(0.078)
Log-Likelihood	-1645.245	-1639.811	-1640.848	-1645.783	-1640.343	-1641.452
Observations	583	583	583	583	583	583
			Panel C. Women who live past 1901			
Effect of Pension on Marriage Rate	-0.295	-0.503	-0.421	-0.303	-0.520	-0.434
	(0.276)	(0.319)	(0.284)	(0.279)	(0.320)	(0.289)
Pension X Age at widowhood		-0.039			-0.040	
		(0.032)			(0.032)	
Pension X Number of children			-0.256**			-0.260**
			(0.129)			(0.131)
Log-Likelihood	-792.605	-791.894	-790.646	-792.818	-791.960	-790.830
Observations	253	253	253	253	253	253
			Panel D. Linked Only			
Effect of Pension on Marriage Rate	-0.309*	-0.567***	-0.398**	-0.296	-0.577***	-0.389**
	(0.180)	(0.206)	(0.184)	(0.182)	(0.209)	(0.182)
Pension X Age at widowhood		-0.048**			-0.049**	
		(0.020)			(0.020)	
Pension X Number of children			-0.170**			-0.183**
			(0.081)			(0.080)
Log-Likelihood	-1378.258	-1375.465	-1376.082	-1378.330	-1375.857	-1375.396
Observations	464	464	464	464	464	464
			Panel E. Linked Only: immigrant status control			
Effect of Pension on Marriage Rate	-0.331*	-0.579***	-0.411**	-0.312*	-0.581***	-0.411**
	(0.181)	(0.207)	(0.184)	(0.181)	(0.206)	(0.187)
Pension X Age at widowhood		-0.046**			-0.047**	
		(0.020)			(0.020)	
Pension X Number of children			-0.161**			-0.175*
			(0.081)			(0.091)
Log-Likelihood	-1375.578	-1372.947	-1373.639	-1374.753	-1373.372	-1372.712
Observations	464	464	464	464	464	464
			Panel F. Husband died during war			
Effect of Pension on Marriage Rate	-0.246	-0.448**	-0.263	-0.249	-0.460**	-0.263
	(0.181)	(0.189)	(0.181)	(0.188)	(0.194)	(0.185)
Pension X Age at widowhood		-0.059***			-0.062***	
		(0.020)			(0.020)	
Pension X Number of children			-0.116			-0.119
			(0.081)			(0.082)
Log-Likelihood	-1338.634	-1334.480	-1337.623	-1339.148	-1334.662	-1338.131
Observations	502	502	502	502	502	502

All specifications include the full set of controls from table 3; see notes to this table for explanation. The top panel replicates the baseline results. Panel B restricts the sample to women who have children under 16 at the time of initial pension application. Panel C restricts the sample to women who live at least to 1901. Panel D restricts the sample to women whose marital status is verified independently by links to the census. Panel E includes the sample from panel D, and includes a control for immigrant status. Panel F restricts the sample to women widowed during the war. The first three columns include no correction for correlated unobserved heterogeneity, and the last three colums include this correction.

Figure 1: Possible Outcomes for Widows in Sample

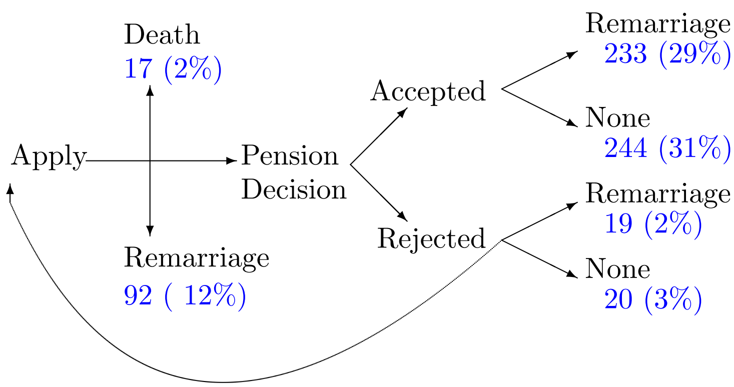
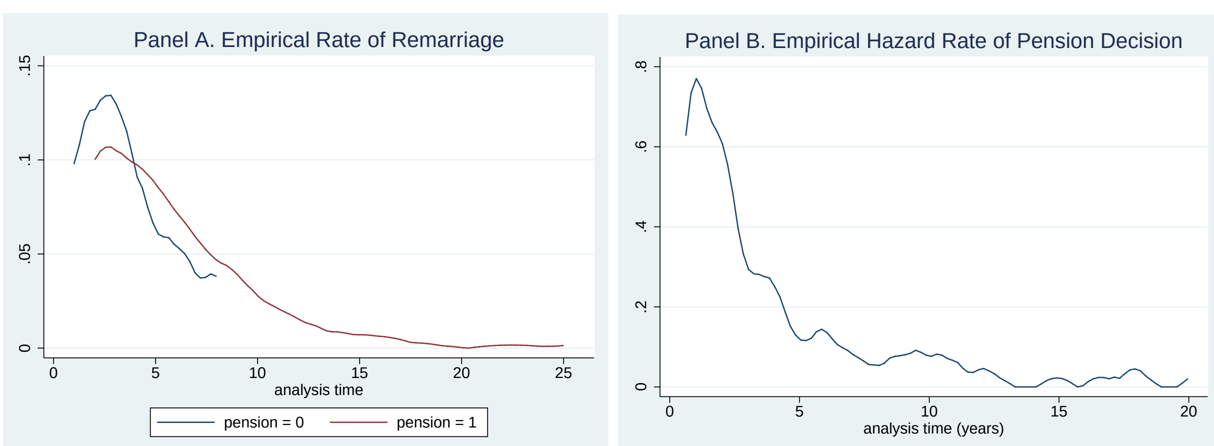


Figure 2: Empirical Hazard Rate of Remarriage and Pension Decision



A Theory Appendix: A Search Model of Marriage and Pensions

Suppose there are three otherwise identical types of widows: those who are receiving a pension (indexed by P), those who never receive a pension (N), and those who have pending claims (denoted with tildes). Married women are indexed by M. Assume for simplicity that there is no divorce. A marriage generates flow utility θ, which is drawn from a distribution $F(\theta)$, and discounting occurs at a rate r. Each state, married or single, is associated with a lifetime expected value, V. For all women, the value of being in a marriage with match quality θ is given by:

$$rV^M = \theta$$

In words, this is the present discounted value of receiving utility θ forever. The value of being single is different for pensioned and unpensioned women. Suppose remaining single generates a flow utility s, and women with pensions receive additional utility p. Marriage proposals have a poisson arrival rate α, which depends on search effort. Specifically, it costs a widow $c(\alpha)$ in utility to obtain a rate of proposals α. I assume that costs are increasing and convex in α, so $c'(\alpha) > 0$ and $c''(\alpha) > 0$. Then, the value to a pensioned woman of remaining single with proposal rate α_P^* can be written

$$rV^P = s + p - c(\alpha_P^*) + \alpha_P^* E[\max\{V^M - V^P, 0\}] \tag{3}$$

This is composed of two elements: the instantaneous utility a woman receives ($s+p-c(\alpha_P^*)$) and a term that reflects additional value, over and above the value of remaining single, from anticipated future proposals of marriage. It is a standard result that these unmarried women will have a reservation match quality, θ_P, which means they will accept any match carrying quality $\theta \geq \theta_P$. This has the property that $V^M(\theta_P) = V^P = \theta_P/r$. In other words, the reservation match quality is such that the woman is indifferent between remaining single and accepting the match. Substituting this into (1), and re-writing the expectation as an integral, we get the following equation that implicitly defines this reservation match quality:

$$\theta_P = s + p - c(\alpha_P^*) + \frac{\alpha_P^*}{r} \int_{\theta_P}^{\infty} (\theta - \theta_P) dF(\theta)$$

Women will choose α_P^* that maximizes the value of being unmarried. The maximizing level α_P^* will solve the following first order condition (Mortensen 1986):

$$rc'(\alpha_P^*) = \int_{\theta_P}^{\infty} (\theta - \theta_P) dF(\theta)$$

Similarly, for women who do not receive pensions, the reservation match quality is

$$\theta_N = s - c(\alpha_N^*) + \frac{\alpha_N^*}{r} \int_{\theta_N}^{\infty} (\theta - \theta_N) dF(\theta)$$

It is straightforward to show that θ_P is increasing and α_P^* is decreasing in p (Rogerson et al 2005); therefore, $\theta_P > \theta_N$ and $\alpha_P^* < \alpha_N^*$. In other words, women with pensions should be more selective and should spend less effort on search in the marriage market.

The above results are a straightforward application of search theory to this particular problem (Rogerson et al 2005). I now derive the value of being unmarried for women with pending pension claims. Suppose that the (endogenous) arrival rate of marriage proposals for a woman with a pending claim is $\tilde{\alpha}^*$, and the arrival rate of pension decisions is λ. The probability that the decision will be favorable is π. Then, the value of being a widow with a pending pension claim ($\tilde{V}$) can be written:

$$r\tilde{V} = s - c(\tilde{\alpha}^*) + \tilde{\alpha}^* \Big(E[\max(V^M - \tilde{V}, 0)] \Big) + \lambda \Big(\pi V^P + (1-\pi) V^N - \tilde{V} \Big) \tag{4}$$

Proof. This follows Rogerson et al (2005). Suppose the arrival rate of pension decisions is λ, the arrival rate of marriage proposals is α, and the probability of an acceptance is π. Take Δ to be an arbitrarily small period of time, and note that, for search effort $c(\alpha)$, the probability of receiving a marriage proposal during this interval is $\alpha\Delta$; similarly, the probability of receiving a decision from the pension bureau is $\lambda\Delta$. Call V^S the expected value of being single, which will be a weighted average of the value of being single in each potential state of “singlehood”. Then, it must be that

$$\begin{aligned}
\tilde{V} &= \Delta(s-c(\alpha)) + \frac{\Delta\alpha}{1+\Delta r}\Big(E[\max(V^M, V^S)]\Big) + \frac{1-\Delta\alpha}{1+\Delta r}E[V^S] \\
&= \Delta(s-c(\alpha)) + \frac{\Delta\alpha}{1+\Delta r}\left(\Delta\lambda\Big(\pi E[\max(V^M, V^P)] + (1-\pi)E[\max(V^M, V^N)]\Big) + (1-\Delta\lambda)E[\max(V^M, \tilde{V})]\right)+ \\
&\quad + \frac{1-\Delta\alpha}{1+\Delta r}\left(\Delta\lambda\Big(\pi V^P + (1-\pi)V^N\Big) + (1-\Delta\lambda)\tilde{V}\right) \\
&= \Delta(s-c(\alpha)) + \frac{\Delta\alpha}{1+\Delta r}\left(\Delta\lambda\Big(\pi E[\max(V^M - V^P, 0)] + (1-\pi)E[\max(V^M - V^N, 0)]\Big)+\right. \\
&\quad \left. + (1-\Delta\lambda)E[\max(V^M - \tilde{V}, 0)]\right) + \frac{\Delta\lambda}{1+\Delta r}\Big(\pi V^M + (1-\pi)V^N - \tilde{V}\Big) + \frac{1}{1+\Delta r}\tilde{V}
\end{aligned}$$

Re-arranging, dividing by Δ, and taking the limit as $\Delta \to 0$, we get (4). □

Because V^M is strictly increasing in θ, the right hand side of this equation is also strictly increasing in θ. This implies that there exists a reservation match quality $\tilde{\theta}$ for women with pending pension applications:

$$\tilde{\theta} = s - c(\tilde{\alpha}^*) + \frac{\tilde{\alpha}^*}{r}\int_{\tilde{\theta}}^{\infty}(\theta - \tilde{\theta})dF(\theta) + \frac{\lambda}{r}\Big(\pi\theta_P + (1-\pi)\theta_N - \tilde{\theta}\Big) \tag{5}$$

The optimal $\tilde{\alpha}^*$ will be defined similarly to those of the other two groups.

Proposition. For $\pi \in [0,1]$, $\tilde{\theta} < \theta_P$ and $\tilde{\alpha}^* > \alpha_P^*$.

Proof. Throughout, I use the well known result that $\int_{\theta_i}^{\infty}(\theta - \theta_i)dF(\theta) = \int_{\theta_i}^{\infty}(1 - F(\theta))d(\theta)$ First notice that $\tilde{\theta}$ is strictly increasing in π:

$$\begin{aligned}
\frac{\partial\tilde{\theta}}{\partial\pi} &= -\frac{\tilde{\alpha}^*}{r}(1 - F(\tilde{\theta}))\frac{\partial\tilde{\theta}}{\partial\pi} + \frac{\lambda}{r}(\theta_P - \theta_N) \Rightarrow \\
\frac{\partial\tilde{\theta}}{\partial\pi} &= \frac{\lambda(\theta_P - \theta_N)}{r + \tilde{\alpha}^*(1 - F(\tilde{\theta}))} > 0
\end{aligned}$$

Now, define $\tilde{\theta}^1 = \tilde{\theta}$ when $\pi = 1$. Because $\tilde{\theta}$ is strictly increasing in π, if $\theta_P > \tilde{\theta}^1$, then $\theta_P > \tilde{\theta}$ for every $\pi \leq 1$. When $\pi = 1$:

$$\tilde{\theta} = s - c(\tilde{\alpha}^*) + \frac{\tilde{\alpha}^*}{r}\int_{\tilde{\theta}}(1 - F(\theta))d(\theta) + \frac{\lambda}{r}(\theta_P - \tilde{\theta})$$

Suppose $\tilde{\theta} \geq \theta_P$. Because the optimal α^* is decreasing in reservation θ (see below), it follows that

$\alpha_P^* \geq \tilde{\alpha}^*$. Two inequalities follow from this: First,

$$\frac{1}{r}\int_{\tilde{\theta}}(1-F(\theta))d(\theta) \leq \frac{1}{r}\int_{\theta_P}(1-F(\theta))d(\theta)$$

And, from convexity of $c(\alpha)$, we get the following inequality:

$$-c(\tilde{\alpha}^*) \leq -c(\alpha_P^*) + c'(\alpha_P)(\alpha_P^* - \tilde{\alpha}^*)$$

This implies the following:

$$\begin{aligned}
\tilde{\theta} &= s - c(\tilde{\alpha}^*) + \frac{\tilde{\alpha}^*}{r}\int_{\tilde{\theta}}(1-F(\theta))d(\theta) + \frac{\lambda}{r}(\theta_P - \tilde{\theta}) \\
&\leq s - c(\tilde{\alpha}^*) + \frac{\tilde{\alpha}^*}{r}\int_{\theta_P}(1-F(\theta))d(\theta) \\
&\leq s - c(\alpha_P^*) + c'(\alpha_P^*)(\alpha_P^* - \tilde{\alpha}^*) + \frac{\tilde{\alpha}^*}{r}\int_{\theta_P}(1-F(\theta))d(\theta) \\
&= s - c(\alpha_P^*) + \frac{1}{r}\int_{\theta_P}(1-F(\theta))d\theta(\alpha_P^* - \tilde{\alpha}^*) + \frac{\tilde{\alpha}^*}{r}\int_{\theta_P}(1-F(\theta))d(\theta) \\
&= s - c(\alpha_P^*) + \frac{\alpha_P^*}{r}\int_{\theta_P}(1-F(\theta))d(\theta) \\
&= \theta_P - p < \theta_P
\end{aligned}$$

This is a contradiction. So, it must be that, when $\pi = 1$, $\theta_P > \tilde{\theta}$, which further implies that $\theta_P > \tilde{\theta}$ for all $\pi \leq 1$.

The result that $\alpha_P^* < \tilde{\alpha}^*$ follows from the fact that α^* is decreasing in reservation match quality. Recall that, for reservation match quality θ_i, α^* is defined by the following condition:

$$rc'(\alpha^*) = \int_{\theta_i}^{\infty}(1-F(\theta))d(\theta)$$

Then, $\partial\alpha^*/\partial\theta_i$ is given by:

$$\frac{\partial\alpha^*}{\partial\theta_i} = \frac{-(1-F(\theta_i))}{rc''(\alpha^*)} < 0$$

This follows from the convexity of search costs. □

It is a well known result that lower reservation match qualities and greater search effort cause the hazard rate of remarriage to be greater. So, this model predicts that women with pending pension claims should marry at a faster rate than women with claims in hand.

B Data Appendix

B.1 Detailed Data Description

The sample of widows is drawn from Union Army (UA) database created by the Center for Population Economics (CPE) at the University of Chicago (Fogel et al 2000). The data are drawn from three principal sources: the military, pension and medical records are compiled from sources at the National Archives including military service records and Civil War pension records; data from the Surgeons Certificates contain detailed information about veterans' health status, which was used to determine pension eligibility; further socioeconomic information is gathered by linking veterans to the Federal Censuses of 1850, 1860, 1900 and 1910. These data have primarily been used to study health and aging in the late 19th and early 20th centuries. See for example Costa 1997, 1995, 1993; Fogel 2004; Eli 2010. They have also been used to analyze group dynamics in military settings (Costa and Kahn 2003, 2008). The data contain information about every soldier who enlisted in 303 randomly sampled companies of white volunteer infantry regiments. The database contains 39,341 observations and 3,230 variables (Fogel et al. 2000).

Information on widows' pension and marital outcomes are compiled from pension records at the National Archives in Washington, DC. Using the indices to the Civil War pension files available on ancestry.com and fold3.com, I compile a list of all pension applications made and certificates issued on behalf of soldiers married to the women in my sample. Then, I request these files from the National Archives. In approximately 93 percent of cases, these files are successfully located, and I am able to collect digital images of them. Files that could not be located had either been taken out by another user (30% of cases), or the file number was incorrectly recorded, and the record puller was unable to find it (70% of cases). Where possible, I make use of digital images of widows' pensions from the website fold3.com. This website is in the process of uploading images of accepted widows' pensions, which they are doing chronologically. It is not possible to make exclusive use of this resource for several reasons. First, this project is expected to take several years to complete. Second, they do not include rejected pension applications. In total, 33 percent of my sample can be collected from this resource.

Because of the importance of these variables to the paper, I describe the source of information on pension outcomes and marriages in the body of the text. However, there are other important variables collected from the pension files. Other available information includes the widow's age and place of residence, as she had to furnish this information in her pension application. If a remarried widow applied to be restored to the pension rolls under the act of March 3, 1901, her file will contain further information about her second husband. For example, she had to provide proof of her husband's death, which usually meant furnishing a death certificate. In some cases, these death certificates contain the age, birthplace, and occupation of the husband.

B.2 Variables

Variable	**Source**	**Notes**
Date of first husband's death	Union Army database (Fogel et al 2000)	Based on dependents' pension applications or military death records
Date of pension application	Widows' pension database (Salisbury)	Date at which widow filled out pension declaration form; if missing, date at which pension application received by pension bureau
Date of pension receipt	Widows' pension database	Date of issuance on pension certificate; if missing, date of pension approval on pension brief
Date of remarriage	Widows' pension database	Based on marriage certificates or affadavits rendered in support of minors' pension application or application for widow to be restored to the pension rolls under a later act.
Date of death	Widows' pension database	Based on pension drop cards, or death records filed in support of minors' pension application.
Age at widowhood	Widows' pension database	Deduced from widow's first pension declaration, in which age and date of application are both provided.
Number of children	Union Army database	Equal to number of children under the age of 16 when widow first filed for pension.
Potential minor pension	Union Army database	Calculated as $8/mo until youngest child turns 16, or $8/mo plus $2/mo for each child under 16 if widowed after July 25, 1866.
No pension attorney	Widows' pension database	Equal to one if the widow did not hire an attorney at the time of filing her first claim
Washignton pension attorney	Widows' pension database	Equal to one if the widow first hired an attorney from a Washington firm at the time of filing her first claim
First husband: height	Union Army database	Soldier's height at enlistment
First husband: log occupational wage	Union Army database; Preston and Haines (1991); United States Census of Agriculture (1900)	Based on soldier's occupation at enlistment
First husband: age at death	Union Army database	Based on implied birth year from age at enlistment
County of residence	Widows' pension database	County listed on first pension application form
County male-to-female ratio	Haines and ICPSR (2010)	Weighted mean of male-to-female ratio in 1860, 1870 and/or 1880, depending on date of application.
County percent urban	Haines and ICPSR (2010)	See above.
County population density	Haines and ICPSR (2010)	See ablve.
Name homogeneity index	Ruggles et al (2010); Atack and Bateman (1992)	Herfindahl index of concentration of unique spellings within phonetic surname groups among household heads in 1 percent IPUMS sample from 1860-1880. Phonetic groups created using NYIIS algorithm.

Last name: mean occupational income	Ruggles et al (2010); Preston and Haines (1991); United States Census of Agriculture (1900)	Mean occupation status of household head, calculated using 1900 wage distribution, by phonetic name group in IPUMS 1 percent sample from 1860-1880.
Last mean: mean immigrant status	Ruggles et al (2010)	Mean literacy of household head by phonetic name group in IPUMS 1 percent sample from 1860-1880.
Last name: mean literacy	Ruggles et al (2010)	Mean immigrant status of household head by phonetic name group in IPUMS 1 percent sample from 1860-1880.
Immigrant statuss	Linked widow sample; ancestry.com	Immigrant in census of 1870 or 1880