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Common law marriage and couple formation

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Abstract

The Current Population Survey is used to investigate effects of Common Law Marriage (CLM) on whether young US-born adults live in couples in the US CLM effects are identified through cross-state and time variation, as some states repealed CLM over the period examined. Analysis based on Gary Becker's marriage economics helps explain why CLM affects couple formation and does so differently depending on education, sex ratios, age, and parent status. CLM reduces in-couple residence, and more so for childless adults and where there are fewer men per woman. Effects are larger for college-educated men and women without college. CLM effects on likelihood of marriage and cohabitation and likelihood of being divorced if ever-married are also estimated.

JEL: J10; J12; J16

Keywords: Common-law marriage; Couple; Couple formation; Marriage; Cohabitation; Gary Becker

1 Introduction

The last few decades have witnessed a movement from marriage towards unmarried cohabitation. For example, 48% of over 12,000 women interviewed in 2006–2010 for the National Survey of Family Growth cohabited with a partner at a first union, compared with 34% of women in 1995.[1] About 41% of all children in this country are born to unmarried mothers, including a quarter of all births to mothers who cohabit with their partner (National Vital Statistics Reports 2014). In our sample, couple formation rates have decreased during this period: in 1995, 43% of men ages 18 to 35 and 50% of women the same age resided in couple; by 2011, these percentages had dropped to 37% and 44%.

There are many benefits of living in a couple rather than alone: larger households enjoy economies of scale in expenses; partners can specialize in various tasks which leads to higher productivity of both partners; household tasks require no transaction costs, partners can enjoy joint consumption; and one partner's income can serve as insurance against the other partner's income shocks. However the main reason why couples are formed is to provide a protective environment for children. Children born to parents living in a couple are better off than children of single parents, in part because they grow up with more of a father's presence in their lives (see Mincy and Oliver 2003). Relative to children raised by a mother and a father, children raised by a single parent often achieve less in terms of school performance (McLanahan and Sandefur 1994, McLanahan and Sigle-Rushton 2004) and have higher rates of depression and crime participation (Hobcraft 1998, Sigle-Rushton et al. 2005). In addition, whether individuals form couples or not has an impact on the demand for goods and services such as housing and childcare.

The choice between marriage and (non-marital) cohabitation [hence simply called 'cohabitation'] is a more subtle one. The two institutions differ, e.g., in their tax implications, access to a spouse's Social Security and other government benefits, and access to employer benefits such as health insurance and pension plans. Consequently, marriage provides more asset protection to the spouse who earned less (possibly because they engaged more in household production) in the event of dissolution due to separation or death.

In this paper, we examine whether the likelihood of in-couple residence, marriage, cohabitation, and likelihood of being divorced among those ever-married are associated with variation in the state-level availability of Common Law Marriage (CLM). Several US states offer their heterosexual residents this additional way of organizing their living-together arrangements. CLM does not require a marriage certificate or ceremony, it can be established when couples cohabit and hold themselves out as spouses by calling each other husband and wife in public, using the same last name, filing joint tax returns, or declaring their marriage on applications, leases, birth certificates and other documents. Cohabiting couples who have a child are almost certainly considered "married" in a CLM state. In the event of separation, such couples go through a regular divorce. Once established, CLM is no different from marriage, including its acceptance by all other states and government institutions dealing with tax collection and redistribution of income.

In most states with CLM, there are no rules regarding cohabitation time required for common law marriage. A short term cohabiting relationship may also be called "marriage" if both spouses agree. One peculiar feature of CLM that distinguishes it from a regular marriage is that it can be claimed *ex-post* by one partner after the relationship ends due to separation or death. This fluid definition of marriage in CLM states is one of the reasons why this paper's focus is on in-couple co-residence—married or not–more than on marriage and cohabitation as two separate states.

Most US states recognized CLM in the past but don't any more. The historical vagaries of why some states have the law and others don't are described in Bowman (1996) and Lind (2008). Informal marriage was originally brought to the colonies by British settlers, and it remained in the U.S. long after its repeal in the U.K. in 1753. Among the factors that explain early popularity of these laws in the US are the philosophy of non-interference of the state in its citizens' private affairs and the practical advantages of establishing marriages in frontier conditions when priests were not available. A first wave of repeals of CLM laws occurred between 1875 and 1917; a second wave in the years 1921 to 1959. These repeals have been explained, e.g., in terms of transportation improvements (it became easier to register marriages), the expanded system of government benefits with spousal entitlements (making it more costly for states to recognize CLM), and growth in the number of contested marriage cases (Bowman 1996).

As of 2014, common-law marriage remains legal in: Alabama, Colorado, Iowa, Kansas, Montana, New Hampshire (only posthumously for purposes of inheritance), Oklahoma, Rhode Island, South Carolina, Texas, as well as in the Navajo Nation and in the District of Columbia. CLM laws are also found in Utah where they were repealed in 1888 and reinstituted in 1987. Most recently, CLM was repealed by Ohio (1991), Idaho (1996), Georgia (1997), and Pennsylvania (2005), thus providing us with a quasi-experiment. The internet contains plenty of information and legal advice for couples about CLM, so we know it is practiced.[2] Some indication of CLM's prevalence can be derived from the fact that about one hundred legal CLM–related judgments have been issued each decade in each state with CLM laws at the federal level, according to legal historian Goran Lind (2008).

Data availability on CLM is a problem. There is virtually no official data on CLM marriages published by state or local governments, even though some counties encourage residents to register their CLMs.[3] We circumvent this data problem by exploiting cross-state variation in CLM and changes over time.

Our analysis includes the derivation of predictions regarding the effect of CLM on couple formation based on Gary Becker's (1973) demand and supply model of marriage.[4] Becker's pioneering economic theory of marriage is part of his economic approach to the family, one of his contributions highlighted by the Nobel prize committee. Our model leads us to predict that CLM will be associated with lower couple formation rates and that the extent of CLM's effect will depend on male and female education, sex ratios, parental status, and age. The model assumes traditional gender roles, with women more involved in home production than men, and helps explain gender differences in the effects of CLM on subgroups of the population.

Previous research has linked variation in legal regimes to outcomes related to couple formation, dissolution, and fertility. The focus of much of that U.S.-based research has been on the effect of no-fault or unilateral divorce on divorce rates (e.g. Peters 1986, Friedberg 1998, and Wolfers 2006), marriage rates (e.g. Alesina A and Guiliano P 2007, Divorce, Fertility and the Value of Marriage, Harvard Institue of Economics Research Discussion Paper No 2136, April, unpublished manuscript, Rasul 2006), and couples' investments in marriage-specific capital such as spouse's education and children (Stevenson 2007). Halla (2013) showed that the introduction of joint custody increased marriage rates, overall fertility (including a shift from non-marital to marital fertility), and divorce rates for older couples. Leturcq (2011) has studied how the introduction of PACS (civil unions) in France in 1999 has affected marriage, and Guttierez and Becerra (2012) studied its effect on fertility. Variations in laws regulating division of property in the case of dissolution have helped explain the likelihood that women are out-of-couple when they give birth (Ekert-Jaffe and Grossbard 2008). However, to the best of our knowledge, ours is the first study examining the effect of CLM laws on couple formation. Elsewhere (Grossbard S, Vernon V (forthcoming) Common law marriage, labor supply and time use: a partial explanation for gender convergence in labor supply. Res Labor Econ), we analyzed effects of CLM on labor supply.

We include sex ratios in all our models explaining in-couple residence, marriage, cohabitation, and likelihood of being divorced if ever-married. This variable was included in Becker's (1973) Demand and Supply models of marriage and in earlier empirical research on couple formation (e.g. by Heer and Grossbard-Shechtman 1981, Lichter et al. 1992, and Angrist 2002). Higher sex ratios tend to be associated with higher marriage rates relative to single status and relative to cohabitation (Grossbard-Shechtman 1993).

Our models assume that state laws regarding CLM and changes in those laws are known to state residents. The wealth of information online supports this assumption. We use individual-level data for US-born men and women ages 18-35 based on the Current Population Surveys (CPS) for the period 1995-2011, focusing on US-born individuals to increase the likelihood that respondents in our sample are familiar with this law, possibly based on the experience of older community members. We exclude foreign-born respondents who may have been unaware of CLM laws when they decided to form a couple. In some of our models, we assume that such legal knowledge is transmitted immediately. In other models, we relax that assumption and either exclude three years of data after the repeal of CLM in each of the relevant states or control for years before and after the repeal. We limit our analyses to respondents under age 36, as they are most likely to be entering marriage.[5] We also assume that inter-state migration is minimal, and that if people move, they don't do so based on whether a state has CLM or not.

Consistent with our predictions, we find that CLM is associated with lower rates of in-couple residence, especially when sex ratios are low. We also find similar CLM effects on marriage probabilities when sex ratios are low. CLM's negative effect on in-couple residence is seen for less educated men and women mainly because of its negative effect on cohabitation. Another finding consistent with our model's prediction is that CLM only has a negative effect on in-couple residence of men with a college education, and not on their female counterparts. This effect is found separately for the likelihood of marriage and that of cohabitation. We also find that CLM has a larger negative effect on in-couple residence of women with no college education than on that of women who are college-educated. Childless men and women are more likely to respond to changes in CLM. Our demand and supply model of marriage helps explain these findings.

2 Why would CLM affect couple formation?

The following predictions are based on a demand and supply model of marriage inspired by Becker (1973) and on the concept of *Work-In-Household (WiHo)*, defined as household production work of benefit to a spouse/partner. WiHo may include activities such as parenting, meal preparation, and house cleaning (see Grossbard-Shechtman 1984).[6] In a couple, individuals may sometimes exchange WiHo, but if one partner works relatively more in WiHo, that partner is the WiHo-worker. The WiHo-worker may not necessarily be working at WiHo full time: she or he may also be employed in the labor force. The other spouse may 'pay' the WiHo-worker in the form of an intra-couple transfer of money or goods.[7] A higher price for WiHo means that the individual WiHo-worker obtains a larger transfer for the same amount of WiHo[8].

Marriage markets are viewed as markets for WiHo, in part to facilitate reliance on labor market analysis.[9] An example of such a WiHo market is shown in Figure 1. h denotes WiHo; y denotes its price. When traditional gender roles and heterosexuality prevail, women tend to be the WiHo-workers. Women's willingness to form a couple is expressed as a supply of WiHo, i.e., a willingness to supply different amounts of WiHo at different 'prices'.[10] Women's supply is upward-sloping: higher pay gives incentives to work more at such household production. With traditional gender roles, those benefiting from WiHo and possibly 'paying' for it are men. Men thus have a demand for WiHo, reflecting their willingness to pay for different amounts of women's WiHo. The demand is downward-sloping: the more expensive WiHo, the more men will look for substitute ways of fulfilling their needs for clean clothing, meals, etc. Men will prefer to pay less for women's WiHo; women will prefer to earn more for that kind of work.

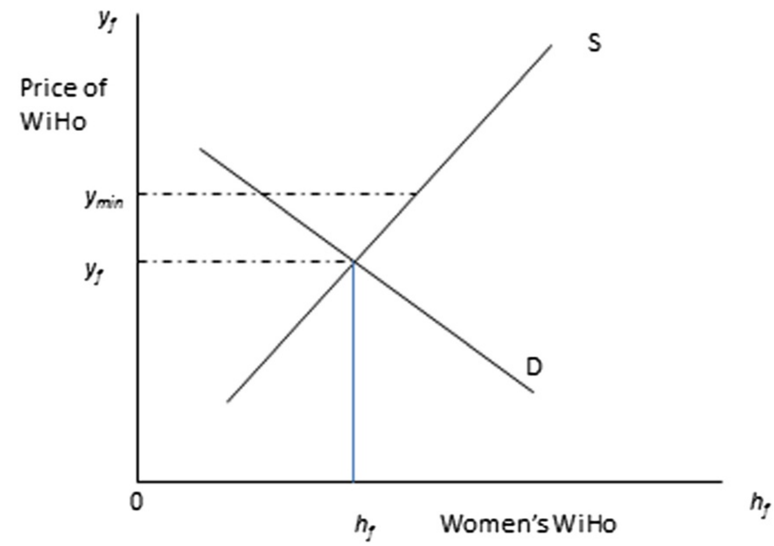

Figure 1 **Market for women's work-in-household, comparing CLM states to non-CLM states.**

These conflicting interests possibly lead couples to bargain over the price of WiHo. Equilibrium prices for WiHo are established at the intersection of aggregate demand and supply in markets for WiHo. In each market, women are sufficiently alike to be substitutable and the same is true for the men in that market. There are many interrelated WiHo markets defined by personal characteristics of men and women (such as education and age).

The price for WiHo when a couple cohabits (Work-in-a-cohabiting-household or cohabitation-WiHo) may be lower than that for WiHo supplied in marriage. This may be reflected, e.g., in the form of fewer benefits for WiHo-workers in case of death or divorce (Mincy et al. 2005). To the extent that a CLM law imposes the higher price for WiHo that prevails among married couples on cohabiting couples, it can be viewed as the equivalent of a minimum wage law. Interpreting the WiHo in Figure 1 as cohabitation-WiHo and assuming traditional gender roles, CLM laws set a minimum price of WiHo, i.e. *ymin.* In Figure 1, that minimum price lies above the market-clearing *y*. At this higher *y*, women are willing to work more at WiHo: they move up their supply. However, at the higher *ymin*, men are less willing to obtain WiHo than they were at the market-clearing price. The law prevents men from 'hiring' women doing WiHo work for them in return for a cohabitation contract with low legal protection. It has to be marriage whenever WiHo workers want it. The higher price of women's WiHo implies both a price effect and an income effect, leading men to move up their demand curve.

At the new equilibrium associated with CLM and *ymin*, the total amount of WiHo supplied will be the amount demanded and will be less than the amount of WiHo supplied in market equilibrium without CLM to the extent that the demand is downward-sloping and the supply is upward-sloping. By definition, less total employment in WiHo corresponds to a smaller number of cohabiting couples. The number of married couples is not expected to grow as a result of CLM; thus,

Prediction 1. *Under CLM, there will be a lower probability that individuals live in couples.*

As long as monogamy prevails, overall effects of CLM on couple formation of men and women are likely to be similar. However, in submarkets of men and women with particular characteristics, CLM may have different effects on the status of specific types of men and women. It is assumed that gender roles are traditional, and women get paid for their WiHo. Expected sensitivity to CLM laws will depend on how effective the *ymin* is relative to the market price of WiHo in a particular market. CLM laws are expected to have little impact if the market value of women's WiHo in equilibrium is already high (in such case the minimum *y* is not likely to be effective). A minimum *ymin* is more likely to be effective and therefore to cause a drop in couple formation where market-clearing WiHo prices *y* are relatively low.

Educated women interested in being (part-time or full-time) WiHo workers are likely to be more in demand in marriage markets than their counterparts without a college education. This higher demand is also reflected in higher marriage rates: for example, in 2012, only 73% of 35- to 39-year-old adults without a bachelor's degree had ever married, but 81% of their college-educated counterparts had (Fry 2014; this education gap also holds for women separately).[11] Higher demand for educated WiHo workers is also expected to raise the unobserved price of WiHo relative to that of spouses with low education. Therefore, to the extent that women do more WiHo than men, when a minimum price of *y* is imposed, it is expected to be more effective in markets for the WiHo of women with low education relative to markets for women with higher education whose *y* may already be above the *ymin*. Therefore, a minimum *y* due to CLM is likely to affect WiHo workers with low education more, and:

Prediction 2. *Drops in couple formation associated with CLM are likely to be larger for women without a college education than for women who are college-educated.*

If gender roles are traditional and men are paying for WiHo, and if they are more educated and their income is higher, they may find that by levying more potential claims on their future earnings, CLM entails higher costs: they may have to share a higher income in case of separation, for example, if CLM is available in their state, but not if they were cohabiting in a non-CLM state. Therefore, CLM entails more of a movement along men's demand curve for WiHo if men are college-educated than if they are not. Therefore,

Prediction 3. *Drops in couple formation associated with CLM are likely to be larger for men with a college education than for men without that education.*

Assuming traditional gender roles, it thus follows that the combined effect of CLM and education will differ for men and women. More generally, Predictions 2 and 3 regarding education depend on whether men and women are predominantly WiHo workers or employers of WiHo. The more they are likely to be on the employer side, the more the combined effect of education and CLM on couple formation is likely to be negative. The more they are likely to be WiHo workers, the less CLM is likely to affect couple formation of the college-educated.

Gender asymmetry in the predicted effect of CLM is related to possible gender differences in preferences for relationships with partners with a similar education. We postulate that given traditional gender roles, men tend to be on demand side and women on the supply side in WiHo markets. CLM effects are related to the likelihood that men pay women for their WiHo. Matches involving two college educated spouses who both participate full-time in the labor force are not expected to entail much of a payment for WiHo. Instead, partners may exchange WiHo, or they may rely on hired help. The effect of CLM on couple formation is more likely to be observed for couples trading WiHo in return for money or goods, which involves a monetary or quasi-monetary price *y*. This is more likely to occur when men are more educated and earn more than their spouse than when both have similar earnings.

Another group of women who are likely to obtain low market prices for their WiHo are women living in areas with low sex ratios. Low sex ratios mean that there are relatively few men for every woman wanting to marry, implying a relatively low demand for WiHo. Consequently, the price of women's WiHo is expected to be lower than when sex ratios are high. In cultures where brideprice and/or dowry are observed and varying with demand and supply, a lower sex ratio may involve a shift from brideprice (men pay) to dowry (women pay). In that vein, Francis (2011) found that when sex ratios dropped in Taiwan, brideprices became less common and dowries became more common. In the US, we only have implicit WiHo prices, but nevertheless, a minimum price for women's WiHo, *ymin,* is expected to have a more negative effect on couple formation prospects of women in markets with low sex ratios and therefore low price *y* relative to its effect on women in markets with high sex ratios and high *y*.

Prediction 4. *Drops in women's couple formation associated with CLM are likely to be larger in low sex ratio areas than in high sex ratio areas.*

The next distinction we make is between parents and childless adults. It is expected that relative to childless respondents, parents will do more WiHo work and are likely to get paid more for their WiHo by their partner or spouse. Given that their market-clearing *y* is likely to be high (even under cohabitation), women with children may not be affected by a minimum 'price' *ymin* implied by CLM. Therefore, a minimum *y* taking the form of CLM laws is likely to be more effective for childless women supplying WiHo to men than to women who share children with a man. It follows that

Prediction 5. *CLM is more likely to be associated with a lower likelihood of being in couple in the case of childless respondents than in the case of parents whose children are present.*

A final distinction is between respondents who are relatively younger and those who are older. Here we set the dividing line at 25 and distinguish those 18 to 25 and those 26 to 35. It is expected that CLM will have more effect on women who can expect a lower *y* in their markets for WiHo and who are therefore more likely to find that *ymin* is an effective minimum price. This could possibly apply more to women ages 18 to 25 than to women ages 26 to 35. As for CLM's effect on men of different ages, it will depend on age differences in men's demand for WiHo. Young men may be particularly sensitive to an increased price of cohabitation WiHo. It follows that

Prediction 6. *CLM is more likely to be associated with a lower likelihood of being in couple in the case of the youngest respondents than in the case of those somewhat older. This may be gender-specific.*

The variables that we selected to formulate these predictions may also affect other related outcomes, such as the likelihood that an individual is married if he/she resides in couple, cohabits if single, or is ever divorced at the time of the survey. CLM and some of these variables also help explain labor supply: the women who do have a 'WiHo job'—and therefore are observed to be in couple–are expected to get paid more for their WiHo if CLM is available than if it is not. Consequently, they will have a higher reservation wage and lower labor supply if CLM is available. Evidence supporting this prediction for a number of labor supply measures is reported in Grossbard and Vernon (forthcoming).

It also follows from this analysis that to the extent that in the US blacks and whites participate in separate WiHo markets in the case marriage and cohabitation, we could expect some differences in the effect of CLM on blacks and whites. However, we don't have a sufficient number of black respondents to test for such differential effects of CLM.

3 Data and sample means

We analyze micro data from the March Current Population Surveys[12] (CPS) for the period 1995-2011 to estimate individual probabilities of being in couple, being married, and cohabitation. This is a large nationally representative dataset with information on demographic characteristics, labor market status, and identifiable cohabiting relationships. A drawback of the CPS is that not all cohabiting couples can be identified prior to 2007: until that date only relationships between household heads and their partners were recorded, while other household members were assigned either married or single status. Therefore, our sample will underestimate the share of cohabiting couples in the population for 1995-2006. This should not be a problem because our variable of interest is not the time trend but the difference between CLM and non-CLM states, as long as the designation of a household head and the composition of other family members do not vary systematically by CLM status. CPS surveys prior to 1995 do not distinguish between partners and roommates; hence, we draw data starting in 1995.

Three states repealed CLM over the period covered by this data set: Idaho (1996), Georgia (1997), and Pennsylvania (2005). Had we been able to use earlier CPS data, we could also have examined the effect of repealing CLM in Florida (1968) and Ohio (1991).

We select all US-born men and women for we want to exclude individuals who possibly made their marriage decision in another country. Excluding non-US citizens resulted in a disproportionate loss of married individuals since first generation immigrants are more likely to be married and less likely to cohabit compared to the rest of the US population. This selection affected the Hispanic sample the most: it shrank by more than one-third.

We choose to focus on young individuals aged 18 to 35. Younger people are more likely to be affected by the change in the marriage law as they are more likely to transition in and out of marriage and cohabitation. We also drop same-sex cohabiting couples. Our sample includes 321,917 women and 292,376 men, of which around 21.6% live in CLM states. Sample means are presented in Table 1. It can be seen that CLM states have a higher proportion of married and a lower proportion of cohabiting residents. Respondents from CLM states are on average less educated, less likely to be enrolled in college, and are more likely to have children and be Hispanic. CLM states have lower unemployment rates, lower median household income, less generous welfare payments and slightly lower sex ratios. All differences by CLM status are statistically significant at 5% due to large samples.

Table 1 Sample means

	Women		Men	
	CLM 21.6%		CLM 21.4%	
	Non-CLM	**CLM**	**Non-CLM**	**CLM**
Individual characteristics				
Married	0.380	0.427	0.317	0.362
Cohabiting	0.087	0.073	0.080	0.069
Age	26.4	26.5	26.4	26.5
No high school diploma	0.115	0.128	0.138	0.151
Some college	0.366	0.358	0.329	0.326
College degree	0.183	0.175	0.158	0.151
Graduate degree	0.052	0.041	0.040	0.038
Black	0.159	0.151	0.137	0.131
Hispanic	0.082	0.132	0.085	0.133
Asian	0.019	0.007	0.020	0.008
Other race	0.013	0.013	0.013	0.012
Presence of children <6	0.288	0.315	0.182	0.211
Children 6-17	0.149	0.167	0.069	0.080
Number of children	0.833	0.935	0.464	0.556
Metro: central city	0.267	0.266	0.259	0.260
Metro: outside central	0.557	0.525	0.565	0.525
Unearned income	53,331	49,324	49,601	43,524
State characteristics				
Sex ratio	0.997	0.994	0.998	0.995
College educated adults	25.7	24.5	25.8	24.6
Unemployment rate	5.9	5.4	6.0	5.4
Median household income	51,957	49,083	52,046	49,216
Welfare	707	611	710	614
N	*243,926*	*78,063*	*222,034*	*70,427*

CPS 1995-2011. US born women and men, ages 18-35.
Notes. All differences are statistically significant. Means are weighted using survey weights.
Median household income is in Table H-8 at http://www.census.gov/hhes/www/income/data/historical/household/.
Unemployment rates are annual averages by state and are obtained from BLS (http://www.bls.gov/data/).
Sex ratios are calculated from Census state population estimates by dividing the number of men in each 5-year age group by the number of women who are 2 years younger. For example, in order to get a sex ratio for women aged 18-22, we divide the number of men aged 20-24 by the number of women aged 18-22. Sex ratios are calculated separately for white and black population and for the total population. Other races as assigned sex ratios for total population. Population data by age are obtained from http://www.census.gov/popest/data/intercensal/state/state2010.html and https://www.census.gov/popest/data/state/asrh/2012/SC-EST2012-ALLDATA6.html.
Welfare benefits are maximum TANF + SNAP benefits for a family of two, in 2010 dollars, and are obtained from the University of Kentucky Center for Poverty Research http://www.ukcpr.org/data.

Figure 2 presents women's age profiles of marriage, cohabitation, and in-couple residence (either marriage or cohabitation), extending beyond our sample's age to age 44. Relative to non-CLM states, for all age groups, there is a higher share of married and lower share of cohabiting women in CLM states. Cohabitation rises from 3% at age 18, to about 12% at age 24, and then declines back to 4% by age 44. The marriage profile is much steeper: starting at 3% at age 18, one third of women by age 24, and covering over 60% by age 32. After age 32, the marriage profile flattens and only grows by 7% percentage points between ages 32 and 44. The age-marriage and age-cohabitation profiles indicate that at all ages, women are more likely to be married and less likely to

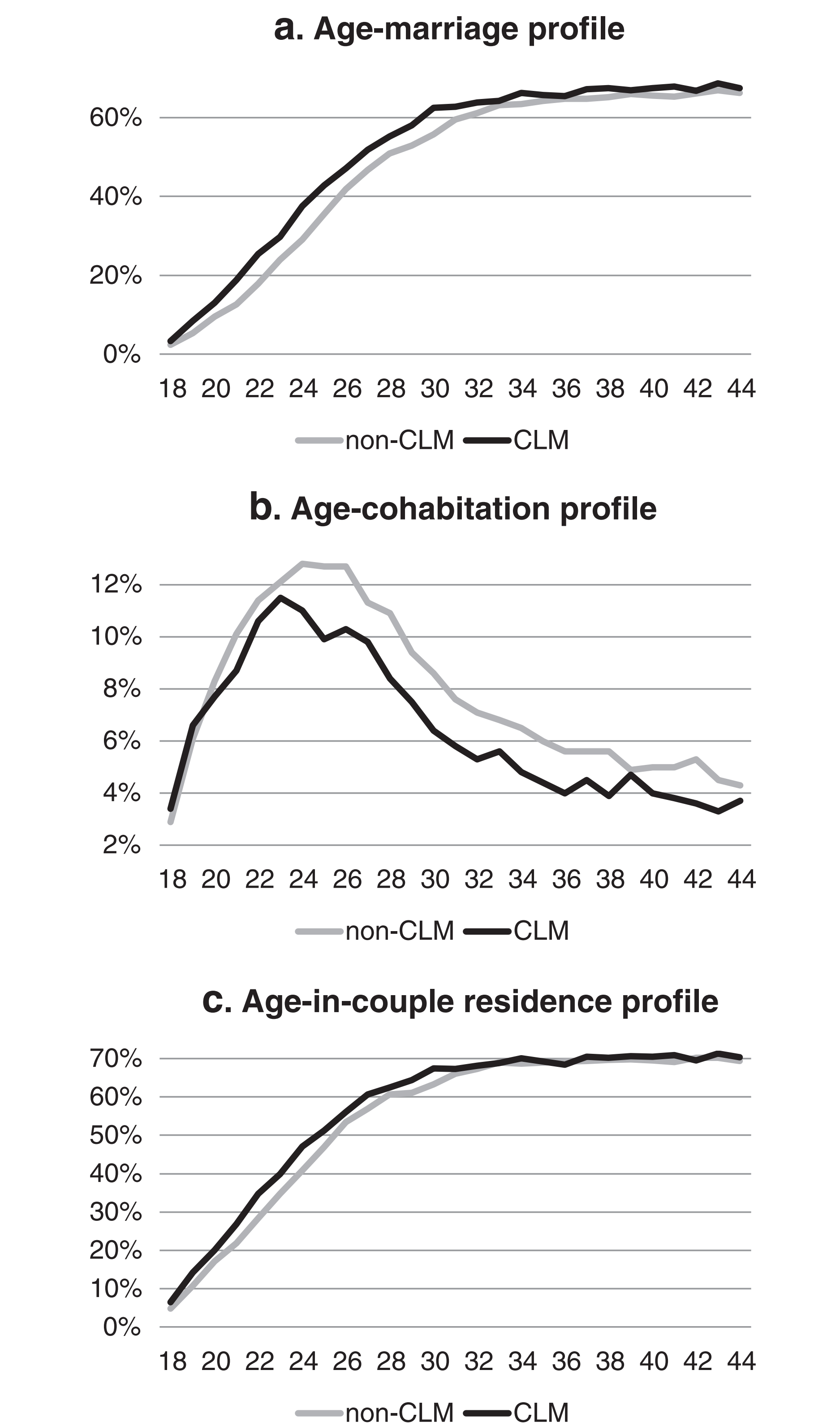

Figure 2 (See legend on next page.)

(See figure on previous page.)

Figure 2 Cohabitation and marriage profile by age and presence of CLM; all US-born women aged 18-44 in CPS 1995-2011. Results in Panel **a** are obtained from the full sample. Results in Panel **b** are obtained by restricting the corresponding subsamples to respondents from three transition states. Models in Panel **a** and Panel **b** are otherwise identical; other controls are shown at the bottom of the table.

cohabit in CLM states. The last panel indicates that in CLM states, women are more likely to be in couple than in non-CLM states until around age 32. After that, there is no visible CLM differential in in-couple residence.

Between 1995 and 2011 the percentage of women ages 18-35 residing in couples decreased in all three states that repealed CLM; the percentage of men in couples decreased in Georgia and (very slightly) in Pennsylvania. It grew slightly in Idaho (from 48% to 50%). Figure 3 presents shares of women and men who reside in couples. That share decreased slightly over time in non-CLM states for men and women. CLM states have higher shares of men and women who live in couples. We notice above-average rates for Idaho and below-average rates for Pennsylvania, which underscores the importance of including state fixed effects or state-specific trends into our estimations. The graphs suggest that couple formation may have increased in Idaho and decreased slightly in Georgia and Pennsylvania after the repeal of CLM.

Figure 4 shows that relative to Hispanic and black women, white women are more likely to be in couple. In non-CLM states, the in-couple residence rate for black women stands at 25%, amounting to less than half the rate for white women. Hispanic in-couple residence rates are much closer to those of whites. Similar proportions are obtained for CLM states.

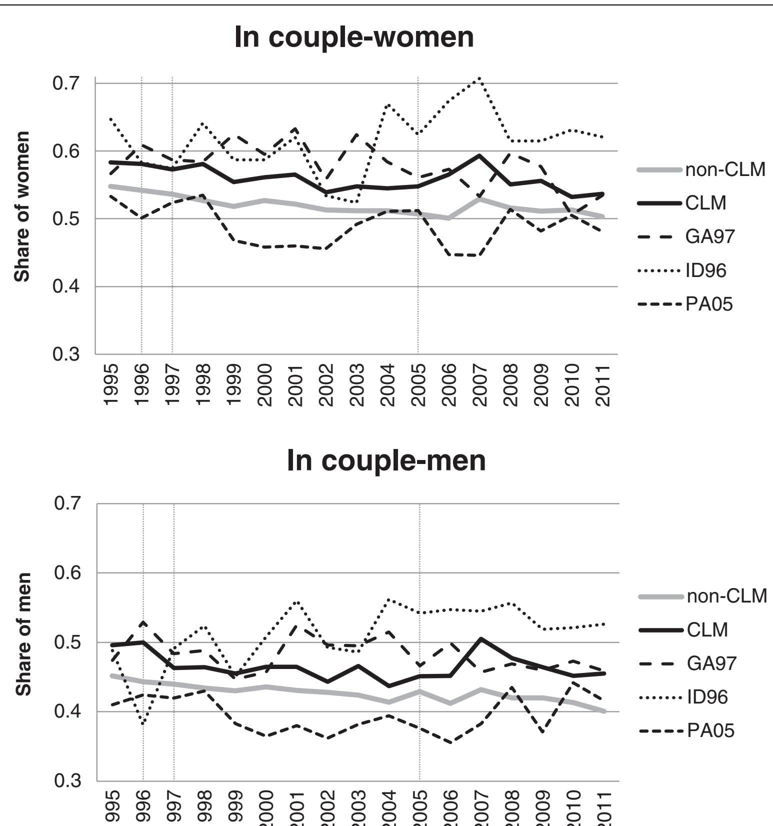

Figure 3 Evolution of in-couple residence rates over time, by presence of CLM.

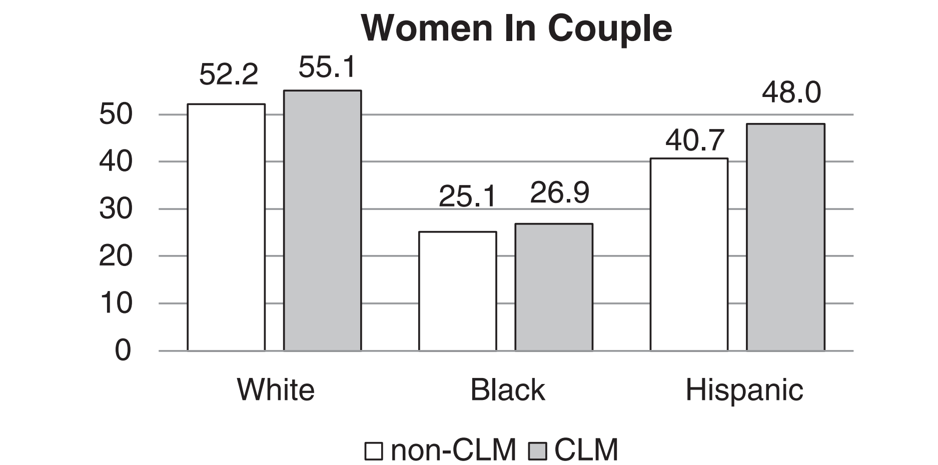

Figure 4 **Share of women in couple by race, CPS.**

4 Empirical strategy

Our empirical strategy is to use the individual-level CPS data to estimate a series of models where *Y*, the outcome of interest, is a function of CLM and other determinants of a decision. For individual *i* from state *s* in year *t*, outcome *Y* is:

$$Y_{ist} = \alpha CLM_{st} + \beta X_{ist} + \delta_s + \gamma_t + u_{ist} \tag{1}$$

where *Y* is one of the following probabilities for the entire sample: probability of (1) being in a couple (either married or cohabiting), or (2) being married (versus unmarried). In addition, for a sample of unmarried respondents, we estimate the probability of (3) cohabiting (versus being single). A fourth outcome we analyze is the probability of being divorced if ever-married. We estimate probit regressions for these three outcomes. Furthermore, as a robustness test, we estimate multinomial logit regressions of the log odds of being married or cohabiting relative to being single.

CLM, our variable of interest, indicates whether the state of residence recognizes CLM in year *t*;

δ_s are state fixed effects to account for unobservable differences in economic, legal, demographic and cultural environment that may affect individual choices, such as laws regarding child custody and religiosity;

γ_t are time dummies to capture the time trend; and u_{ijt} are i.i.d. error terms.

The vector of controls *X* consists of:

a. *Individual characteristics:* a quadratic function of age, 4 dummies for educational level, dummies for black, Hispanic, Asian and other ethnicity, two indicators for metropolitan residence (central city and outside central city with non-metropolitan with non-identifiable as a reference group), and log of personal non-labor income. We chose not to include potentially endogenous total household income and presence of children.
b. *State characteristics:* sex ratios calculated by respondents' age and ethnicity to reflect that most marriages are between people of the same ethnicity and that marriage market conditions vary by ethnicity [13]; unemployment rate to account for

economic conditions that may have had an impact on couple formation; log of median household income to capture aggregate economic conditions and the cost of living; share of college-educated, urban, Hispanic and black population to adjust for differing marriage market conditions.

Some specifications also include indicators for the individual being in one of the three transition states in the year immediately before the repeal of CLM, the year of the repeal and in the years immediately following the repeal. The full list of controls in vector X can be found in the Additional file 1, where we show full estimates of equation (1) for the probability of being in couple. We also experiment with including a dummy variable for the state of New York before August 2010, when New York adopted no-fault divorce, given that it was the last state to implement the divorce reform that other states adopted before 1985. The coefficients on this variable are highly significant and positive in most regressions, yet its presence does not change the results, and we don't include it in our final analyses. Results presented below include New Hampshire among CLM states, even though it only offers CLM in the case of a partner's death. Excluding this state does not significantly alter our results. In some of our trial runs, we included indicators for full-and part-time students. Students are significantly less likely to be in a couple than non-students. Yet, we choose to report the results without these controls because they are potentially endogenous.

Standard errors are clustered by state/year to adjust for correlated standard errors that are likely to arise due to common random effects at the state-year level. This is a necessary step because the unit of observation is at the individual-level while the variation is at the state-level (see Moulton 1990). Identification of a CLM effect arises through cross-state variation and variation over time as three states repealed CLM over the period examined.

If the availability of CLM increases couple formation, we will observe positive coefficient α in the equation for the probability of being in a couple. If CLM increases the odds of being married or cohabiting relative to staying single, the corresponding coefficients in the probit regressions for married and cohabiting will be positive. It is an empirical issue whether there is evidence consistent with these predicted effects of CLM.

We estimate all models for the entire sample including all ethnicities. Ideally, we would have liked to also present separate results for blacks, Hispanics, and whites given that most marriages are within the same race, and there are reasons to believe that marriage market conditions differ by ethnicity: means and standard deviations of the dependent variables are significantly different for the white and black subsamples. However, sample sizes of non-white groups are not sufficiently large for race-specific difference-in-difference analyses. For example, there are fewer than 30 Hispanic men and women in each of the transition states in most years, and there are virtually no blacks in the relevant years in Idaho.

We present results separately for men and women. We expect that since monogamy prevails overall, effects of CLM on couple formation of men and women will be similar. However, to the extent that we analyze smaller samples subdivided by personal characteristics, we expect gender differences in the effect of CLM due to different types of matching between men and women with different characteristics. Overall, men aged 18-35 are less likely to be in a couple than women, and if they are, they are more likely to be married and be in relationships with women of the same age group. If we find that despite these differences, CLM affects couple formation among men and women in similar ways it implies that our results are more robust.

5 Results

Table 2 shows estimates of CLM effects on in-couple residence, not distinguishing between marriage and cohabitation. These probit marginal effects show how much the probability of the outcome variable changes when the value of CLM changes from zero to one, holding other variables at their means. We estimate three versions of equation (1) for the full sample (Panel A) and separately for residents of the three transition states (Panel B). The basic model and the model with state characteristics include state and year fixed effects, whereas the third model includes state-specific time trends and excludes 3 years (1996-99, 2005-07) after the repeal of the CLM law in the transition states. These years are excluded in order to relax the assumption of quick adjustment to changes in the law and quick couple formation.

All coefficients are negative, except for one, suggesting that CLM is associated with lower rates of in-couple residence, as was predicted. The basic model's coefficients for all men and for women and men in the three transition states are statistically significant at the 5% level. CLM effects for the full sample and the three-state subsample are negative and statistically significant in the last model with state-specific time trends that excludes the years after the law change. In most regressions, the coefficients for men are larger in absolute value than they are for women, suggesting that relative to women,

Table 2 CLM marginal effects in regressions of individual probability of being in couple

	Basic model		State characteristics added		State-specific time trends added, 3 years after each repeal removed	
	Women	**Men**	**Women**	**Men**	**Women**	**Men**
A. Full sample						
CLM	-0.011	-0.028	-0.012	-0.025	-0.038	-0.065
	[0.011]	[0.012]**	[0.012]	[0.013]*	[0.018]**	[0.017]***
Observations	*321917*	*292376*	*321917*	*292376*	*202309*	*183543*
B. Only three transition states: Idaho (1996), Georgia (1997), Pennsylvania (2005)						
CLM	-0.03	-0.035	0.016	-0.002	-0.035	-0.061
	[0.012]**	[0.014]**	[0.013]	[0.017]	[0.018]*	[0.023]***
Observations	*22176*	*19786*	*22176*	*19786*	*13708*	*12275*
Other controls in the model						
Individual characteristics	Yes	Yes	Yes	Yes	Yes	Yes
State characteristics			Yes	yes	yes	Yes
Year dummies	Yes	Yes	Yes	Yes		
State dummies	Yes	Yes	Yes	Yes	Yes	Yes
State-specific time trends					Yes	Yes

Notes: This table shows estimates of the probit marginal effects from equation (1) estimated at the mean. Each entry comes from a separate regression. All regressions include individual demographic characteristics and state fixed effects. Basic model and model with state characteristics include year fixed effects. The last model does not include year fixed effects; instead it includes state-specific time trends, and it excludes 3 years after the repeal of the law in each transition state (1996-99, 2005-07 are excluded).
Here and in all tables: *significant at 10%; **significant at 5%; ***significant at 1%. Robust standard errors are clustered by state and year, shown in brackets; individuals' survey weights are used.

men responded more to CLM changes and were quicker to adapt to the repeal of CLM. In part, this could be because fewer men are in a couple by age 35, our sample's cutting point (see footnote 5).

Table 3 shows two more versions of equation (1), estimated for the full sample and a sample of respondents from transition states. Here we keep all years in the data set but include an indicator for one year immediately before the repeal of the law in transition states, a dummy for year of abolition in those states, and four dummies for subsequent years in transition states. The dummy for year before repeal is positive, suggesting that couple formation increased in transition states several months before CLM was repealed. This may be either because some couples or individuals who wanted to establish CLM rushed to do so or because the deterring feature of CLM–whatever it may be–ceased to discourage couples from moving in together.

All coefficients for CLM in the year of repeal and in the three following years after the repeal are negative. This could indicate the continuing effect of past availability of CLM that possibly discouraged past couple formation. Alternatively, it could be a reaction to excess couple formation the year before abolition. Again, the coefficients for men tend to be larger in absolute value than those for women. Year four after the repeal shows a small increase in couple formation among women and no effect for men. This could indicate more active couple formation among new couples encouraged by the repeal of CLM. By year five after the repeal, we find no impact of CLM.

Our preferred specification appears in columns 3 and 4 of Table 3: it is estimated for all states and includes state-specific time trends as well as dummies for CLM at different points in time surrounding the date of the repeal in the transition states.

Table 4 shows estimates of CLM effects on the probability that (i) a respondent is married as opposed to either cohabiting or being single (columns 1 and 2), (ii) an unmarried respondent is cohabitating (columns 3 and 4), and (iii) an ever-married respondent is separated or divorced. The models are similar to the first model in Table 3. The CLM coefficients in the marriage and cohabitation regressions are negative and highly significant for men but only marginally significant for women. There is an increase in cohabitation the year before and the year after CLM's repeal, which suggests that cohabiting couple formation responds to CLM changes faster than married cohabitation. Marriage probability is not significantly affected by CLM the year before repeal, or in the year after repeal in the case of women. However the past availability of CLM has a lingering effect on marriage 1-3 years after the repeal: there is a lower probability of being married in states that repealed CLM and had it two or three years before. This holds for both men and women.

Panel B indicates that effects of CLM on probability of being married are also negative when only considering the residents of three transition states. These small sample estimates show limited impact of CLM on cohabitation, although regressions with state-specific time trends (not presented here) show an increase in cohabitation the year before repeal.

The last two columns in the same table show effects of CLM on the probability of being separated or divorced if ever married: Panel A for the full sample and Panel B for respondents from transition states. For men in both panels, some of the CLM effects on probability of being divorced are positive. Based on the whole sample, women from transition states were less likely to be divorced four or five years after CLM's repeal.

Table 3 Individual probability of being in couple as a function of CLM availability at different points in time

	State and time fixed effects		**State-specific time trends**		**Only 3 transition states**		**Only 3 transition states, state-specific time trends**	
	Women	**Men**	**Women**	**Men**	**Women**	**Men**	**Women**	**Men**
	1	**2**	**3**	**4**	**5**	**6**	**7**	**8**
CLM	-0.021	-0.051	-0.028	-0.059	-0.031	-0.033	-0.033	-0.069
	[0.010]**	[0.010]***	[0.020]	[0.021]***	[0.020]	[0.022]	[0.021]	[0.019]***
Year before repeal	0.028	0.035	0.028	0.032	0.004	0.037	0.027	0.033
	[0.011]**	[0.016]**	[0.016]*	[0.012]***	[0.017]	[0.019]*	[0.014]*	[0.009]***
Year of the repeal	0.015	-0.037	0.015	-0.04	-0.052	-0.019	0.011	-0.045
	[0.012]	[0.014]***	[0.017]	[0.018]**	[0.018]***	[0.028]	[0.016]	[0.017]***
One year after repeal	-0.031	-0.046	-0.032	-0.052	-0.096	-0.027	-0.033	-0.054
	[0.011]***	[0.018]**	[0.014]**	[0.023]**	[0.016]***	[0.023]	[0.014]**	[0.021]**
Two years after repeal	-0.046	-0.055	-0.039	-0.045	-0.095	-0.002	-0.037	-0.044
	[0.021]**	[0.011]***	[0.017]**	[0.015]***	[0.017]***	[0.027]	[0.015]**	[0.015]***
Three years after repeal	0.02	-0.005	0.028	0.003	-0.029	0.023	0.033	0.007
	[0.008]**	[0.027]	[0.010]***	[0.030]	[0.016]*	[0.034]	[0.010]***	[0.031]
Four years after repeal	0.022	0	0.023	0.001	-0.02	0.026	0.03	0.008
	[0.017]	[0.016]	[0.016]	[0.018]	[0.016]	[0.021]	[0.017]*	[0.020]
State characteristics	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Year dummies	Yes	Yes			Yes	Yes		
State dummies	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
State-specific time trends			Yes	Yes			Yes	Yes
Observations	*321917*	*292376*	*321917*	*292376*	*22175*	*19782*	*22175*	*19782*

Notes: This table shows estimates of the probit marginal effects from equation (1) estimated at the mean. Each column comes from a separate regression. All regressions include individual demographic characteristics and state fixed effects. Models 1 and 3 include state fixed effects. Models 2 and 4 include state-specific time trends.
*significant at 10%; **significant at 5%; ***significant at 1%.

Table 4 CLM and probability of being married, cohabiting and divorced; probit marginal effects

	Probability of being married among all respondents		Probability of cohabiting among unmarried		Probability of being divorced or separated among ever-married	
	Women	Men	Women	Men	Women	Men
	1	2	3	4	5	6
A. Full sample						
CLM	-0.016	-0.033	-0.016	-0.026	0.011	0.026
	[0.009]*	[0.010]***	[0.009]*	[0.007]***	[0.014]	[0.018]
Year before repeal	0.004	0.001	0.035	0.044	-0.002	0.018
	[0.011]	[0.015]	[0.009]***	[0.008]***	[0.016]	[0.034]
Year of the repeal	-0.004	-0.043	0.023	0.006	-0.012	0.045
	[0.008]	[0.011]***	[0.009]***	[0.006]	[0.009]	[0.019]**
One year after repeal	-0.028	-0.033	-0.013	-0.016	-0.006	0.035
	[0.008]***	[0.007]***	[0.014]	[0.013]	[0.010]	[0.015]**
Two years after repeal	-0.048	-0.042	-0.011	-0.02	0.018	0.044
	[0.017]***	[0.010]***	[0.009]	[0.007]***	[0.021]	[0.029]
Three years after repeal	0.005	-0.014	0.015	0.006	-0.031	0
	[0.014]	[0.015]	[0.014]	[0.014]	[0.017]*	[0.010]
Four years after repeal	0.018	0.002	0.003	-0.003	-0.052	-0.005
	[0.011]*	[0.015]	[0.010]	[0.005]	[0.012]***	[0.009]
Observations	*321917*	*292376*	*190655*	*186020*	*161687*	*123368*
B. Only three transition states						
CLM	-0.046	-0.058	0.02	0.024	0.078	0.069
	[0.018]***	[0.015]***	[0.013]	[0.014]*	[0.020]***	[0.028]**
Year before repeal	0.011	0.025	-0.014	0.014	0.006	-0.012
	[0.017]	[0.016]	[0.008]*	[0.012]	[0.013]	[0.025]

Table 4 CLM and probability of being married, cohabiting and divorced; probit marginal effects *(Continued)*

Year of the repeal	-0.044	-0.035	-0.013	0.018	0.057	0.05
	[0.015]***	[0.019]*	[0.011]	[0.015]	[0.023]**	[0.029]*
One year after repeal	-0.085	-0.04	-0.017	0.016	0.099	0.044
	[0.015]***	[0.014]***	[0.009]*	[0.014]	[0.021]***	[0.029]
Two years after repeal	-0.095	-0.028	-0.01	0.024	0.126	0.025
	[0.017]***	[0.016]*	[0.009]	[0.019]	[0.024]***	[0.038]
Three years after repeal	-0.03	-0.001	-0.01	0.022	0.029	-0.011
	[0.016]*	[0.020]	[0.012]	[0.019]	[0.020]	[0.020]
Four years after repeal	-0.004	0.019	-0.017	0.009	-0.003	0.009
	[0.014]	[0.013]	[0.008]**	[0.010]	[0.016]	[0.017]
Observations	*22176*	*19786*	*12662*	*12218*	*11465*	*8758*
Other controls: demographics, state characteristics, year & state fixed effects	Yes	Yes	Yes	Yes	Yes	Yes

*significant at 10%; **significant at 5%; ***significant at 1%.

However, when only the transition states are considered (Panel B), CLM is associated with a higher probability of being divorced or separated, especially for women, suggesting either that CLM facilitated couple formation among some individuals who were not well-matched or that CLM discouraged remarriage/recoupling among some of the women who were separated or divorced from a previous partner.

5.1 Testing predictions 2 to 5

We test the predictions that were derived from the theoretical model in Section 2 for the outcome "probability of being in couple". From a policy standpoint, couple formation is a more interesting outcome than marriage or cohabitation. It has been shown that children benefit from growing up with two parents, and children born to single mothers typically grow up with less of a father's presence in their lives than children born to a cohabiting unwed couple (see Mincy and Oliver 2003). Relative to children raised by a mother and a father, children raised by a single parent often achieve less in terms of school performance (McLanahan and Sandefur 1994, McLanahan and Sigle-Rushton 2004) and have higher rates of depression and crime participation (Hobcraft 1998, Sigle-Rushton et al. 2005). However, we also examine separate effects of CLM on the likelihood of marriage (vs. cohabitation and single status) and the probability of cohabitation (for unmarried respondents) for separate samples corresponding to our predictions. All regressions use our preferred model with state and year fixed effects and indicators for the years before and after the repeal as in model 1 of Table 3.

In the first two rows and first two columns of Table 5, CLM marginal effects on in-couple residence are estimated for subsamples of respondents split by education. Prediction 2 implies that we should find larger effects of CLM on women's probability of being in-couple for women who are less educated. It can be seen from Table 5 that the CLM coefficient is negative and significant in regressions for women without a college degree, but it is zero for college-educated women. We also find negative effects of CLM on cohabitation probability of these women without a college education.

However, we predicted that couple formation among men with a college degree would respond more to CLM changes than among men without a degree (prediction 3). We find that all men have a lower probability of in-couple residence where CLM is available. However, in accordance with the prediction, we find a stronger response to CLM law changes for college educated men (the coefficient of CLM for college educated men is -.108) than for men without college (-.038). We conclude that there is evidence for prediction 3 indicating that men who have a college degree have more to loose from CLM in case of divorce.

CLM's effect of men's likelihood of being in a couple seems to originate from a negative effect of CLM on both the likelihood of marriage and the likelihood of cohabitation. CLM's effect on the likelihood of marriage of men with a college education is much larger than its effect on marriage among men without college. However, effects of CLM on cohabitation versus being unmarried and single are stronger for those without a college education than on those who are college educated, reflecting the much higher rates of marriage among the college-educated in the US (Fry 2014).

Prediction 4 stated that drops in women's couple formation resulting from CLM would be larger in low sex ratio states than in high sex ratio states. It can be seen from

Table 5 CLM marginal effects by subsample

	In couple probit		Married probit		Cohabiting, among unmarried probit	
	Women	**Men**	**Women**	**Men**	**Women**	**Men**
No college	-0.031	-0.038	-0.016	-0.018	-0.027	-0.024
	[0.012]***	[0.010]***	[0.010]	[0.011]*	[0.009]***	[0.008]***
College educated	-0.006	-0.108	-0.034	-0.103	0.036	-0.03
	[0.025]	[0.025]***	[0.025]	[0.022]***	[0.028]	[0.018]*
Low sex ratio	-0.042	-0.084	-0.024	-0.046	-0.01	-0.031
	[0.017]**	[0.020]***	[0.012]**	[0.012]***	[0.013]	[0.012]***
High sex ratio	-0.024	-0.029	-0.024	-0.025	-0.032	-0.021
	[0.017]	[0.018]	[0.015]	[0.020]	[0.013]**	[0.007]***
No children present	-0.035	-0.041	-0.027	-0.017	-0.015	-0.023
	[0.012]***	[0.009]***	[0.009]***	[0.007]**	[0.010]	[0.007]***
Children present	-0.015	-0.012	-0.007	-0.013	-0.018	-0.044
	[0.016]	[0.010]	[0.019]	[0.012]	[0.019]	[0.051]
Age 18-25	-0.011	-0.046	0.001	-0.015	-0.012	-0.031
	[0.013]	[0.009]***	[0.011]	[0.007]**	[0.010]	[0.007]***
Age 26-35	-0.022	-0.038	-0.021	-0.039	-0.02	-0.01
	[0.015]	[0.016]**	[0.014]	[0.016]**	[0.017]	[0.020]

Notes: The table reports CLM marginal effects in regressions for probability of being in a couple. High sex ratio areas are those with sex ratio >1 and low are sex ratio < =1.
The table shows full estimates of the first model from Table 3, our preferred specification with demographics, state characteristics, year & state fixed effects, and six dummy variables for years before and after repeal of the law.
All pairs of coefficients estimated using two subsamples are statistically different from each other at less than 1% level.
Sample sizes in the first four columns vary between N = 55,908 for men with college degrees to N = 249,659 for women with less than college education.
Sample sizes among unmarried adults in the last two columns vary between N = 10,450 for men with children to N = 159,883 for men with less than college education.
*significant at 10%; **significant at 5%; ***significant at 1%.

rows 3 and 4 in Table 5 that the coefficient of CLM is negative in most regressions for both women and men when sex ratios are low, the only exception being that CLM does not affect women's likelihood of cohabitation when sex ratios are low. In contrast, CLM effects on likelihood of in-couple residence and likelihood of marriage are zero when sex ratios are high. This is consistent with what we argued: where sex ratios are low, CLM potentially boosts the price of women's WiHo, but not where sex ratios are high. That we find this effect for both men and women makes sense given that monogamy prevails. We also find significant negative CLM effects on the probability of cohabitation among the unmarried where sex ratios are high.

Prediction 5 was that CLM effects on couple formation are more likely to be found for childless respondents than for those with children. It can be seen from rows 5 and 6 in Table 5 that when no children are present, CLM coefficients are negative for both men and women and for all outcomes (except for the cohabitation vs. single regression for women). When there are no children, less WiHo is likely to be performed, and CLM as a form of minimum compensation for WiHo is expected to have more impact on couple formation. In contrast, in all regressions of Table 5, the coefficient of CLM is insignificant for both men and women with children. To the extent that presence of children captures individuals with children from a previous relationship, this finding could reflect the fact that WiHo workers may be getting a lower y for their WiHo in the marriage market if they have children from previous relationships.

According to Prediction 6, CLM is more likely to have an effect on the likelihood of being in couple in the case of the youngest respondents than in the case of those somewhat older. This may be gender-specific. The last two rows in Table 5 indicate that younger men below age 26 are less likely to cohabit in CLM states, whereas men age 26-35 are less likely to get married. The impact of CLM does not vary by age for women.

5.2 Robustness checks

Our results suggest negative effects of CLM on couple formation. In order to ascertain that these results are not spurious, we conduct three additional robustness checks. Table 6 reports log odds of being in a married or cohabiting relationship relative to being single from the multinomial logit model over three outcomes: married, cohabiting, and single. These equations include fixed time and state effects along with personal demographic and state characteristics. The odds are negative for men and women. Living in a CLM state reduced the log-odds of being married or cohabiting relative to staying single while holding all other variables in the model constant. Cohabitation rises the year before the repeal as evidenced by large and positive coefficients on the corresponding dummy variable for both men and women. Cohabitation increases for women during the repeal year. Marriages decline temporarily 1-3 years after the repeal with men experiencing a larger decline in marriages compared to women.

Table 6 Robustness check: multinomial logit log odds of cohabiting or being married relative to being single

	Women		Men	
	Married	**Cohabiting**	**Married**	**Cohabiting**
CLM	-0.084	-0.188	-0.214	-0.33
	[0.043]*	[0.081]**	[0.056]***	[0.085]***
Year before repeal	0.062	0.32	0.056	0.469
	[0.053]	[0.065]***	[0.077]	[0.070]***
Year of the repeal	0.012	0.193	-0.25	0.063
	[0.052]	[0.076]**	[0.072]***	[0.067]
One year after repeal	-0.14	-0.158	-0.209	-0.231
	[0.035]***	[0.147]	[0.056]***	[0.189]
Two years after repeal	-0.227	-0.103	-0.262	-0.207
	[0.093]**	[0.081]	[0.060]***	[0.079]***
Three years after repeal	0.055	0.127	-0.044	0.065
	[0.054]	[0.098]	[0.108]	[0.144]
Four years after repeal	0.099	0.032	0.014	-0.027
	[0.068]	[0.098]	[0.085]	[0.056]
Demographics, state characteristics, year & state fixed effects	Yes	Yes	Yes	Yes
Observations	*321917*	*321917*	*292376*	*292376*

Notes: Multinomial logit model shows log odds of being married or cohabiting relative to being single; estimated separately for men and women. Full set of controls included into the regressions.
*significant at 10%; **significant at 5%; ***significant at 1%.

Our next test is based on placebo laws. The difference-in-difference (DD) approach that we used for identification often suffers from a serial correlation problem. Bertrand et al. (2004) have shown that standard errors of DD estimators are often underestimated, and thus the statistical significance of the coefficients is overestimated. We repeat their experiment with 'placebo laws' in order to assess whether our results are reliable or could be due to random coincidence. We remove the three actual transition states from the sample and randomly assign any three states to be CLM states from 1995 until a random year between 1996 and 2006. We then estimate the impact of these fake laws on our outcomes of interest: in couple residence, the probability of being married, and the probability of cohabiting. For this experiment, we use our preferred model with state and year fixed effects and six dummies for years before/after abolition. We repeat this procedure 100 times and record the number of times the null hypothesis of no effect is rejected as well as the direction of the estimated 'effect'. We find that the non-existent laws are significant 22% and 17% of the time in in-couple regressions for women and men, as well as correspondingly 20%, 21% in marriage probability regressions, and 23% and 24% of the simulations in cohabitation regressions. In all simulations, significantly positive and negative effects were equally likely, for example, out of 17 significant effects of fake laws on in-couple residence among men, 8 are negative and 9 are positive. These non-rejection rates are higher than the 5% that can be conventionally attributed to randomness. Thus, we conclude that this test is inconclusive. In some simulations, the CLM dummy possibly picks up the average differences between existing CLM and non-CLM states or coincides with other changes that may have occurred in randomly chosen states. We therefore design an additional test to check the robustness of our findings of negative effects of CLM on the probability of being married and cohabiting.

According to Bertrand et al. (2004), one of the best ways to deal with serial correlation in standard errors, if the problem is suspected, is to estimate a panel data model using individual-level data aggregated by gender/year/state cells. We compute these estimates and record the results of this in Table 7. First, we regress the binary data on whether the person is in a couple on personal characteristics. Then we calculate means of residuals by year/state and regress the mean of residuals on CLM, state characteristics, and state and year fixed effects. These are linear probabilities, not probits, yet we once again obtain negative effects of CLM on the probability of being in a couple and being married, negative coefficients for two years following the repeal, and larger overall effects for men than for women. None of the coefficients of interest in the cohabitation equations are statistically significant.

5.3 Other coefficients, based on the Additional file 1

A higher sex ratio, i.e., a higher ratio of men per woman (using as two year age difference as explained in Section 4) strongly increases the odds that both men and women are in couple. More specifically, a unit increase in sex ratio from 1 to 2 men per woman increases women's probability of being in a couple by 22.5%, according to the sex ratio coefficient in the first column. In other words, a 10% increase in sex ratio (from 1 to 1.1) is expected to increase women's probability of being in couple by 2.25%. A higher unemployment rate in the last year causes a slight increase in the odds that the person is in a couple. Men and women who live in higher income states (as measured by median state income) are more likely to have a spouse or a partner. The generosity of

Table 7 Robustness check: OLS estimates of linear probability models with data aggregated by state/year

	In couple		Married		Cohabiting	
	Women	Men	Women	Men	Women	Men
	1	2	3	4	5	6
CLM	-0.017	-0.046	-0.022	-0.04	-0.002	-0.019
	[0.016]	[0.016]***	[0.015]	[0.015]***	[0.016]	[0.014]
Year before repeal	0.009	0.016	-0.008	-0.01	0.021	0.03
	[0.022]	[0.022]	[0.020]	[0.020]	[0.022]	[0.019]
Year of the repeal	0	-0.072	-0.015	-0.081	0.009	-0.013
	[0.020]	[0.020]***	[0.018]	[0.018]***	[0.020]	[0.017]
One year after repeal	-0.03	-0.038	-0.036	-0.041	-0.005	-0.007
	[0.020]	[0.020]*	[0.018]**	[0.018]**	[0.020]	[0.017]
Two years after repeal	-0.02	-0.027	-0.03	-0.022	0.002	-0.02
	[0.020]	[0.020]	[0.018]	[0.018]	[0.020]	[0.017]
Three years after repeal	0.002	-0.024	-0.004	-0.02	-0.005	-0.021
	[0.020]	[0.020]	[0.018]	[0.018]	[0.020]	[0.017]
Four years after repeal	0.011	0.001	0.007	0.001	-0.003	-0.008
	[0.020]	[0.020]	[0.018]	[0.018]	[0.020]	[0.017]
Demographics included in the first stage, state characteristics, year & state fixed effects	Yes	Yes	Yes	Yes	Yes	Yes
Observations	*867*	*867*	*867*	*867*	*867*	*867*

Notes: To obtain these estimates, we first regress data on individual-level controls and collect residuals. Then we compute average residuals by year and state, and regress them on year and state fixed effect.
*significant at 10%; **significant at 5%; ***significant at 1%.

welfare payments has a small negative impact on union formation. A higher percentage of black residents in the state is associated with a significant drop in the odds that the respondent is in couple; a higher share of Hispanics in associated with a reduced probability that men are in couple. A higher share of educated residents increases men's odds of being in couple. Household formation has a concave age profile. Education has a strong positive effect on the odds of being in a couple. Living in a metropolitan area, especially in central city, reduces the odds of being in couple. Men and women of all other ethnic groups are less likely to be in a relationship than their white counterparts. Higher own unearned income is positively related to men's probability of being in couple but negatively related to that same probability for women.

6 Discussion and conclusions

This paper examined whether the availability of Common-Law Marriage (CLM) helps explain in-couple residence, marriage, and cohabitation among young men and women in the US. A difference-in-difference analysis was performed given that during the period examined, three states repealed CLM. Results using native CPS respondents under age 36 reveal that CLM reduces in-couple residence and probability of marriage among both women and men. Effects are larger for men than for women. We obtain these results when considering all states and for a sample of the three states that repealed CLM. Couple formation increases several months before the repeal and then decreases for up to 3 years during the adjustment time following the repeal.

We presented a model based on Becker's theory of marriage that considers CLM as setting a minimum price for women's Work-In-Household (WiHo). Assuming traditional gender roles, we derived gender-specific predictions regarding differential effects of CLM by male and female education, sex ratio, parental status, and age. We predict that college-educated men are more likely to experience reduced couple formation under CLM than their counterparts without a college education, but that the opposite is the case when comparing CLM's effect on college educated-women with that on women without a college education. We find that CLM has asymmetric effects on couple formation and cohabitation for men and women in accordance with our predictions.

Other predictions we find evidence for are that the couple formation effects of CLM are stronger in states with low sex ratios (ratios of men to women) and in the case of childless respondents.

Our basic finding implies that repeal of CLM encouraged couple formation. Since overall couple formation rates have decreased in the US during this period, it follows that other reasons have led to this drop in couple formation. One of our recommendations aimed at encouraging couple formation is for the remaining 11 states to repeal CLM. This may be especially beneficial in areas from which men migrate more than women, leaving populations with low sex ratios. We found larger negative effects of CLM on in-couple residence of women without a college education and on that of men with a college education. As a result, the repeal of CLM has led to larger increases in couple formation by low educated women (relative to those with high education) and by high educated men (relative to those with a low education).

In this analysis, we have assumed that the repeal of CLM is an exogeneous change. We realize that changes in legislation are not spurious: factors that have led to increases in in-couple residence rates may also have pushed states to repeal CLM laws. One of these factors may be social norms that are increasingly tolerant of cohabitation and accepting of an egalitarian division of labor within the household. The more egalitarian a society's gender norms, the more households are formed (Sevilla-Sanz 2010). CLM goes against that trend: by providing marriage-like protection to those who perform the household production (typically women), it discourages men from cohabitating.

This has been an exploratory study. It is the first to suggest that couple formation, marriage, and cohabitation are affected by Common Law Marriage legislation. More research on these laws' effects on couple formation is needed, including further econometric evidence for the United States and other countries that underwent similar legal changes. Endogenizing legal change would also be a welcome direction for future research. It is hoped that new conceptual contributions about marriage, cohabitation, couple formation, and CLM will be offered and that they will continue to be inspired by Gary Becker's economic theories of marriage.

Endnotes

[1]In 2002, 40% of first premarital cohabitations among women transitioned to marriage within 3 years (Goodwin et al. 2010).

[2]Popular sites with information on CLM:

http://video.about.com/marriage/How-to-Qualify-for-a-Common-Law-Marriage.htm;

http://www.unmarried.org/common-law-marriage-fact-sheet/;

Common Law Marriage Handbook for government employees who handle claims: http://www.dol.gov/owcp/energy/regs/compliance/PolicyandProcedures/CommonLaw_Marriage.pdf.

[3]Travis county in Texas offers CLM couples to fill out a Vital Statistics form, which suggests that some CLMs are recorded among New Marriages: http://www.co.travis.tx.us/dro/common_law.asp.

[4]Only one of Becker's Demand and Supply models of marriage appears in Becker's (1981) *Treatise on the Family*.

[5]The percentage of men ever married by age 34 was 73.8% in 1990 and 63% in 2010. For women, these percentages were 81.8% and 72.8% (UN world marriage data at http://www.un.org/esa/population/publications/WMD2012/MainFrame.html) for 1990. For the US in 2010, the percent ever married was computed using the US census table creator at http://www.census.gov/cps/data/cpstablecreator.html.

[6]WiHo may also benefit the self. More on this Beckerian theory of marriage can be found in Grossbard (2015).

[7]In this model, men and women remain individual decision-makers after marriage. No *gain from marriage* is calculated as is the case in Becker's theory of marriage. The larger the transfer of goods or money from the spouse employing WiHo to the WiHo-worker, the larger share of the gain from marriage obtained by the WiHo-worker.

[8]Spouses could also pay WiHo-workers with love, but the compensation for WiHo is expected to mostly contain a material component.

[9]Markets for WiHo are similar to the markets for wives or husbands found, e.g., in Becker (1973) and Choo and Siow (2006).

[10]Men may also supply WiHo to their wife, especially if they earn less than she does.

[11]Also see Brien and Sheran (2003). A possible reason why educated women are more in demand in WiHo markets and benefit from a higher equilibrium y than women without a college education is that they are more productive in WiHo. For example, they may be more productive at educating shared children and better household managers. All women with a higher y (including college educated women) have more bargaining power within their relationships and are less likely to be affected by CLM laws.

[12]https://cps.ipums.org/cps/.

[13]Sex ratios are computed by dividing the number of men in age groups 20-24, 25-29, 30-34, 35-39 by the number of women two years younger (see Grossbard 2015, Chapter 6).

Additional file

Additional file 1: Estimates of the probability of being in couple. Full model.

Competing interest
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