Do Innovations in Birth Control Technology Increase the Welfare of Women? *

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Abstract

Birth control pills and legal abortions enable single women to participate in sexual activity with much lower risk of unwanted pregnancy or childbearing. The standard view is that these innovations increase opportunities for women and therefore increase their welfare. An alternative view is that these innovations cause more single women to participate in sexual activities, reducing the bargaining power of women in marriage. This paper investigates both views in an integrated model. When the predictions of the model on the average age of first marriage and the number of never married individuals are consistent with behavior in the US since the seventies, the standard view is correct, that improvements in birth control technology increase the welfare of women. Across states and years variation in access to legal abortions and birth control pills show that easier access to reproductive control technologies did not increase, and may have reduced, the out-of-wedlock birth rate.

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In the United States, legal abortions and birth control pills became widely available to single women in the seventies. How did these innovations in birth control technology affect the welfare of American women?¹

The conventional wisdom among economists is that these innovations increased the choices available to women and therefore enhanced their welfare (E.g. Cook, et. al. (1999), Kane and Staiger (1996), Lundberg and Plotnick (1995)). Goldin and Katz (2000) provide a nice application of this reasoning. They argue that when birth control pills became widely available to single women, they could be sexually active and not have to worry about unwanted pregnancy. A cost of postponing marriage, either sexual abstinence or sexual activity and the threat of unwanted pregnancy, was removed. Single women became more likely to postpone marriage and accumulate more human capital. Goldin and Katz show that the rise in the entry rates of women into graduate and professional schools in the seventies coincided with the increasing availability of the pill in the United States.

In their provocative paper, Akerlof, Yellin and Katz (1996; hereafter AYK) provide a less optimistic view of these innovations in birth control technology. They argue that the availability of legal abortions reduced the bargaining power of women in marriage. Without legal abortion, men who wanted to have sex with their partners had to offer to marry their partners. With legal abortions, many women became willing to be sexually active without demanding marriage in return. The increase in the supply of sexually active single women

¹Marks (2001), Solinger (1998), and Tone (1997, 2001) provide social histories of the development and impact of these technologies.
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reduced the willingness of men to marry. Thus women who wanted to marry suffered a loss in bargaining power. AYK argue that this reasoning is partly responsible for the simultaneous rise in both the abortion rate and the rate of single motherhood. This reasoning can also potentially explain the coincidence in the rise in human capital accumulation by women and the increasing availability of the pill. In this view, women accumulated more human capital when the pill became available because they expected less rents from marriage. Although AYK’s insight is likely to be more general, they employ a random search model of the marriage market with non-transferable utilities and men who never prefer to marry.

The difference in welfare implications for women between the two views is striking. The conventional view emphasizes the welfare gains from an individual woman’s perspective. AYK take an equilibrium perspective and argue that marriage market considerations may overturn the conventional wisdom.

This paper provides a model of the marriage market which nests both the conventional view and AYK’s equilibrium considerations. The model assumes that men value marriage, allows transfers between men and women, and dispenses with search frictions. The objective of the model is to study the circumstances in which the conventional wisdom is valid and when it is not. Predictions of the model will be compared with observed marital behavior to see which view is empirically relevant.

My analysis shows that the relative supply of marriageable men to marriageable women is critical to the welfare calculation. When women are relative scarce in the marriage market, innovations in birth control technologies improve the welfare of women. A major rationale
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for marriage is to have children. In general, because women are fecund for a shorter period of their lives than men, marriageable women are relatively scarce. Thus in the “normal” case, the conventional view is valid, that innovations in birth control technology improve the welfare of women.

When marriageable men are very scarce relative to marriageable women, perhaps as in the environment discussed by Wilson (1987), the alternative view applies, that innovations in birth control technology decrease the welfare of married women. If the discount rate is low enough, all women, independent of their gains from cohabitation, may be made worse off. If marriageable men are relatively scarce, but not extremely so, the welfare of married women are unaffected by the innovations. So from an analytic perspective, AYK’s assumption that marriageable men are extremely scarce is necessary to generate the alternative view. Search frictions and non-transferable utilities are unnecessary.

Depending on relative scarcity, the model generates different predictions about gender differences in the average ages of first marriage and the fractions of never married individuals in the population. These predictions also vary with the state of birth control technology. I compare these predictions with the observed behavior by whites and blacks from the seventies to nineties. Since the inter-racial marriage rate is extremely low, blacks and whites essentially participate in two different marriage markets. The paper shows that the observed marital behavior of whites conformed to the “normal” case. On the other hand, Wilson and other observers argue that there is a relative scarcity of marriageable black men compared with marriageable white men (E.g. Tucker and Mitchell-Kernan (1995)). This scarcity is necessary
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but not sufficient for the alternative equilibrium to be valid. I show that the observed marital behavior of blacks shows that black marriageable men are not sufficiently scarce relative to black marriageable women to negate the welfare enhancing effects of innovations in birth control technology for black women.

I also investigate whether easier access to reproductive control technologies increased the out-of-wedlock childbearing rate. I use across states and years variation in access to legal abortions and the earliest age at which a woman may acquire contraceptive services without parental consent to see how these factors affect the out-of-wedlock childbearing rate in the seventies. Both these variables have been shown to affect mean childbearing, marital and sexual behavior of women in the seventies (Gruber, Levine and Staiger (1999), Goldin and Katz). Using these variables, I show that easier access to reproductive control technologies did not increase, and may have decreased, the out-of-wedlock childbearing rate. Thus although there was an increase in aggregate abortions and out-of-wedlock childbearing in the seventies, it is likely that this coincidence was spurious.

Based on the evidence in this paper, innovations in birth control technology in the seventies benefited both white and black women. My conclusions apply to the aggregate marriage market in the US. It does not rule out the possibility that there are some local marriage markets in which marriageable men may be very scarce and the AYK effect is valid. While my conclusion differs from AYK, their concern for marriage market effects is well placed.

Aside from the above cited papers, the paper also builds on two literature. In his path breaking article, Trivers (1972) started the literature on differential fecundity and gender
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The two views studied in this paper were anticipated by the popular debate over the desirability of premarital sex. In that debate, the conventional view of economists expressed here is the modern popular view. As Posner (1992) argues, “There is no good reason to deter premarital sex, a generally harmless source of pleasure and for some people an important stage of marital search”. The alternative view expressed here is the traditional popular view. In her survey on courtship in early twentieth century America, Bailey (1988) writes: “For their collective good, women must continue to define their value in terms of virtue (sexual scarcity). Therefore, the woman who rejected “virtue” was not gaining sexual power or equality with men but was breaking the sexual “trust” and so threatening the precarious position of woman in society”. Kass (1997) specifically links the availability of effective female birth control techniques to the breakdown of traditional values and bemoans its consequence
for women: “Once female modesty became a first casualty of the sexual revolution, even women eager for marriage lost their greatest power to hold and to discipline their prospective mates”. The analysis in this paper may be informative in the earlier debate.

Finally, the theoretical argument that innovations may sometimes generate perverse welfare effects is well known from the international trade literature (E.g. Krugman and Obstfeld (2000), Chapter 5). As in that literature, the interest here is whether these perverse effects are empirically relevant.

1 The Model

In every period, an equal number of males and females enters the marriage market. We normalize the number of entrants of each gender to 1. We denote gender by subscript \( g \) which is either \( m \) for male or \( f \) for female. An individual’s participation in the marriage market is divided into two phases. In the young phase, some individuals value cohabitation more than marriage. In the mature phase, each individual values marriage more than cohabitation. The value of being single in each period is normalized to zero. In each period, an individual has to decide whether to remain single, cohabitate without marriage or to marry.

All young adults will eventually exit the young phase by becoming mature adults. Let \( p \) be the probability that a young adult will remain as a young adult in the next period. \( (1 - p) \) is the probability that a young adult will exit the young phase.

In each period, the total number of young adults of each gender is

\[
y = \sum_{i=0}^{\infty} p^i = \frac{1}{1 - p}
\]  

(1)
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A young adult exits the young phase either by death or into the mature phase. Let $\sigma$ be the per period survival probability for a young or mature adult. Conditional on surviving in the next period, let $r$ be the probability that a young adult in the current period will remain a young adult in the next period. Then $1 - r$ is the probability that a surviving young adult will become a mature adult in the next period. So

$$p = \sigma r$$

Mature adults exit the marriage market by death or becoming single. Let $\pi_g$ be the probability that a current mature adult will remain a mature adult in the next period. Conditional on surviving into the next period, a mature adult may leave the marriage market by becoming single in the next period. Exit as a single is irreversible. Let $\gamma_g$ be the per period probability that a surviving mature adult of gender $g$ will remain in the marriage market in the next period. I will discuss the rationales for gender differences in the exit rates later. Then

$$\pi_g = \sigma \gamma_g$$

In each period, the total number of remaining mature males is

$$n_m = y\sigma (1 - r) \sum_{i=0}^{\infty} \pi_m^i = \frac{y(\sigma - p)}{1 - \pi_m}$$

(2)

$y\sigma (1 - r)$ is the number of new mature males in a current period. The stock of remaining mature males in a period that is generated by one entrant per period, for every previous period, is $\sum_{i=0}^{\infty} \pi_m^i$. 

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Similarly, in each period, the total number of remaining mature females is

$$n_f = \frac{y(\sigma - p)}{1 - \pi_f}$$  \hspace{1cm} (3)

From hereon, remaining mature adults are referred to as mature adults. Using (1), (2) and (3),

$$n_m - n_f = \frac{(\pi_m - \pi_f)y(\sigma - p)}{(1 - \pi_m)(1 - \pi_f)}$$  \hspace{1cm} (4)

In any period, the value of being single is normalized to zero. In the young stage, men and women differ in their attitudes towards cohabitation. Let the range of per period intrinsic values from cohabitation, $\theta$, for men range uniformly from 0 to 1. Let the range of per period intrinsic values from cohabitation, $\lambda$, for women range uniformly from 0 to $\hat{\lambda}$. $0 < \hat{\lambda} < 1$.

In general, young women value cohabitation less than young men because of the cost of unintended pregnancy and childbearing. Let $\mu > 0$ be the per period intrinsic value of marriage for every young adult. $\hat{\lambda} > \mu$. Some young men and women prefer to cohabitate rather than marry.

Let $\Gamma$ be the per period intrinsic value of marriage for every mature adult where $\Gamma + \mu > 1 + \hat{\lambda}$. That is, the total per period value of a marriage between a mature adult and a young adult exceeds the per period gain from any cohabitating young couple. A rationale for this large value of marriage for mature adults is that they are wealthier than young adults. Let the value of cohabitation for a mature adult be the same as when he or she was young. In the transferable framework used here, since the largest value of cohabitation for a person is 1, all mature adults will marry.\footnote{This claim was checked directly after the equilibrium transfers were calculated.} So the distinction between mature and young adults is
that mature adults can afford to and want to marry. On average, mature adults are older than young adults. However, there will be some old adults who are not mature.

Cohabitants or spouses will immediately reenter the marriage market when their partners leave the marriage market. Individuals in the marriage market do not discriminate among other individuals of the opposite gender based on their maturity or marital history.\(^3\) Since individuals in marriage do not discriminate between young or mature spouses, let \(t\) be the per period equilibrium transfer that men make to women in marriage. A mature married man will obtain \(\Gamma - t\) per period. A mature married woman will receive \(\Gamma + t\) per period. A young married man will obtain \(\mu - t\) per period. A young married woman will obtain \(\mu + t\) per period. Cohabitants are all young individuals. Let \(k\) be the equilibrium per period transfer that men make to women in cohabitation. A man with a intrinsic value of cohabitation of \(\theta\) will obtain \(\theta - k\) per period from cohabitation. A woman with a intrinsic value of cohabitation \(\lambda\) will obtain \(\lambda + k\) per period from cohabitation.

Finally, individuals who leave the marriage market obtain no further payoff.\(^4\)

2 Excess supply of mature men

Let \(\pi_m > \pi_f\). Then \(n_m - n_f > 0\), and there are more mature men than mature women in the marriage market. I consider this the standard case because one of the main reasons for marriage is to have children. Menopausal women will not marry to have children whereas men of the same age cohort may still marry to have children. Thus women are more likely

\(^3\) Clearly an analytic convenience.
\(^4\) Death should receive lower payoff than being single but it does not affect behavior and so is ignored.
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than men to exit the marriage market earlier.

Empirical evidence that women are more likely to leave the marriage market earlier than men are provided by Chamie and Nsuly (1981). They show that for 47 countries for which there is data, the remarriage rate of divorced men is much higher than that of divorced women. In order to fit the aggregate marital behavior of 18th century Quebeckers, Hamilton and Siow estimated the extra exit rate of women at slightly more than 1% per annum. Integrating much of the previous anthropological literature, Siow shows that gender differences in both marital and labor supply behavior are consistent with a model in which fecund women are relatively scarce in the marriage market.

When \( n_m - n_f > 0 \), some mature men have to marry young women. All mature men will succeed in marrying because \( \Gamma + \mu > 1 + \lambda \) and they will always out bid younger men for partners. Let

\[
Z = \frac{n_m - n_f}{y} = \frac{(\pi_m - \pi_f)(\sigma - p)}{(1 - \pi_m)(1 - \pi_f)}
\]  

(5)

While \( Z \) may be larger than one, this case is not interesting because mature men will marry all of the women and all young men will be single. There will be no cohabitation. Since we want to study a society with cohabitation, let \( Z \) be a fraction. \( Z \) is the fraction of young women who will marry mature men.

Let \( \lambda^i \) be the intrinsic value of cohabitation for the marginal young woman who is indifferent between marriage and cohabitation. Then \( \frac{\lambda^i}{\lambda} \) is the fraction of young women who will marry. Since \( Z \) fraction of young women will marry mature men, \( \frac{\lambda^i}{\lambda} \geq Z \).

Let \( \theta^i \) be the intrinsic value of cohabitation for the marginal young male who is just
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indifferent between cohabiting and remaining single. Equating the per period return to
being single or cohabitating for the marginal young male gives the equilibrium value of $k$, $k^i$

$$\theta^i - k^i = 0$$ (6)

The marginal woman to marry, with intrinsic value of cohabitation $\lambda^i$, is indifferent
between marriage and cohabitation. The marginal young woman is indifferent when the
equilibrium value of $t$, $t^i$ satisfies

$$\mu + t^i = \lambda^i + k^i$$ (7)

Using (7), the per period net gain from marriage for a young man is

$$\mu - t^i = 2\mu - k^i - \lambda^i$$ (8)

The fraction of young men who cohabitate is $1 - \theta^i$. The fraction of young women who
cohabitate is $1 - \frac{\lambda^i}{\lambda}$. Since the fraction of male cohabitants must be equal to the fraction of
female cohabitants,

$$\theta^i = \frac{\lambda^i}{\lambda}$$ (9)

There are two cases to consider. In case A, no young man marries. Let the intrinsic
value of cohabitation for the marginal young woman be $\lambda^A$. Then $\lambda^A = Z\lambda$. In case B, some
young men marry. Let the intrinsic value of cohabitation for the marginal young woman be
$\lambda^B$. In this case, $\lambda^B > Z\lambda$. 

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3 Case A: No young man marries

In this case, young women only marry mature men and so

\[ \lambda^A = Z \hat{\lambda} \]  
(10)

Let i equal to A. (6), (7), (9) and (10) consist of four equations with four unknowns, \( \theta^A \), \( \lambda^A \), \( t^A \) and \( k^A \). Solving these equations give:

\[ \lambda^A = Z \hat{\lambda} \]  
(11)

\[ \theta^A = Z \]

\[ k^A = Z \]

\[ t^A = Z(1 + \hat{\lambda}) - \mu \]

Using (11), the per period net gain from marriage for a young woman is

\[ \mu + t^A = Z(1 + \hat{\lambda}) \]

Using (11) and (8), the per period net gain from marriage for a young man is

\[ \mu - t^A = 2\mu - Z(1 + \hat{\lambda}) \]

(12)

Since no young man marries, this net gain must be non-positive which implies

\[ \frac{\mu}{.5 \ast (1 + \hat{\lambda})} \leq Z \]

(13)

Thus the weak inequality (13) must hold for case A to apply. The denominator in (13) may be interpreted as the expected per period gain from cohabitation for a randomly chosen
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young adult. Then (13) implies that case A holds when $Z$, the fraction of young women who will marry mature men, is weakly larger than the ratio of the per period return to marriage for a young adult relative to the expected per period gain from cohabitation for a randomly chosen young adult.

The fraction of young men who will marry is

$$h^A_m = 0$$

The fraction of young women who will marry is

$$h^A_f = Z$$

The fraction of young adults who cohabit is

$$x^A = 1 - Z$$

In this paper, young adults are often indifferent between marriage and remaining single. When this is this case, I assume that young adults who marry, remain married until they become mature. In other words, young adults who are indifferent between marriage and remaining single do not rotate between being single and married.

I will now calculate the average age of first marriage for men. No young man marry. All mature men marry. So men marry as soon as they become mature. Consider a cohort of new young men of total size one. No one will marry at age one. $(\sigma - p)$ of this cohort will reach maturity at age two and then marry. $p(\sigma - p)$ of the cohort will reach maturity at age three and then marry. $p^{i-2}(\sigma - p)$ of the cohort will reach maturity at age $i$ and then marry.
Thus the average age of first marriage for this cohort of men is:

\[ a_m^A = \frac{\sum_{i=2}^{\infty} ip^{i-2}(\sigma - p)}{\sum_{i=2}^{\infty} p^{i-2}(\sigma - p)} = 1 + \frac{1}{1 - p} > 2 \]

Now consider a cohort of new young women of total size one. \( Z \) of them will marry at age one. The rest, \((1 - Z)\), will cohabit. They will continue to cohabit until they reach maturity. When they become mature, they will marry. Thus the average age of first marriage for a new cohort of women is:

\[ a_f^A = \frac{Z + (1 - Z) \sum_{i=2}^{\infty} ip^{i-2}(\sigma - p)}{Z + (1 - Z) \sum_{i=2}^{\infty} p^{i-2}(\sigma - p)} = \frac{Z(1 - p) + (1 - Z)(\sigma - p)a_m^A}{Z(1 - p) + (1 - Z)(\sigma - p)} < a_m^A \]

So in case \( A \), the average age of first marriage is lower for women than men.

The number of never married males in the population may be calculated as follows. Only young men never marry. So the number of never married males in the population is

\[ q_m^A = y = \frac{1}{1 - p} \quad (14) \]

\( Z \) number of age one females marry. \((1 - Z)\) of them do not. Young women who do not marry in the first period remain unmarried until they reach maturity. Following the derivation for never married men, the total number of never married women in the population is then

\[ q_f^A = \frac{1 - Z}{1 - p} \quad (15) \]

Comparing (14) and (15),

\[ q_m^A - q_f^A > 0 \]

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That is, there are more never married men than never married women in the population.

The average intrinsic value of cohabitation for cohabitating women is

$$\lambda^A = \frac{(1 + Z)\hat{\lambda}}{2}$$

The average intrinsic value of cohabitation for cohabitating men is

$$\theta^A = \frac{(1 + Z)}{2}$$

The average per period net payoff for cohabitating women is

$$\lambda^A + k^A = \frac{(1 + Z)\hat{\lambda}}{2} + Z$$

The average per period net payoff for cohabitating men is

$$\theta^A - k^A = \frac{(1 - Z)}{2}$$

4 Case B: Some young men marry

In this case, some young women marry young men and so

$$\lambda^B > Z\hat{\lambda} \quad (16)$$

Since some young men are marrying and other young men are single, the per period net gain from marriage for a young man must be zero. So

$$\mu - t^B = 0 \quad (17)$$
Let $i$ equal to $B$. (6), (7), (9) and (17) consist of four equations with four unknowns, $\theta^B$, $\lambda^B$, $t^B$ and $k^B$. Solving these equations give:

\[
\begin{align*}
\lambda^B &= \frac{2\mu}{1 + \lambda} \\
\theta^B &= \frac{2\mu}{1 + \lambda} \\
k^B &= \frac{2\mu}{1 + \lambda} \\
t^B &= \mu
\end{align*}
\]

Substituting the value of $\lambda^B$ from (18) into (16) gives:

\[
\frac{\mu}{.5(1 + \lambda)} > Z
\]

When (19) holds, case $B$ applies. Comparing (13) and (19), the parameter space of the model is partitioned as follows. When (19) applies, case $B$ follows. When $\frac{\mu}{.5(1 + \lambda)} = Z$, $\lambda^B = \lambda^A$, $\theta^B = \theta^A$, $k^B = k^A$, and $t^B = t^A$. When (13) holds, case $A$ applies. The economic intuition is straightforward. Holding $Z$ constant, when the gains from marriage for young adults is large relative to the expected gains from cohabitation, some young men will marry in equilibrium. Otherwise, young men will only cohabitate with young women.

Using the equilibrium value of $\lambda^B$, the fraction of young men who marry is

\[
h_m^B = \frac{\lambda^B}{\lambda} - Z = \frac{2\mu}{(1 + \lambda)} - Z
\]

The fraction of young women who marry is

\[
h_f^B = \frac{\lambda^B}{\lambda} = \frac{2\mu}{(1 + \lambda)}
\]
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Since the fraction of young men who cohabitate is equal to the fraction of young women who also do so, the fraction of young adults who cohabitate is

\[ x^B = 1 - \frac{\lambda^B}{\lambda} = 1 - \frac{2\mu}{(1 + \lambda)} \]  

(22)

The fraction of young men who marry in age one is \( h^B_m \). The rest of the age cohort, \( 1 - h^B_m \), will remain single or cohabitate. When they become mature, they will marry. The average age of first marriage for a new cohort of men is:

\[ a^B_m = \frac{h^B_m(1 - p) + (1 - h^B_m)(\sigma - p)a^A_m}{h^B_m(1 - p) + (1 - h^B_m)(\sigma - p)} \]  

(23)

The fraction of young women who marry in age one is \( h^B_f \). The rest of the age cohort will marry when they become mature. So the average age of first marriage for a new cohort of women is:

\[ a^B_f = \frac{h^B_f(1 - p) + (1 - h^B_f)(\sigma - p)a^A_m}{h^B_f(1 - p) + (1 - h^B_f)(\sigma - p)} \]  

(24)

Comparing (20) and (21), \( h^B_f > h^B_m \) and so the average age of first marriage is lower for women than men.

\( 1 - h^B_m \) age one men do not marry. Following the reasoning in the previous section, the number of never married men in the population is:

\[ q^B_m = \frac{1 - h^B_m}{1 - p} \]  

(25)

Likewise, the number of never married women in the population is

\[ q^B_f = \frac{1 - h^B_f}{1 - p} \]  

(26)
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By (20) and (21), \((1 - h_f^B) < (1 - h_m^B)\). So \(q_m^B > q_f^B\). There are more never married men than never married women.

In cases \(A\) and \(B\), the average age of first marriage for men is higher than that for women, and there are more never married men than never married women. These two results are standard in models of differential fecundity where there are more men than women in the marriage market (e.g. Siow (1998)). The results obtain when young men are squeezed out of marriage by mature (older and more wealthy) men.

The average intrinsic value from cohabitation and net per period payoff for cohabitating women are

\[
\Lambda^B = \frac{\lambda^B + \tilde{\lambda}}{2}
\]

\[
\Lambda^B + k^B = \mu + \frac{\mu}{1 + \lambda} + \frac{\tilde{\lambda}}{2}
\]

The average intrinsic value from cohabitation and net per period payoff for cohabitating men are:

\[
\theta^B = \frac{\mu}{1 + \lambda} + \frac{1}{2}
\]

\[
\theta^B - k^B = \frac{1}{2} - \frac{\mu}{1 + \lambda}
\]

5 Innovations in birth control technology

The diffusion of the use of birth control pills by single women and the availability of legal abortion in the 1970’s make unintended pregnancy and/or childbirth much less likely for
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sexually active women. If fear of unintended pregnancy or childbirth is a major drawback of cohabitation for single women, these innovations in birth control technology raise the gains from cohabitation. To study how these innovations affected the behavior and welfare of individuals in our society, we model the innovations as an increase in $\hat{\lambda}$. Since the intrinsic value of cohabitation for women is uniformly distributed between 0 and $\hat{\lambda}$, an increase in $\hat{\lambda}$ increases the intrinsic value of cohabitation probabilistically for every woman.

For case $A$,

$$\frac{\partial x^A}{\partial \lambda} = \frac{\partial h_f^A}{\partial \lambda} = 0$$

There is no change in the fraction of young adults who cohabitate or the fraction of young women who marry. So the ratio of the stock of cohabitating couples to the stock of married couples remains unchanged.

Since the average ages of first marriage for men and women are independent of $\hat{\lambda}$,

$$\frac{\partial a_m^A}{\partial \lambda} = \frac{\partial a_f^A}{\partial \lambda} = 0$$

That is, the average ages of first marriage for men and women do not change when $\hat{\lambda}$ changes.

Likewise

$$\frac{\partial q_m^A}{\partial \lambda} = \frac{\partial q_f^A}{\partial \lambda} = 0$$

The number of never married men and women do not change when $\hat{\lambda}$ changes.

Equilibrium transfers change as

$$\frac{\partial t^A}{\partial \lambda} = Z > 0$$
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\[ \frac{\partial k^A}{\partial \lambda} = 0 \]

Since equilibrium marital transfers increase, married women, young and mature, are unambiguously better off. Cohabitating women do not get higher transfers.

The welfare of young men is unaffected. Married (mature) men are worse off because they have to pay a higher equilibrium transfer.

For case B,

\[ \frac{\partial x^B}{\partial \lambda} = \frac{2\mu}{(1 + \lambda)^2} > 0 \]
\[ \frac{\partial h_f^B}{\partial \lambda} = \frac{-2\mu}{(1 + \lambda)^2} < 0 \]
\[ \frac{\partial h_m^B}{\partial \lambda} = \frac{-2\mu}{(1 + \lambda)^2} < 0 \]

The number of cohabitating couples increases. The number of young women who marry falls. The number of young men who marry falls. So the ratio of the stock of cohabitating couples to the stock of married couples rises.

And from (23) and (24) and \( a_g^A > a_g^B > 1 \),

\[ \frac{\partial a_m^B}{\partial \lambda} = \frac{-2\mu}{(1 + \lambda)^2}[\frac{a_m^B(1 - \sigma) + a_m^A(\sigma - p) - (1 - p)}{h_m^B(1 - p) + (1 - h_m^B)(\sigma - p)}] > 0 \]
\[ \frac{\partial a_f^B}{\partial \lambda} = \frac{-2\mu}{(1 + \lambda)^2}[\frac{a_f^B(1 - \sigma) + a_f^A(\sigma - p) - (1 - p)}{h_f^B(1 - p) + (1 - h_f^B)(\sigma - p)}] > 0 \]

Since \( h_f^B(1 - p) > (\sigma - p)h_m^B, h_f^B(1 - p) + (1 - h_f^B)(\sigma - p) > h_m^B(1 - p) + (1 - h_m^B)(\sigma - p) \).

Together with \( a_m^B > a_f^B, \frac{\partial a_m^B}{\partial \lambda} > \frac{\partial a_f^B}{\partial \lambda} \). So the average ages of first marriages for men and women rise with \( \hat{\lambda} \) with an increasing gender gap.
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From (25) and (26),

\[
\frac{\partial q^B_m}{\partial \lambda} = \frac{2\mu y}{(1 + \lambda)^2}
\]

\[
\frac{\partial q^B_f}{\partial \lambda} = \frac{2\mu y}{(1 + \lambda)^2} = \frac{\partial q^B_m}{\partial \lambda}
\]

The number of never married men increases at the same rate as the number of never married women when \( \hat{\lambda} \) increases.

Equilibrium transfers change as follows:

\[
\frac{\partial t^B}{\partial \lambda} = 0
\]

\[
\frac{\partial k^B}{\partial \lambda} = \frac{-2\mu}{(1 + \lambda)^2} < 0
\]

Married women are unaffected. Even though equilibrium transfers have decreased, the welfare of the average cohabitating woman increases by

\[
\frac{\partial (\lambda^B + k^B)}{\partial \hat{\lambda}} = \frac{1}{2} - \frac{\mu}{(1 + \lambda)^2} > 0
\]

The welfare of single and married men is unaffected. The welfare of the average cohabitating man increases by

\[
\frac{\partial (\hat{\theta}^B - k^B)}{\partial \lambda} = \frac{\mu}{(1 + \lambda)^2} > 0
\]

Summarizing the results for cases A and B, when there is an excess supply of mature men relative to mature women in the marriage market, the standard view that innovations in birth control technology improve the welfare of women is supported.
Do Innovations in Birth Control Technology Increase the Welfare of Women?

6 Wilson Equilibria

In some inner city black communities, Wilson argues that the supply of marriageable women exceeds the supply of marriageable men. He attributes this relative scarcity of marriageable men to their lack of jobs, their high incarceration rates and other problems. While some of unmarriageable men may be willing to marry, women are unwilling to marry them. From the perspective of this model, unmarriageable men have left the marriage market. Under Wilson’s assumption, $\pi_f > \pi_m$.

Let $\pi_f > \pi_m$. Then from (4), $n_f - n_m > 0$, and there are more mature women than mature men in the marriage market. Some mature women have to marry young men. All mature women will succeed in marrying because $\Gamma + \mu > 1 + \hat{\lambda}$ and they will always outbid younger women for partners. Let

$$\frac{n_f - n_m}{y} = \frac{(\pi_f - \pi_m)(\sigma - p)}{(1 - \pi_m)(1 - \pi_f)} = -Z$$

(27)

$-Z$ is the fraction of young men who will marry mature women. As before, we will assume $-Z$ is a fraction. The analysis of Wilson Equilibria is similar to the standard case except here mature women compete with young women for young men in the marriage market. As before, there are two cases to consider. In case $C$, some young women marry. Let the intrinsic value of cohabitation for the marginal young man be $\theta^C$. In this case, $\theta^C > -Z$. In case $D$, no young woman marries. Let the intrinsic value of cohabitation for the marginal young man, who is indifferent between marriage and cohabitation, be $\theta^D$. Then $\theta^D = -Z$.

\footnote{For empirical assessments of the Wilson hypothesis, see the edited volume by Tucker and Mitchell-Kernan and the references therein.}
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Let $i = C$ or $D$. Let $\lambda^i$ be the intrinsic value of cohabitation for the marginal young woman who is just indifferent between cohabiting and remaining single. Let $k^i$ be the equilibrium per period transfer from a cohabitating man to his partner, and let $t^i$ be the equilibrium per period transfer from a man to his wife.

Following the derivation in the earlier sections, Table (1) shows the equilibrium quantities for cases $C$ and $D$. The equilibrium values are as expected. In cases $C$ and $D$, some young women will be single. In case $C$, there are more young married men than young married women. In case $D$, no young woman marry; they only cohabitate or remain single. The average ages of first marriages is higher for women than men in both cases.

Since both $k^i$ and $t^i$ are negative in both of these cases, women transfer resources to men in marriage and cohabitation.

Table (1) can be used to calculate the changes in equilibrium values due to a marginal increase in $\dot{\lambda}$, where again an innovation in birth control technology is modelled as an increase in $\dot{\lambda}$. To ease comparisons, the changes in equilibrium values for all four cases are presented in Table (2).

In case $C$, some young women marry. When $\dot{\lambda}$ increases, the expected intrinsic value from cohabitation rises. The equilibrium transfer in cohabitation, $k^C$, falls since less transfers are needed to induce women to cohabitate. But if less $k^C$ is needed, men who cohabitate will be better off. More men will also be willing to cohabitate. Since the intrinsic value of cohabitation for women has also gone up, more women are also willing to cohabitate. The increase in cohabitating couples is made up by a decrease in young married couples. So in
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<table>
<thead>
<tr>
<th>Case</th>
<th>i</th>
<th>C</th>
<th>D</th>
</tr>
</thead>
<tbody>
<tr>
<td>Restrictions on ( Z )</td>
<td>( 0 &gt; Z &gt; \frac{2\mu}{1+\lambda} )</td>
<td>( 0 &gt; -\frac{2\mu}{1+\lambda} \geq Z )</td>
<td></td>
</tr>
<tr>
<td>( \lambda ) for marginal ( f )</td>
<td>( \lambda^i )</td>
<td>( \frac{2\mu_\lambda}{1+\lambda} )</td>
<td>(-Z\lambda)</td>
</tr>
<tr>
<td>( \theta ) for marginal ( m )</td>
<td>( \theta^i )</td>
<td>( \frac{2\mu}{1+\lambda} )</td>
<td>(-Z\lambda)</td>
</tr>
<tr>
<td>Cohabitation transfer from ( m ) to ( f )</td>
<td>( k^i )</td>
<td>(-\frac{2\mu_\lambda}{1+\lambda} )</td>
<td>(Z\lambda)</td>
</tr>
<tr>
<td>Marital transfer from ( m ) to ( f )</td>
<td>( t^i )</td>
<td>(-\mu )</td>
<td>(\mu + (1 + \lambda)Z)</td>
</tr>
<tr>
<td>Fraction of young ( m ) married</td>
<td>( h^i_m )</td>
<td>( \frac{2\mu}{(1+\lambda)} )</td>
<td>(-Z\lambda)</td>
</tr>
<tr>
<td>Fraction of young ( f ) married</td>
<td>( h^i_f )</td>
<td>( \frac{2\mu}{(1+\lambda)} + Z )</td>
<td>0</td>
</tr>
<tr>
<td>Fraction young adult cohabit</td>
<td>( x^i )</td>
<td>(1 - \frac{2\mu}{(1+\lambda)} )</td>
<td>(1 + Z\lambda)</td>
</tr>
<tr>
<td>Avg. age of first marriage for ( m )</td>
<td>( a^i_m )</td>
<td>( \frac{h^i_m (1-p_h) + (1-h^i_m) (\sigma - p_h) a^A_m}{h^i_m (1-p_h) + (1-h^i_m) (\sigma - p_h)} )</td>
<td>(-Z(1-p) + (1+Z)(\sigma - p_h) a^A_m )</td>
</tr>
<tr>
<td>Avg. age of first marriage for ( f )</td>
<td>( a^i_f )</td>
<td>( \frac{h^i_f (1-p_h) + (1-h^i_f) (\sigma - p_h) a^A_m}{h^i_f (1-p_h) + (1-h^i_f) (\sigma - p_h)} )</td>
<td>(a^A_m)</td>
</tr>
<tr>
<td>Number of never married ( m )</td>
<td>( q^i_m )</td>
<td>((1-h^C_m) y)</td>
<td>((1+Z) y)</td>
</tr>
<tr>
<td>Number of never married ( f )</td>
<td>( q^i_f )</td>
<td>((1-h^C_f) y)</td>
<td>(y)</td>
</tr>
<tr>
<td>Mean ( \lambda^i ) for cohabitating ( f )</td>
<td>( \lambda^i )</td>
<td>(\frac{\mu_\lambda}{1+\lambda} + \frac{\lambda}{2})</td>
<td>(\frac{(1-Z)\lambda}{2})</td>
</tr>
<tr>
<td>Mean ( \theta^i ) for cohabitating ( m )</td>
<td>( \theta^i )</td>
<td>(\frac{\mu}{1+\lambda} + \frac{1}{2})</td>
<td>(\frac{(1-Z)}{2})</td>
</tr>
<tr>
<td>Mean payoff for cohabitating ( f )</td>
<td>( \lambda^i + k^i )</td>
<td>(\frac{\lambda}{2} - \frac{\mu_\lambda}{1+\lambda})</td>
<td>(\frac{(1+Z)\lambda}{2})</td>
</tr>
<tr>
<td>Mean payoff for cohabitating ( m )</td>
<td>( \theta^i - k^i )</td>
<td>(\mu + \frac{\mu_\lambda}{1+\lambda} + \frac{1}{2})</td>
<td>(\frac{(1-Z)}{2} - Z\lambda)</td>
</tr>
</tbody>
</table>

Table 1: Equilibrium values
Do Innovations in Birth Control Technology Increase the Welfare of Women?

<table>
<thead>
<tr>
<th>Case</th>
<th>A</th>
<th>B</th>
<th>C</th>
<th>D</th>
</tr>
</thead>
<tbody>
<tr>
<td>Restrictions on Z</td>
<td>$Z \geq \frac{2\mu}{(1+\lambda)} &gt; 0$</td>
<td>$\frac{2\mu}{(1+\lambda)} &gt; Z &gt; 0$</td>
<td>$0 &gt; Z &gt; -\frac{2\mu}{(1+\lambda)}$</td>
<td>$0 &gt; -\frac{2\mu}{(1+\lambda)} \geq Z$</td>
</tr>
<tr>
<td>cohab. transfer $\frac{\partial k_i}{\partial \lambda}$</td>
<td>0</td>
<td>$-$</td>
<td>$-$</td>
<td>$-$</td>
</tr>
<tr>
<td>marr. transfer $\frac{\partial e_i}{\partial \lambda}$</td>
<td>$+$</td>
<td>0</td>
<td>0</td>
<td>$-$</td>
</tr>
<tr>
<td>no. young $m$ marry $\frac{\partial h_i}{\partial \lambda}$</td>
<td>0</td>
<td>$-$</td>
<td>$-$</td>
<td>0</td>
</tr>
<tr>
<td>no. young $w$ marry $\frac{\partial h_i}{\partial \lambda}$</td>
<td>0</td>
<td>$-$</td>
<td>$-$</td>
<td>0</td>
</tr>
<tr>
<td>no. young cohab. $\frac{\partial x_i}{\partial \lambda}$</td>
<td>0</td>
<td>$+$</td>
<td>$+$</td>
<td>0</td>
</tr>
<tr>
<td>age first marr. $m$ $\frac{\partial a_i}{\partial \lambda}$</td>
<td>0</td>
<td>$+$</td>
<td>$+$</td>
<td>0</td>
</tr>
<tr>
<td>age first marr. $f$ $\frac{\partial a_i}{\partial \lambda}$</td>
<td>0</td>
<td>$+$</td>
<td>$+$</td>
<td>0</td>
</tr>
<tr>
<td>No. of never married $m$ $\frac{\partial q_i}{\partial \lambda}$</td>
<td>0</td>
<td>$+$</td>
<td>$+$</td>
<td>0</td>
</tr>
<tr>
<td>No. of never married $f$ $\frac{\partial q_i}{\partial \lambda}$</td>
<td>0</td>
<td>$+$</td>
<td>$+$</td>
<td>0</td>
</tr>
<tr>
<td>Avg. value cohab. $f$ $\frac{\partial v_i}{\partial \lambda}$</td>
<td>$+$</td>
<td>$+$</td>
<td>$+$</td>
<td>$+$</td>
</tr>
<tr>
<td>avg. value cohab $m$ $\frac{\partial v_i}{\partial \lambda}$</td>
<td>0</td>
<td>$-$</td>
<td>$-$</td>
<td>0</td>
</tr>
<tr>
<td>net gain cohab $f$ $\frac{\partial (\lambda + k_i)}{\partial \lambda}$</td>
<td>$+$</td>
<td>$+$</td>
<td>$+$</td>
<td>$+$</td>
</tr>
<tr>
<td>net gain cohab $m$ $\frac{\partial (\theta - k_i)}{\partial \lambda}$</td>
<td>0</td>
<td>$+$</td>
<td>$+$</td>
<td>$+$</td>
</tr>
</tbody>
</table>

Table 2: Changes in Equilibrium Values
Do Innovations in Birth Control Technology Increase the Welfare of Women?

In case $C$, an increase in $\hat{\lambda}$ will result in more cohabitations relative to marriages. The average ages of first marriages increase at the same rate for both men and women. The number of never married women increase faster than the number of never married men.

Marital transfers are unaffected because the net payoff from marriage has to make young women indifferent between remaining single and marrying. An increase in $\hat{\lambda}$ does not change the intrinsic value of marriage. So the welfare of married women are unaffected by an improvement in birth control technology. The welfare of cohabitating women increase. Cohabitating men are better off.

In case $D$, no young woman marries. Young women either cohabitate or remain single. When $\hat{\lambda}$ increases, the expected intrinsic value from cohabitation rises. The equilibrium transfer in cohabitation, $k^D$, falls since less transfers are needed to induce women to cohabitate. But if less $k^D$ is needed, men who cohabitate will be better off. Less men will be willing to marry. In order to have the same number of young men to be willing to marry the mature women, the equilibrium transfers that men make to women in marriage, $t^D$, must also fall. Put another way, the net per period payoff for a married woman falls. The average intrinsic value from cohabitation, $\lambda^D$, rises more than the fall in $k^D$. So the average payoff per period for cohabitating women, $\lambda^D + k^D$, rises. So in case $D$, married women are made worse off when there is an improvement in birth control technology but cohabitating women are made better off. Single women are unaffected. Men are made better off. The explanation for the decline in the welfare of married women is in the spirit of AYK.

The fall in the welfare of married women can result in lower expected lifetime utility
for all women. The reason is that all surviving women eventually become mature and will marry. For a sufficiently low subjective discount rate, even young women who highly value cohabitation are made worse off in terms of discounted lifetime utility.

In case $D$, only equilibrium transfers adjust when $\lambda$ changes. The number of cohabitating couples and married couples do not change. The average ages of first marriages for men and women also do not change. The number of never married men and women do not change.

Comparing all four cases, $A$, $B$, $C$ and $D$, case $D$ is the only one in which the welfare of married women fall when there is an improvement in birth control technology.

7 Marital Behavior since the Seventies

As discussed in the introduction, legal abortions and birth control pills became widely available to single women in the United states in the seventies. I will analyze the marital behavior of whites and blacks separately because the two groups may be in different regimes.

The average age of white males who married for the first time in 1970 was 22.9. The average age of white females who married for the first time was 21.2 years old. The average ages of first marriage for men were higher than that for women for all years between 1968 to 1988. In 1970, 27% of white males age 15 and over never married. 22% of white females age 15 and over never married. In every decade between 1950 to 1990, the fraction of never married men exceeded the fraction of never married women. The gender differences

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6 Source for all age at first marriage data: Vital Statistics from National Center for Health Statistics.
8 Even though there are more women than men, the number of never married men also exceeded the number of never married women for every decade.
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in average age of first marriage and fraction of never married individuals are consistent with cases A and B where there is a relative scarcity of marriageable men. Cases C and D are rejected.

Figure 1 shows the evolution of the average ages of white men and women who married for the first time between 1968 to 1988. The figure shows clearly that both ages have been increasing over the time period. A linear regression of the age difference against time shows that the difference is increasing by 0.005 years per year with a t-statistic of 2.54. The increases in the ages of first marriages is inconsistent with case A. Case B predicts increases in the ages of first marriages with an increasing gender gap due to an innovation in birth control technology. Figure 1 is consistent with these predictions.

Figure 2 shows the evolution of the fractions of never married white men and women between 1950 to 1998. The fraction of never married white men increased in each decade from 1970. On the other hand, the fraction of never married white women fell between 1970 and 1990. It started to increase after 1990. The gender difference in the fractions never married grew between 1970 to 1990 and narrowed since. Case A predicts no change in the fraction of never married men or women which is inconsistent with the data. Case B predicts increases in the fraction of both never married men and women with no change in the gender gap. Thus it is consistent with the increase in the fraction of never married white men. However it is not consistent with the decline in the fraction of never married white women between 1970 and 1990.

Taken as a whole, the marital behavior of white men and women is most consistent with
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Case B. A caveat is in order. The increases in the ages of first marriages and fractions of never married individuals are also consistent with increasing educational attainment of individuals since the seventies. The percent of white men who completed college rose from 14.4% in 1970 to 25.3% in 1990.\textsuperscript{9} The percent of white women who completed college rose from 8.4% in 1970 to 19.0% in 1990. The decline in the fraction of never married white women between 1970 and 1990 is also problematic for the educational hypothesis. Moreover the rate of increase of college attainment by white women is faster than that of white men. But the gender gap in the age of first marriage slowly increased. In fact, the convergence in educational attainment for men and women potentially mitigated the divergence in the ages of first marriage.

Figure 3 shows the evolution of the average ages of black men and women who married for the first time between 1968 to 1988. The average age of first marriage for black men is always higher than that for black women. Both ages are trending up since 1970. The estimated coefficient of the age gap on time is -0.006 years per year with a t-statistics of -1.06. Thus one cannot reject the hypothesis of a small positive or no trend in the gender gap.

Figure 4 shows the evolution of the fractions of never married black men and women between 1950 to 1998. While both fractions have been increasing since the seventies, black men are more likely to be never married than black women. The gender gap has closed over time.

\textsuperscript{9} Educational attainment figures are from U.S. Census Bureau, \textit{Statistical Abstract of the United States: 2000}. 

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The gender differences in the ages of first marriage and fractions never married are clearly inconsistent with cases C and D. Thus there is no support for the hypothesis that innovations in birth control technology since the seventies affect black women adversely. With two exceptions, Figures 3 and 4 are most consistent with case B. The exceptions are (i) the gender gap in ages of first marriages should be rising rather than falling and (ii), the gender gap in the fractions never married should be constant rather than falling.

Instead of viewing Wilson’s hypothesis as an absolute scarcity of marriageable black men relative to marriageable black women, one can interpret the Wilson hypothesis as an increasing scarcity of marriageable black men since the sixties. This interpretation may be modelled as a fall in \( \gamma_m \), the probability that a mature man will remain in the marriage market in the next period, over time. It is beyond the scope of this paper to do a dynamic analysis of falling \( \gamma_m \). As a first pass, by focusing on case B, straightforward calculations show:

\[
\frac{\partial(a^B_m - a^B_f)}{\partial\gamma_m} > 0
\]

If \( \gamma_m \) was falling between 1968 and 1988, the gender gap in average ages of first marriage for blacks should be falling as is consistent with the evidence in Figure 3.

Also for case B,

\[
\frac{\partial(q^B_m - q^B_f)}{\partial\gamma_m} > 0
\]

As \( \gamma_m \) falls, the gender gap in the fractions of never married blacks should also fall which is also consistent with Figure 4. Thus a dynamic interpretation of the Wilson hypothesis and the effects of innovations in birth control technology can contribute towards understanding
the evolution of recent black marital behavior.

The above evidence applied to the aggregate marriage market in the US. It does not rule out the possibility that there may be some local marriage markets where marriageable men are extremely scarce, where the AYK effect is valid.

8 Unwed Parenthood in the Seventies

This section considers how changes in legal access to new reproductive control technologies affected the out-of-wedlock birth rates in different states in the US in the seventies. Following AYK, I will interpret an increase in the rate of out-of-wedlock childbearing due to easier access to new reproductive control technologies as a fall in the bargaining power of women in marriage. The converse is less clear. If the rate of out-of-wedlock childbearing fell due to increased availability of new technologies, there are two non-mutually exclusive explanations. First, the bargaining power of women in marriage may have increased as discussed in this paper. Second, the rate of out-of-wedlock childbearing may have fallen directly due to usage of the newly available technologies.

I will study how across states and years variations in legal access to contraceptives and legal abortions affected out-of-wedlock childbearing. The dependent variable, UNWED, is the ratio of the total number of out-of-wedlock births to the total number of births in a state-year.\textsuperscript{10} Altogether, there are 9 years of data from 1970-1977, & 1979 with 36 states per year that have consistent data.\textsuperscript{11} I also use LUNWED which is log(UNWED). The


\textsuperscript{11}The states are RI, IL, PA, WI, MA, MI, UT, CO, MO, AZ, AK, DC, WY, MN, WV, WA, VA, SC, OK, TN, SD, ND, NJ, OR, HW, NB, AL, LO, FL, NC, IN, NH, DE, KA, IO, KN.
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mean of UNWED across the 36 states in 1970 was 0.100. It increased steadily over the seventies and reached 0.159 in 1979.

The coincidental increase in UNWED over the seventies and the legalization of abortion in 1973 led AYK to argue that the legalization of abortion at least partially caused the increase in UNWED. However there were other factors that could have also led to an increase in UNWED over the seventies. In order abstract from national trends, I use across states variations in legal changes over time to estimate the effect of those changes on out-of-wedlock childbirths.

I consider two legal changes. First, I consider the change in access to legal abortions. Prior to Roe vs. Wade, the Supreme Court decision in 1973 which legalized abortions in all states, legal abortions were available in Alaska, California, Hawaii, New York and Washington. The dummy variable ABORTION for a particular state-year takes the value of zero when abortion was not legally available in that state-year. It takes a value of 1 when abortions were legal in that state-year.

Second, I consider changes in the earliest age in a state in which a woman may obtain contraceptive services (birth control pills) without parental consent. The variable LEGAL is the earliest age in a state-year in which a woman in that state-year may obtain contraceptive services without parent consent. This variable is obtained from Table 3 of Goldin and Katz. The variable is not updated for changes in values after 1974. The mean value in 1970 was 20.4 years with a standard deviation of 1.46 years. The mean value in 1974 was 16.2 years with a standard deviation of 2.20 years. Within four years, LEGAL fell by more than 20%.
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The across states variations in legal access to abortion over time have been shown by Gruber, Levine and Staiger to affect birth rates and average child outcomes in different state-years. The across states variations in changes in LEGAL have been shown by Goldin and Katz to affect the age of first marriage for college women and usage of the birth control pill. Thus state level variations in the two variables affected mean childbearing, marital and sexual behavior of women in the seventies.

Table 3 shows different regressions of UNWED or LUNWED on the two legal variables and other controls. In specifications [1] to [6], I control for state and year effects. In column [1], there is no statistically significant effect of either of the two variables on UNWED. The point estimates are also not economically significant. In columns [2] to [3], each legal variable is entered in a regression by itself. None of the estimates is economically or statistically significant. In column [4], LUNWED is the dependent variable. In this case, there is again no evidence to support the hypothesis that easier access to abortions increased the rate of out-of-wedlock childbearing. Increasing LEGAL, the legal age by which a woman may obtain contraceptives without parental consent, by one year increased the rate of out-of-wedlock childbearing by 0.4% with a t-statistic of 1.26. Although statistically insignificant, this estimate is not supportive of the hypothesis that easier access to contraceptives lowers the bargaining power of women in marriage. In columns [5] and [6], each legal change is entered by itself in a regression. The estimated coefficient on ABORTION remains statistically insignificant. The estimated coefficient and t-statistic on LEGAL remain the same as in column [4].
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<table>
<thead>
<tr>
<th>Dep. var.</th>
<th>1</th>
<th>2</th>
<th>3</th>
<th>4</th>
<th>5</th>
<th>6</th>
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</thead>
<tbody>
<tr>
<td>UNWED</td>
<td></td>
<td></td>
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<td></td>
<td></td>
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<tr>
<td>LUNWED</td>
<td></td>
<td></td>
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</tr>
<tr>
<td>ABORTION</td>
<td>0.00156 (0.36)</td>
<td>0.00170 (0.40)</td>
<td>0.0193 (0.55)</td>
<td>0.0126 (0.37)</td>
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</tr>
<tr>
<td>LEGAL</td>
<td>−0.00008 (−0.21)</td>
<td>−0.00010 (−0.27)</td>
<td>0.00357 (1.26)</td>
<td>0.00329 (1.18)</td>
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</tr>
<tr>
<td>DUMMIES</td>
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<td>state, yr</td>
<td>state, yr</td>
<td>state, yr</td>
<td>state, yr</td>
<td>state, yr</td>
</tr>
<tr>
<td>$R^2$</td>
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<td>0.969</td>
<td>0.969</td>
<td>0.973</td>
<td>0.973</td>
<td>0.973</td>
</tr>
<tr>
<td>$N$</td>
<td>324</td>
<td>324</td>
<td>324</td>
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<td>324</td>
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</tr>
<tr>
<td>$\sqrt{MSE}$</td>
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<td>0.00931</td>
<td>0.00931</td>
<td>0.0736</td>
<td>0.0735</td>
<td>0.734</td>
</tr>
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</table>

<table>
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<td>−0.020 (−1.55)</td>
<td>−0.239 (−1.89)</td>
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<td>LEGAL</td>
<td>0.00037 (1.47)</td>
<td>0.00052 (2.15)</td>
<td>0.0032 (1.37)</td>
<td>0.0041 (1.90)</td>
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<td>DUMMIES</td>
<td>state*T</td>
<td>state*T</td>
<td>state*T</td>
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Table 3: Out-of-wedlock Childbearing Regressions
In columns [7] to [12], I allow for state specific linear trends in the regressions rather than year effects. In general, the estimated effects of the variables of interest are stronger. In column [7], legalization of abortion reduced the out-of-wedlock birth rate by a statistically significant 0.003. Evaluated at the mean, it reduces UNWED by 2.4%. An increase in LEGAL by one year increased UNWED by 0.0004 with a t-stat of 1.47. In column [8], LEGAL is entered by itself. Now the point estimate is 0.0005 and the t-stat is 2.15. Evaluated at the mean, a one year increase in LEGAL increased UNWED by 0.4%. Note that this estimate is quantitatively similar to the estimates using LUNWED without state specific trends (i.e. in columns [4] & [5]). In column [9], ABORTION is entered by itself in the regression. Here the point estimate is -0.004 and it remains statistically significant. Taken together, the estimates in columns [7] to [9] support the hypothesis that easier access to reproductive control technologies decreased the out-of-wedlock childbearing rate. Columns [10] to [12] use LUNWED as the dependent variable. While the estimated coefficients are not as statistically significant than when the dependent variable is measured in levels, the estimated quantitative effects are similar. So again, the estimates in columns [10] to [12] support the hypothesis that easier access to reproductive control technologies decreased the out-of-wedlock childbearing rate.

The results in Table 3 may be summarized as follows. Without state specific trends in the regressions, there is largely no statistically significant effect of ABORTION or LEGAL on the rate of out-of-wedlock childbearing. There is some weak evidence that increasing LEGAL increased the out-of-wedlock childbearing rate when LUNWED is the dependent
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variable. With state specific trends in the regressions, the estimated effects show that easier access to reproductive control technologies reduced, both quantitatively and statistically, the out-of-wedlock childbearing rate. There is no evidence to support the hypothesis that easier access to reproductive control technologies increased the out-of-wedlock childbearing rate. Thus consistent with the rest of the evidence in this paper, there is no support for the AYK hypothesis that easier access to reproductive control technologies decreased the bargaining power of women in marriage.

9 Conclusion

Innovations in birth control technologies potentially have two very different effects on the welfare of women. This paper shows that the AYK assumption of extreme scarcity of marriageable men is necessary for welfare to fall. Empirical evidence on the recent evolution of black and white marital behavior suggests that the innovations in birth control technologies in the seventies improved the welfare of women. Using across states and years variation in access to reproductive control technologies, I show that there is no evidence to support the hypothesis that easier access to reproductive control technologies increased the out-of-wedlock childbearing rate. Although my substantive conclusion differs from AYK,12 their

12The circumstantial evidence also support my conclusion. Cook, et. al., Kane and Staiger, Lundberg and Plotnick, Goldin and Katz present evidence that women did use new birth control technologies when it became available. As for the transferable utility assumption, Chiappori, Fortin and Lacroix (2002) present direct evidence on marriage market tightness and transfers within marital households. Lundberg, Pollak and Wales (1997), Stevenson and Wolfers (2000) present evidence on changes in household behavior which may be linked to changes in the bargaining power of husbands and wives. As for the relative scarcity of marriageable women, Chamie and Nsuly show that the remarriage rate of divorced men exceeds that of divorced women for 47 countries for which there is data. Hamilton and Siow show that a marriage market model similar to the one discussed here, with an extra exit rate of women from the marriage market, can rationalize aggregate marital behavior in 18th century Quebec. The gender differences in the mean age of
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concern for the marriage market effects of innovations in reproductive technologies is as relevant as before.

A problem remains. Since recent innovations in birth control technologies are unlikely to have decreased the bargaining power of women in marriage nor increased the out-of-wedlock birth rate, the increase in the incidence of out of wedlock childbirths since the seventies remains a puzzle.\(^{13}\)

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13See Aiyagari, et. al., Neal and Willis for some models.
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References


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Figure 1: Whites' ages of first marriage

Figure 1:
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Figure 2: Fractions of whites never married

Figure 2:
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Figure 3: Blacks' ages of first marriage

Figure 3:
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Figure 4: Fractions of blacks never married

Figure 4: