DISCUSSION PAPER SERIES

IZA DP No. 8480

Common Law Marriage and Couple Formation

Shoshana Grossbard
Victoria Vernon

September 2014

Forschungsinstitut
zur Zukunft der Arbeit
Institute for the Study
of Labor

Common Law Marriage and Couple Formation

Shoshana Grossbard
San Diego State University
and IZA

Victoria Vernon
Empire State College, New York

Discussion Paper No. 8480
September 2014

IZA

P.O. Box 7240
53072 Bonn
Germany

Phone: +49-228-3894-0
Fax: +49-228-3894-180
E-mail: iza@iza.org

IZA Discussion Paper No. 8480
September 2014

ABSTRACT

Common Law Marriage and Couple Formation

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 U.S. CLM effects are identified through cross-state and time variation, as some states abolished 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 and parent status. CLM reduces in-couple residence, and more so for childless whites and where there are fewer men per woman. Effects are larger for college-educated men and women without college.

JEL Classification: J10, J12, J16

Keywords: Common-Law Marriage, couple, couple formation, marriage, cohabitation, Gary Becker

Corresponding author:

Shoshana Grossbard
Department of Economics
San Diego State University
San Diego, CA 92182-4485
USA
E-mail: shosh@mail.sdsu.edu

Common Law Marriage and Couple Formation

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 Report, 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%.

A major reason why couples are formed is to provide a protective environment for children. From the point of view of child advocacy the distinction between parents living in couple and single parents is crucial. 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). Couple formation also matters due to its impact on the demand for goods and services such as housing and childcare. In this paper we examine whether in-couple residence, marriage, and cohabitation rates can be 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. 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 is that it does not exist until claimed by one of the partners. Otherwise CLM is like marriage, including its acceptance by all other states and government institutions dealing with tax collection and redistribution of income.

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.[2] The internet contains a lot of information and legal advice for couples about CLM, so we know it is practiced.[3] Some indication of CLM's prevalence can be derived

---

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

[2] Travis county in Texas offers CLM couples to full 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

[3] Popular sites with information on CLM:
http://video.about.com/marriage/How-to-Qualify-for-a-Common-Law-Marriage.htm;
http://www.answers.com/topic/common-law-marriage ;
http://www.unmarried.org/common-law-marriage-fact-sheet.html;
Common Law Marriage Handbook for government employees who handle claims:
http://www.dol.gov/owcp/energy/regs/compliance/PolicyandProcedures/CommonLaw_Marriage.pdf

from a legal historian's reporting of about one hundred legal CLM–related judgments being issued each decade in each state at the federal level (Lind, 2008).

We circumvent this data problem by exploiting cross-state variation in CLM and changes over time. Most US states recognized CLM in the past but have abolished this form of marriage. As of 2014, common-law marriage could be contracted in 11 states: Alabama, Colorado, Iowa, Kansas, Montana, New Hampshire (only posthumously for purposes of inheritance), Oklahoma, Rhode Island, South Carolina, Texas, and Utah, as well as in the Navajo Nation and in the District of Columbia. Over the period covered by our data—1995 to 2011--CLM was abolished by Idaho (1996), Georgia (1997), and Pennsylvania (2005), thus providing us with a quasi-experiment.

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 education, sex ratios, and parental status. The model assumes traditional gender roles, with women more involved in home production than men, and also predicts 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. Empirical U.S.-based research has tested whether the emergence of no-fault and unilateral divorce has affected divorce rates (e.g. Peters 1986, Friedberg 1998, and Wolfers 2006), marriage rates (e.g. Alesina and Guiliano 2007, 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 Suarez Becerra (2012) studied its effect on fertility. Variations in laws regulating division of property in 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 and Vernon forthcoming) we analyzed effects of CLM on labor supply.

Sex ratios play a central role in our model, as they have 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 and shifts from cohabitation to marriage (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 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: no lags are introduced. In other models we relax that

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

assumption and exclude three years of data after the abolition of CLM in each of the relevant states. We limit our analyses to respondents under age 36, as they are most likely to be entering marriage.

Consistent with our predictions, we find that CLM is associated with lower rates of in-couple residence, especially for whites in states with low sex ratios. CLM's effect on women varies negatively with college education: more educated women's probability of being in couple is less affected by CLM. The opposite is the case for men: if they have a college education their probability of being in couple is more affected by CLM. Childless men and women are more likely to respond to changes in CLM. Our demand and supply model of marriage helps us 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). It also incorporates the concept of *Work-In-Household (WiHo)* defined as household production work of benefit to a spouse/partner and that may include activities such as parenting and meal preparation (see Grossbard-Shechtman 1984).[5] It is assumed that people demand and/or supply this type of work. A couple may sometimes exchange WiHo, but if one partner works relatively more in WiHo the other may 'pay' for WiHo in the form of an intra-couple transfer. A higher price means that the individual WiHo-worker obtains more access to the gain from marriage and has higher bargaining power. Marriage markets are viewed as markets for WiHo. It is assumed that heterosexuality prevails and that there are many interrelated WiHo markets defined by personal characteristics of men and women (such as education and age). Each WiHo market establishes an equilibrium implicit "price" as in Becker' (1973) markets for wives or husbands.[6]

The price a partner pays to a cohabiting WiHo worker may be lower than the price a spouse pays to a married WiHo worker due e.g. to fewer benefits in case of death or divorce (Grossbard, Mincy and Huang 2005). To the extent that a CLM law offers more material benefits to some WiHo workers than they would otherwise get based on the market price of their work, it can be viewed as the equivalent of a minimum wage law. It is as if the state says: "WiHo workers can't just get the low price cohabitants typically get: it has to be the price paid in marriage if the WiHo worker wants it."

When traditional gender roles prevail women tend to be the WiHo workers. Women's willingness to form a couple is based on their supply of WiHo and is a function of their willingness to work in WiHo at different 'prices'.[7] Those benefiting from WiHo and possibly 'paying' for it are then men. With traditional gender roles these are men. They could pay with love, but most women also expect a material compensation for the WiHo they supply. Men thus have a demand for WiHo reflecting their willingness to pay for women's WiHo in the form of transfers of some of their higher personal income or of goods consumed by the WiHo-worker. 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. These conflicting interests possibly lead to bargaining within traditional heterosexual couples. Equilibrium prices

[5] WiHo may also benefit the self. More on this Beckerian theory of marriage can be found in Grossbard (forthcoming).
[6] Choo and Siow (2006) also have marriage markets establishing prices for men and women interested in marriage.
[7] Men may also supply WiHo to their wife, especially if they earn less than she does.

for WiHo are established at the intersection of aggregate demand and supply in each marriage market defined as a market for WiHo.

In the context of traditional gender roles this implies that CLM laws make som men pay more for women's WiHo than they otherwise would. They can't just 'hire' women doing WiHo work in return for a cohabitation contract offering low legal protection. It has to be marriage whenever the WiHo worker wants it. This forces men to move up their downward-sloping demand for WiHo and ask for a smaller amount of WiHo at the higher WiHo price. At the same time a higher WiHo price is expected to increase women's willingness to enter marriage or cohabitation: they will move up on their upward-sloping supply of WiHo.

Figure 1 represents a market for WiHo supplied by women; *h* denotes WiHo; *y* denotes its price; Supply *S* is by women and demand *D* is by men. It is assumed that many women are sufficiently alike to be substitutable and the same is true for many men. CLM amounts to a minimum price of WiHo *ymin* set above the market-clearing *y*. At this higher *y* women are willing to work more at WiHo in marriage or cohabitation, as they move up their supply of WiHo. However, at the higher *ymin* men are less willing to obtain WiHo than they were at the market-clearing price. Men move up their demand curve as a result of both a price effect and an income effect.

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. The total number of women employed in WiHo is likely to shrink, implying

<u>Prediction 1</u>. Under CLM there will be less couple formation.

As long as monogamy prevails overall effects of CLM on couple formation of men and women are likely to be similar. However, if we examine the probability of couple formation for a more restricted group, we may find differences for men and women as they may not necessarily form couples with others sharing a particular characteristic.

CLM's effect on couple formation is likely to vary with a number of individual characteristics associated with different marriage market conditions and expected sensitivity to CLM laws.

First, CLM laws are expected to have more impact where traditional gender roles are more likely to be in effect. Where women are more likely to be paid for their WiHo, the analysis presented here is more likely to hold. A minimum *ymin* is more likely to cause drops in couple formation rates. In contrast, if heterosexual women don't get paid for their WiHo anyways, CLM will not have much of an effect on couple formation.

In the U.S. blacks and whites tend to participate in separate markets for marriage and cohabitation, interpreted here as WiHo markets. Most economic analyses of marriage in the U.S. are therefore performed solely for whites (e.g. Oreffice 2014) or for blacks (e.g. Hamilton et al. 2009) or samples are separated by ethnicity (e.g. Lichter et al. 1992). The WiHo model will apply more to whites than to blacks to the extent that white women are more likely to be paid for their WiHo than black women.

Time spent on chores is a possible indicator of WiHo time and there is a wider gap between the average amount of time white men and women spend on chores relative to the corresponding gap for blacks (Grossbard, Gimenez and Molina forthcoming). Husbands and partners of white wives also tend to earn more than husbands or partners of black women. Combined, these gender and race differentials suggest that white women are more likely to get paid for their WiHo than black women. In fact, Cherry (1998) has modeled marriage among blacks as involving women paying men. To the extent that the model presented above applies to whites better than to blacks it follows that:

<u>Prediction 2.</u> *Drops in couple formation associated with CLM are likely to be larger for whites than for blacks.*

Educated WiHo workers are likely to be more productive. This helps explain why relative to the non-educated, the educated are more likely to form couples (Brien and Sheran 2003). 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 affect the women with low education more than 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:

<u>Prediction 3.</u> *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.*

However, for men who are paying for WiHo, if they are more educated and their income is higher, by setting more claims to their future earnings CLM entails potentially more costs: they may have to share a higher income in case of separation, for example. Therefore

<u>Prediction 4.</u> *Drops in couple formation associated with CLM are likely to be larger for men with a college education than for men without that education.*

Predictions 3 and 4 regarding education apply to men as well as women, both as WiHo workers and as employers of WiHo. The more they are likely to be on the employer side, the more education is likely to discourage couple formation.

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 WiHo is expected to be lower when sex ratios are lower. In cultures where brideprice and/or dowry are observed, this means a shift from brideprice (men pay) to dowry (women pay). In that vein Francis (2011) found that when sex ratios were lower in Taiwan brideprice was observed less often and dowries were more common. In the U.S. 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 price *y* (and low sex ratios) relative to its effect for women in markets with high *y* and high sex ratios:

<u>Prediction 5.</u> *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 last distinction we make is between parents and those who don't have children. 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. However, if their *y* is high to start out with, they may not be affected by a minimum 'price' *y* implied by CLM.

Therefore a minimum *y* taking the form of CLM laws is most likely to affect couple formation among those who don't have children:

<u>Prediction 6.</u> *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 those who are parents with children present.*

This model can 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 (2014).

3. Data and Sample Means

We analyze micro data from the March Current Population Surveys[8] (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. Three states abolished CLM over the period covered by this data set: Idaho (1996), Georgia (1997), and Pennsylvania (2005). 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.

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 21.5% 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, 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.

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 cohabit in CLM states. The last panel indicates that in CLM states

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

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 even decreased in all three states that abolished 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 white 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. The graphs suggest that couple formation may have increased in Idaho and decreased slightly in Georgia and Pennsylvania after the abolition 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.

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:

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

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). 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, is the indicator for 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, full-time and part-time student status, two indicators for metropolitan residence, central city and outside central city with non-metropolitan and 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. 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; 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.

The full list of controls in vector X can be found in the Appendix where we show full estimates of equation (1) for the probability of being in couple. 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. Identification of a CLM effect arises through cross-state variation and variation over time as three states abolished 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 cohabit relative to staying single, the corresponding coefficients in the probit regressions for married and cohabiting will be positive. While our theory predicts that having a CLM option will discourage couple formation, it is an empirical issue to find out the actual effect of this law.

We estimate two versions of equation (1), Model 1, a basic model, and Model 2, a model where we replace time fixed effects with state-specific time trends and exclude 3 years after the abolition of the law by each state. The latter is done in order to relax the assumption of quick adjustment to changes in the law and quick couple formation (1996-99, 2005-07 are excluded). As an additional robustness test we also estimate the basic model with the same years excluded (but not state-specific time trends) and find that the results are similar to our Model 2, hence they are not presented.

We estimate most models for the entire sample including all ethnicities. In addition, we also present results for white respondents only. Ideally, we would have liked to also present separate results for other ethnic groups because most marriages are within the same race and there are reasons to believe that marriage market conditions differ for blacks, Hispanics, and whites. In our data, 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 our difference-in-difference analyses. For example, there are fewer than 30 Hispanic men and women in each of the transition states in most years. As for black respondents, there are fewer than 100 observations per gender-year cell in Pennsylvania, and virtually none in Idaho. Furthermore, our theoretical predictions that require creating subsamples categorized by ethnicity, education, presence of children, and residence in high or low sex ratio states can only be tested for whites.

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.

Results presented below include New Hampshire among CLM states, even though it only offers CLM in case of a partner’s death. Excluding this state does not significantly alter our results.

5. Results

Our principal analyses estimate probability of in-couple residence, marriage and cohabitation using the CPS micro data. We first discuss results for all respondents and white respondents. Then we present results for subgroups of whites so we can test most of the predictions derived in Section 2.

All ethnicities and all whites. Table 2 shows estimates of CLM effects on in-couple residence, not distinguishing between marriage and cohabitation. The estimates are based on two versions of equation (1), holding all other regressors constant at mean values. Model 1 includes all years. Model 2 excludes three years after the law change and replaces time and year fixed effects with state-specific time trends. 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 variable at their means. It is evident from a comparison of results on in-couple residence for the whole sample and for whites that the results for the whole sample are driven by those for whites; there are relatively few blacks and Hispanics in the states undergoing law changes.

All coefficients are negative, suggesting that CLM is associated with lower rates of in-couple residence, as was predicted. The coefficients for all men and all white men are statistically significant at the 5% level, regardless of the model used. CLM effects for all women and all white women are negative and statistically significant if we use model 2, excluding the years after the law change and thus allowing for a lag in adaptation to the law change. This suggests that relative to women men were quicker to adapt to the abolition of CLM. Men also responded more to CLM changes than women if we compare the size of the coefficients. In part this is because fewer men are in couple by age 35, our cutting point, than is the case with women.

Table 3 shows estimates of CLM effects in probit regressions of the probability that the respondent is married and that the respondent is cohabitating, if unmarried. For men, the coefficients of CLM in the marriage regressions are negative and highly significant according to Model 2, and the coefficients of CLM in the cohabitation regressions are negative and significant according to Model 1. This suggests that men's adaptation to the law change first took the form of changes in cohabitation and then changes in marriage. The coefficients of CLM for women in Table 3 are also negative, but only marginally significant in the regression of white women' probability of marriage.

Couple formation is the more interesting outcome from a policy standpoint than marriage or cohabitation. It has been shown that children benefit from growing up with two parents. Whether the parents are married or cohabiting has less impact on children. Therefore, and given that we got strong results on CLM effects on in-couple residence, the rest of this discussion focuses on in-couple residence.

Further results on CLM and in-couple residence. Due to data restrictions, we now focus on white respondents. We are not able to test Prediction 2, comparing blacks and whites. Prediction 3 implies that we should find larger effects of CLM on women's in-couple residence if they are less educated. It can be seen from Table 2 that the CLM coefficient is negative and significant in regressions for women regardless of whether we use model 1 or model 2. This is consistent with prediction 3 and suggests that the elimination of CLM led to more couple formation among white women without a college degree. We also find that according to model 2 men without a college degree have a lower probability of in-couple residence where CLM is available.

Prediction 4 stated that couple formation among men with a college degree would respond more to CLM changes than among men without a degree. We find a strong response of college-educated men to CLM law changes according to both model 1 and model 2. For men without college education we only found a significant CLM coefficient according to model 2. According to that model, the coefficient of CLM for college educated men is -.106 and for men without college -.056. We conclude that there is evidence for prediction 4, indicating that men who have a college degree have more to loose from CLM in case of divorce.

Prediction 5 stated that drops in women's couple formation would be larger in low sex ratio states than in high sex ratio states. It can be seen from Table 2 that the coefficient of CLM is negative in all regressions for both women and men in low sex ratio states, regardless of whether we use model 1 or model 2. In contrast, none of the CLM coefficients are significant in the high sex ratio states. This is consistent with what we argued: where sex ratios are low CLM has more of the potential to boost the price for women's WiHo. That we find this effect for both men and women makes sense given that monogamy prevails.

The final prediction 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 Table 2 that when no children are present CLM coefficients are negative for both men and women: for men according to both model 1 and model 2; for women only according to model 1. 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. However, we also find a large negative coefficient of CLM for women with children according to model 2. This makes sense in terms of our analysis to the extent that this captures young women with children from a previous relationship who may be getting a lower y for their WiHo in the marriage market, and the minimum y implied in CLM therefore has more impact.

Robustness checks: Our results suggest negative effects of CLM on couple formation. In order to ascertain that these results are not random, we conduct three additional robustness checks that are reported in Table 4. First, we run the multinomial logit model for men and women separately over three outcomes: married, cohabiting and single, and we compute log odds of being in a married or cohabiting relationship relative to being single. These equations include fixed time and state effects and are similar to Model 1 with three outcomes. The odds are negative for men and women, and are significant at the conventional level in regressions for men of all ethnicities and for white men. Living in a CLM state reduced the log-odds of being in couple relative to staying single while holding all other variables in the model constant.

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 transition states from the sample and randomly assign any three states to be CLM states till a random year between 1996 and 2006. We remove three years after the 'abolition' of the fake law and estimate the effect of CLM where CLM represents existing CLM states as well as fake transition states. 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'. These numbers are presented in Test 2. We find that the non-existent laws are significant in 16-24% of the simulations, with positive and negative effects equally likely. The non-rejection rate is higher than the 5% that can be conventionally attributed to randomness, and thus we conclude that this test is inconclusive. Some simulations possibly result in the CLM dummy picking the average differences between existing CLM and non-CLM states or coincide with other changes that may have occurred in randomly chosen states. We therefore design an additional test 3 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 test under Test 3 in the same table. 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 also obtain negative effects of CLM on the probability of being in a couple and being married. These effects are significant at 5% for men in regressions for the probability of being married, and are significant at 10% for the probability of being in a couple.

Other coefficients, based on the Appendix: 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 26.2%, 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 2.6%. Higher last year unemployment rate reduces the odds that the person is in a couple. Neither median state income nor the generosity of welfare payments have any significant impact on union formation apart from what was captured by state fixed effects. Demographic composition – the share of Hispanic and urban residents in the state – increase the odds that the respondent is in couple; the percentage of black or college educated residents do not affect this probability. Household formation has a concave age profile. Education increases the odds of being in a couple. Living in a metropolitan area reduces the odds of being in couple. Students of both genders are significantly less likely to be in a co-residing relationship than non-students, with large negative coefficients for full-time students. Men and women of all ethnic groups are less likely to be in a relationship than their white counterparts.

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 U.S. A difference-in-difference analysis was performed given that during the period examined three states abolished CLM. Results using CPS respondents under age 36 revealed that CLM reduces in-couple residence among both women and men, whereas it affects the probability of marriage mostly among men. CLM is also associated with a lower probability that unmarried men cohabit with a woman. This holds for all ethnicities and for whites.

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 education, sex ratio, and parental status. We predict and find that college-educated white 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 with college educated-women. Other predictions we find evidence for are that the couple formation effects of CLM are stronger for whites in states with low sex ratios (ratios of men to women) and in the case of childless respondents.

This implies that abolition of CLM in some states encouraged couple formation. Since overall couple formation rates have decreased in the U.S. 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 abolish 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 white women without a college education and on that of white men with a college education. As a result, the abolition of CLM has led to larger increases in couple formation by low education women (relative to those with high education) and by high education men (relative to those with a low education).

In this analysis we have assumed that the abolition of CLM is an exogeneous change. We realize that changes in legislation are not random: factors that have led to increases in in-couple residence rates may also have pushed states to abolish 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 the gender norms in a society 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. The same trend can also explain why CLM laws are becoming increasingly unpopular among voters and politicians: traditional men may have never liked these laws and younger more egalitarian-prone cohorts less favorable to CLM are replacing traditional women who benefit from the laws.

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.

References

Alesina, Alberto and Paola Giuliano. (2007) "Divorce, fertility and the value of marriage." Harvard Institue of Economics Research Discussion Paper No 2136, April.

Angrist, Joshua. (2002). "How Do Sex Ratios Affect Marriage and Labor Markets? Evidence from America's Second Generation," *Quarterly Journal of Economics* 117 (3): 997-1038.

Becker, Gary S. (1973). "A Theory of Marriage: Part I." *Journal of Political Economy* 81(4): 813-846.

Becker, Gary S. (1981). "A Treatise on the Family". Cambridge: Harvard University Press.

Bertrand, Marianne, Esther Duflo, and Sendhil Mullainathan (2004). "How Much Should We Trust Difference-in-Difference Estimates." *Quarterly Journal of Economics* 119 (1): 249-275.

Brien, Michael and Michelle Sheran. (2003). "The economics of marriage and household formation." In *Marriage and the Economy*, ed by Shoshana Grossbard-Shechtman. New York: Cambridge University Press.

Cherry, Robert. (1998). "Rational Choice and the Price of Marriage." *Feminist Economics* 4:27-49.

Choo, Eugene, and Aloysius Siow. (2006). "Who Marries Whom and Why." *Journal of Political Economy* 114(1): 175-201.

Ekert-Jaffe, Olivia and Shoshana Grossbard. (2008). "Does Community Property Discourage Unpartnered Births?" *European J of Political Economy* 24(1):25-40.

Francis, Andrew M. (2011). "Sex ratios and the red dragon: using the Chinese Communist Revolution to explore the effect of the sex ratio on women and children in Taiwan." *Journal of Population Econ* 24(3):813-37.

Friedberg, Leora. (1998). "Did Unilateral Divorce Raise Divorce Rates? Evidence from Panel Data." *The American Economic Review* 88(3):608-627.

Goodwin Paula Y., William D. Mosher, and Anjani Chandra. (2010) "Marriage and cohabitation in the United States: A statistical portrait based on Cycle 6 (2002) of the National Survey of Family Growth". National Center for Health Statistics. Vital Health Stat 23(28).

Grossbard, Shoshana. (forthcoming). *A Price Theory of Marriage*. Springer Publishers.

Grossbard, Shoshana, J. Ignacio Gimenez and J. Alberto Molina. (forthcoming). "Racial Intermarriage and Household Production." *Review of Behavioral Economics*.

Grossbard, Shoshana and Victoria Vernon. (forthcoming). "Common Law Marriage, labor supply and time use: a partial explanation for gender convergence in labor supply." *Research in Labor Economics*.

Grossbard-Shechtman, Amyra. (1984). "A Theory of Allocation of Time in Markets for Labor and Marriage." *Economic Journal* 94:863-82.

Guttierez, Emilio and Pablo Suarez Becerra. (2012). "The relationship between Civil Unions and fertility in France: Preliminary evidence." *Review of Economics of the Household* 10(1): 115-132.

Halla, Martin. (2013). “The Effect of Joint Custody on Family Outcomes.” *Journal of the European Economic Association* 11: 278–315.

Hamilton, Darrick, Arthur Goldsmith, and William Darity Jr. (2009). “Shedding ‘Light’ on Marriage: The Influence of Skin Shade on Marriage of Black Females.” *Journal of Economic Behavior and Organization* 72(1): 30-50.

Heer, David and Amyra Grossbard-Shechtman. (1981). "The Impact of the Female Marriage Squeeze and the Contraceptive Revolution on Sex Roles and the Women's Liberation Movement in the United States, 1960 to 1975." *Journal of Marriage and the Family,* 43(1): 49-65.

Hobcraft, John. (1998). Intergenerational and Life-Course Transmission of Social Exclusion, Influence of Child Poverty, Family Disruption, and Contact with Police CASE Paper 1, ESRC Centre for Analysis of Social Exclusion, London School of Economics.

Leturcq, Marion. (2011). “Competing marital contracts? The marriage after civil union in France, June http://paa2012.princeton.edu/papers/120916

Lichter, Daniel, Diance McLaughlin, George Kephart and David Landry. (1992). “Race and the Retreat from Marriage: A Shortage of Marriageable Men?” *American Sociological Review* 57 (6):781-799

Lind, Goran. (2008). *Common Law Marriage: a Legal Institution for Cohabitation*. New York: Oxford University Press.

McLanahan, Sara S. and Gary Sandefur. (1994). *Growing Up with a Single Parent, What Hurts, What Helps.* Cambridge, Massachusetts, Harvard University Press.

McLanahan, Sara and Wendy Sigle-Ruston. (2004). “Father Absence and Child Wellbeing: A Critical Review.” Pp.116-155 in The Future of the Family, D. Moynihan, L. Rainwater, and T. Smeeding (Eds.). New York: Russell Sage Foundation.

Mincy, Ronald and Helen Oliver. (2003). “Age, Race, and Children’s Living Arrangements: Implications for TANF Reauthorization.” Washington, D.C. The Urban Institute.

Mincy, Ronald, Shoshana Grossbard, and Chien-Chung Huang. (2005). “An Economic Analysis of Co-Parenting Choices: Single Parent, Visiting Father, Cohabitation, Marriage.” Working paper, http://ideas.repec.org/p/wpa/wuwpla/0505004.html

National Vital Statistics Reports. (2014) Volume 63 Number 2, May 29. http://www.cdc.gov/nchs/data/nvsr/nvsr63/nvsr63_02.pdf . Retrieved August 26,2014.

Oreffice, Sonia. (2014). “Season of birth and marital outcomes.” IZA working paper No 8348.

Peters, Elizabeth H. (1986). “Marriage and divorce: informational constraints and private contracting.” *American Economic Review* 76:671-78.

Rasul, Imran. (2006). “Marriage Markets and Divorce Laws.” *Journal of Law, Economics and Organization* 22(1): 30-69.

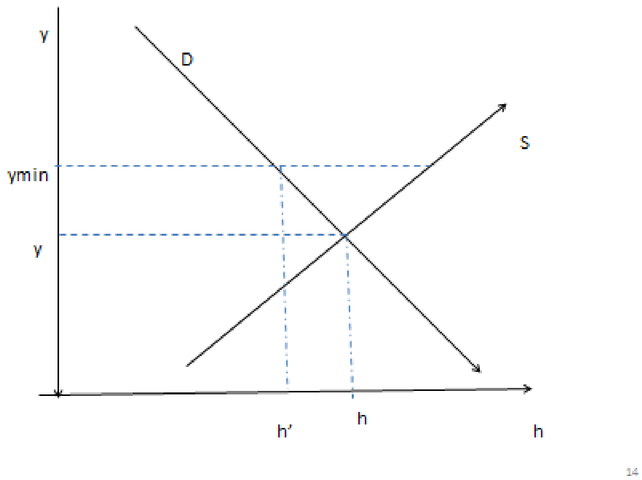
Sevilla-Sanz, Almudena. (2010). “Household division of labor and cross-country differences in household formation rates.” *Journal of Population Economics* 23:225–249

Sigle-Rushton, Wendy, John Hobcraft, and Kathleen Kiernan. (2005). “Parental Disruption and Adult Well-Being, a Cross Cohort Comparison.” *Demography* 423:427--446.

Stevenson, Betsey. (2007). "The Impact of Divorce Laws on Marriage-Specific Capital." *Journal of Labor Economics* 25:75-94.

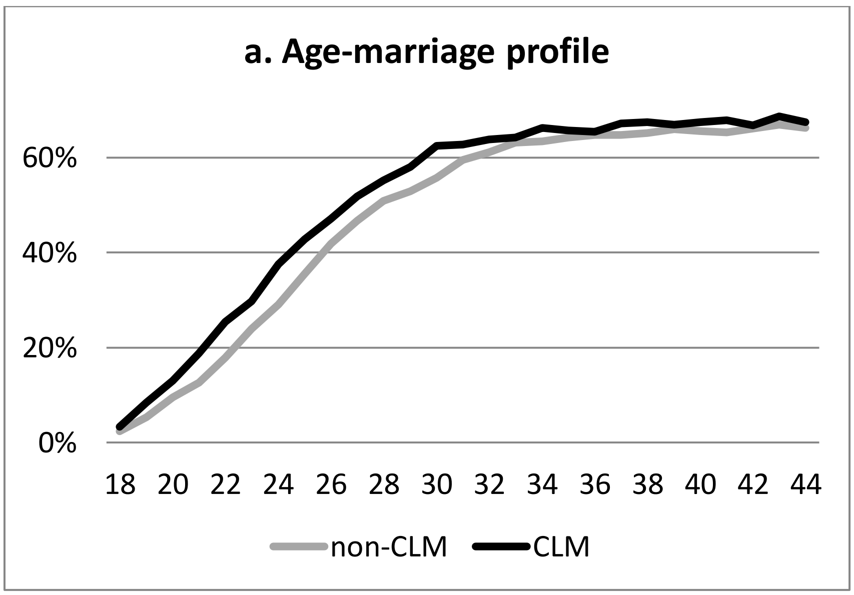
Wolfers, Justin. (2006). "Did Unilateral Divorce Laws Raise Divorce Rates? A Reconciliation and New Results." The *American Economic Review* 96(5):1802-1820.

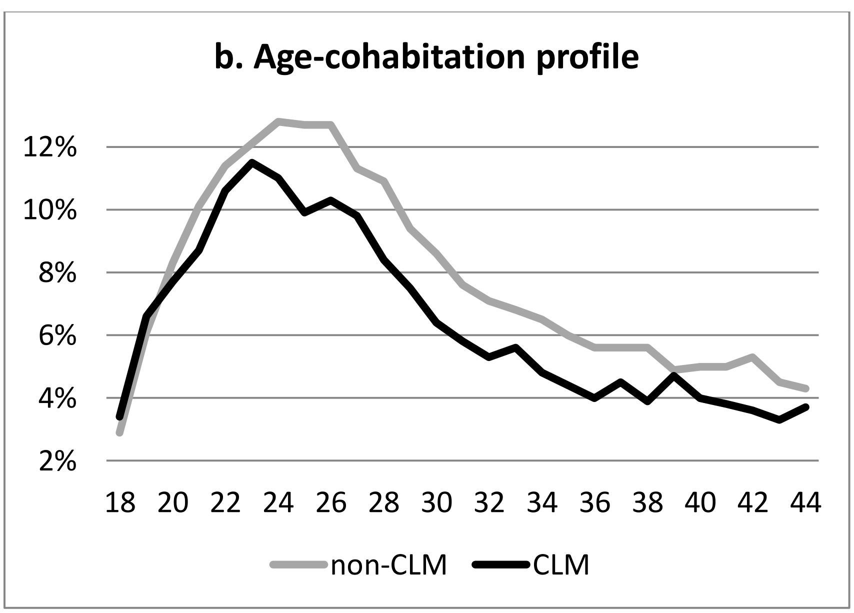
Figure 1. Market for women's Work-In-Household, comparing CLM states to non-CLM states

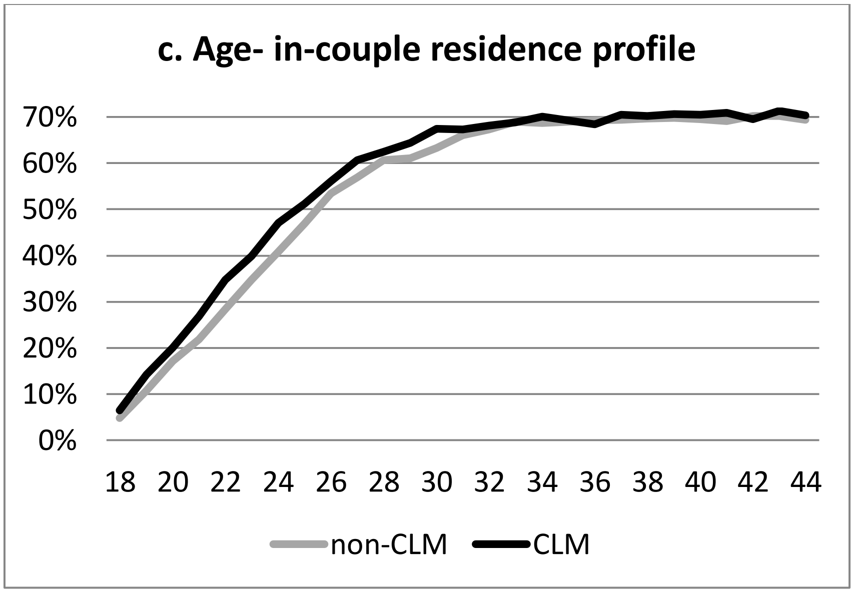


ymin: with CLM

Figure 2. Cohabitation and Marriage Profile by Age and Presence of CLM; All US-born Women aged 18-44 in CPS 1995-2011.

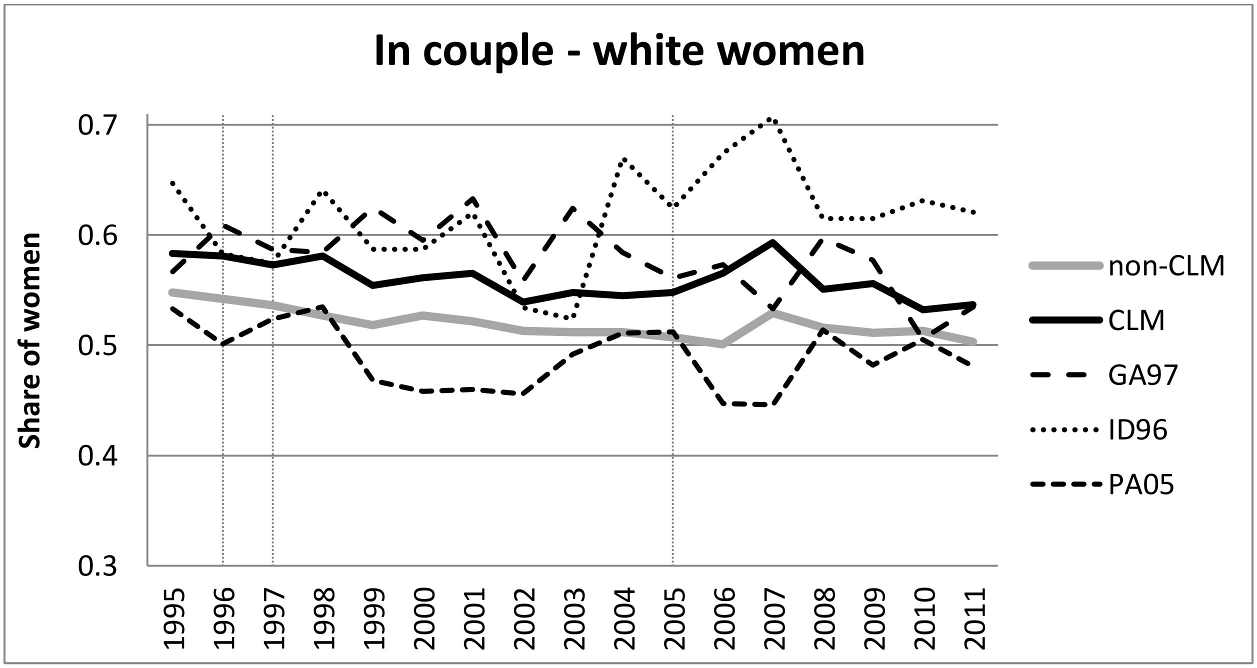


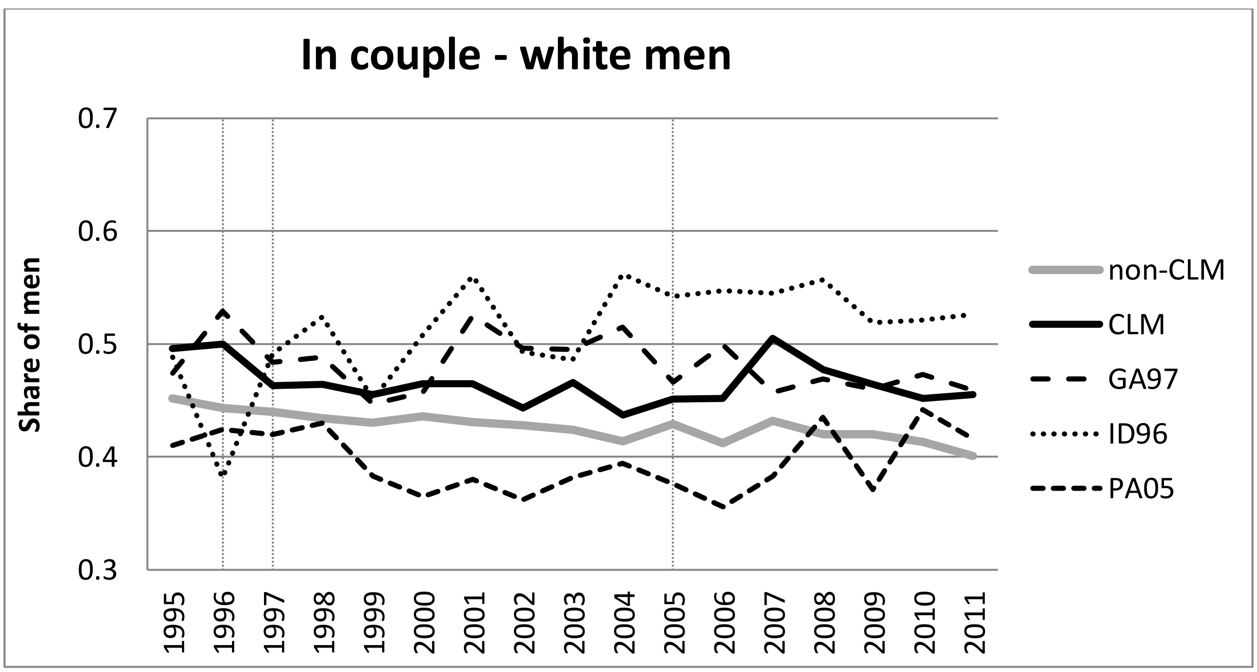




c. Age- in-couple residence profile
70%
60%
50%
40%
30%
20%
10%
0%
18 20 22 24 26 28 30 32 34 36 38 40 42 44
non-CLM
CLM

Figure 3. Evolution of In-Couple Residence Rates over Time, by presence of CLM.





Note: sample size for white women is 228,000, and for white men 213,514.

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

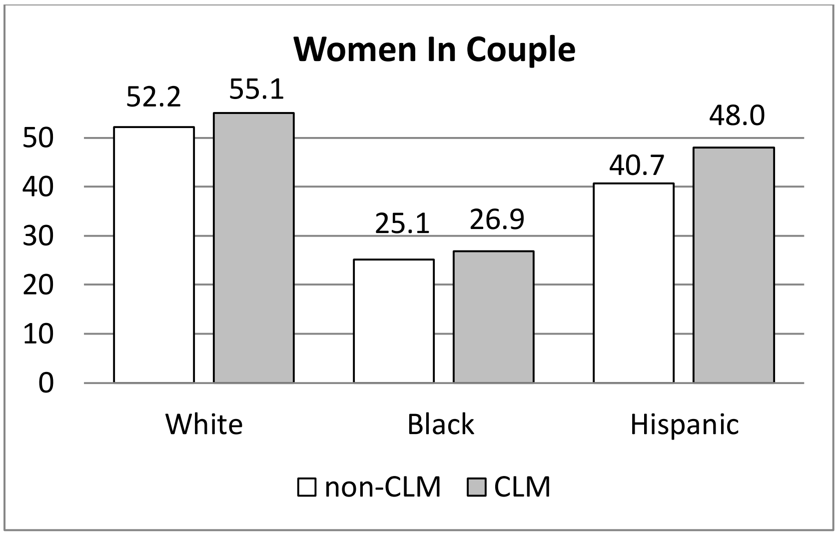


Table 1. Sample means. CPS 1995-2011. US born women and men, ages 18-35.

	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
Student	0.189	0.172	0.178	0.169
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*

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 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 obtained from the University of Kentucky Center for Poverty Research http://www.ukcpr.org/AvailableData.aspx

Table 2. CLM marginal effects in regressions for probability that a person is in a couple.

	WOMEN		MEN		WOMEN		MEN	
	Probit Model 1	Probit Model 2	Probit Model 1	Probit Model 2	N Model 1	N Model 2	N Model 1	N Model 2
	1	2	3	4	5	6	7	8
All ethnicities	-0.010	-0.036**	-0.023*	-0.062***	*321,917*	*202,309*	*292,376*	*183,543*
White	-0.010	-0.043**	-0.026*	-0.071***	*227,991*	*142,292*	*213,499*	*132,747*
White respondents:								
No college	-0.020*	-0.064***	-0.012	-0.056***	*169,811*	*105,167*	*167,049*	*103,372*
College	0.013	0.001	-0.074***	-0.106***	*58,804*	*37,125*	*46,894*	*29,375*
Low sex ratio areas	-0.066***	-0.128***	-0.081***	-0.138***	*48,167*	*32,716*	*47,652*	*32,666*
High sex ratio areas	0.003	-0.008	-0.008	-0.025	*179,832*	*109,576*	*165,859*	*100,081*
Children present	-0.001	-0.114***	0.006	-0.007	*112,674*	*70,678*	*72,094*	*45,195*
No children	-0.027*	-0.003	-0.031***	-0.058***	*115,938*	*71,614*	*141,837*	*87,472*

Notes: Model 1 is the basic model as explained under equation (1). Model 2 excludes 3 years after the abolition of the law in each state (1996-99, 2005-07 are excluded) and includes state-specific time trends. High sex ratio areas are those with sex ratio >1 and low are sex ratio <=1.

Table 3. CLM marginal effects in regressions for probability that a person is married or cohabiting.

	WOMEN		MEN		WOMEN		MEN	
	Probit Model 1	Probit Model 2	Probit Model 1	Probit Model 2	N Model 1	N Model 2	N Model 1	N Model 2
	1	2	3	4	5	6	7	8
Probability of 'Married' (among all respondents)								
All ethnicities	-0.005	-0.024	-0.013	-0.048***	*321,917*	*202,309*	*292,376*	*183,543*
White	-0.006	-0.026*	-0.012	-0.056***	*227,991*	*142,292*	*213,499*	*132,747*
Probability of 'Cohabiting' (among unmarried respondents)								
All ethnicities	-0.011	-0.006	-0.013*	-0.010	*190,655*	*120,822*	*183,608*	*117,517*
White	-0.011	-0.023	-0.015**	-0.022	*121,859*	*76,490*	*127,773*	*79,547*

Table 4. Robustness checks.

	In couple		Married		Cohabiting	
	women	men	women	men	women	men
	1	2	3	4	5	6
Test 1. Multinomial logit model, log odds relative to being single, Model 1.						
All ethnicities			-0.016	-0.109	-0.111	-0.190
			[0.032]	[0.048]**	[0.013]	[0.091]**
White			-0.029	-0.112	-0.115	-0.200
			[0.018]	[0.059]*	[0.019]	[0.090]**
Test 2. Placebo laws, Model 2.						
Number of 'significant' effects, negative/positive	10/12	8/8	12/8	11/10	11/13	11/12
Test 3. Data aggregated by state/year						
All ethnicities	-0.006	-0.021	-0.013	-0.024	0.008	0
	[0.012]	[0.012]*	[0.011]	[0.011]**	[0.013]	[0.010]
White	-0.008	-0.026	-0.015	-0.027	0.01	-0.004
	[0.015]	[0.014]*	[0.014]	[0.013]**	[0.016]	[0.012]

Notes: Multinomial logit model shows log odds of being married and cohabiting relative to single; estimated separately for men and women with state and year fixed effects. Coefficients for women in column (3) and (5) for test 1 are from the same regression, and so are coefficients for men in column (4) and (6).

The impact of 'placebo' laws are estimated using a model with 3 years after each law abolition removed in order to test the power of significant results produced by this model in Table 2.

For test 3 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. Coefficients shown are OLS estimates from these linear probability models.

Appendix. Estimates of probability of being in couple.

	All		White	
	WOMEN	MEN	WOMEN	MEN
CLM	-0.036	-0.062	-0.043	-0.071
	[0.016]**	[0.018]***	[0.020]**	[0.021]***
Sex ratios	0.262	0.187	0.232	0.024
	[0.034]***	[0.033]***	[0.050]***	[0.049]
Lag unemployment rate	-0.002	-0.004	-0.002	-0.004
	[0.001]*	[0.001]***	[0.002]	[0.002]***
Log median hhold income	-0.013	-0.001	-0.032	-0.081
	[0.046]	[0.041]	[0.053]	[0.047]*
Log welfare benefits	-0.009	-0.007	-0.015	-0.007
	[0.018]	[0.018]	[0.023]	[0.022]
Share of black pop	0.145	-0.114	-0.319	-0.482
	[0.570]	[0.570]	[0.745]	[0.757]
Share of Hispanic pop	0.162	0.089	0.358	0.127
	[0.089]*	[0.105]	[0.105]***	[0.178]
Share popul w/ college degree	-0.005	-0.003	-0.004	-0.003
	[0.003]	[0.003]	[0.004]	[0.004]
Share urban population	0.007	0.001	0.007	0.002
	[0.003]**	[0.003]	[0.004]*	[0.003]
Age	0.16	0.184	0.176	0.21
	[0.004]***	[0.005]***	[0.005]***	[0.005]***
Age-squared	-0.235	-0.263	-0.257	-0.304
	[0.008]***	[0.009]***	[0.009]***	[0.010]***
No high school diploma	-0.047	-0.005	-0.044	0
	[0.006]***	[0.006]	[0.008]***	[0.007]
Some college	-0.002	0.01	-0.013	0.004
	[0.004]	[0.004]***	[0.005]***	[0.005]
College degree	-0.008	-0.002	-0.02	-0.001
	[0.005]	[0.005]	[0.006]***	[0.006]
Graduate degree	0.048	0.078	0.033	0.079
	[0.008]***	[0.009]***	[0.009]***	[0.010]***
Metro: central	-0.126	-0.108	-0.142	-0.13
	[0.006]***	[0.005]***	[0.007]***	[0.006]***
Metro: outside	-0.05	-0.04	-0.059	-0.051
	[0.005]***	[0.004]***	[0.005]***	[0.005]***
Student part-time	-0.115	-0.116	-0.115	-0.125
	[0.010]***	[0.011]***	[0.013]***	[0.014]***
Student full-time	-0.287	-0.243	-0.305	-0.255
	[0.006]***	[0.006]***	[0.007]***	[0.007]***
Log unearned income	-0.011	0.003	-0.011	0.004
	[0.000]***	[0.001]***	[0.001]***	[0.001]***
Black	-0.243	-0.109		
	[0.006]***	[0.006]***		
Hispanic	-0.056	-0.019		
	[0.006]***	[0.006]***		
Asian	-0.102	-0.129		
	[0.013]***	[0.010]***		
Other ethnicity	-0.124	-0.041		
	[0.010]***	[0.011]***		
Time trend	-0.007	-0.003	-0.01	-0.004
	[0.002]***	[0.002]	[0.003]***	[0.002]**
State-specific time trends	yes	yes	yes	yes
Observations	*202,309*	*183,543*	*142,292*	*132,747*

Note: The table shows full estimates of Model 2 from Table 2, marginal effects from probit regressions estimated at the mean. 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.