Is the "Best Interests of the Child" Best for Every Child? The Long-Term Implications of Gender-Neutral Custody Laws*

Yang Chen†

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Abstract

I examine the impact of gender-neutral custody laws on child outcomes. Between the 1970s and 1990s, state custody laws moved from maternal preference to the "best interests of the child" doctrine which gives fathers and mothers equal treatment in child custody cases, a change that is independent of divorce law reforms. While standard household bargaining models predict that changes in custody laws give fathers greater bargaining power in marriages, the net effect of the custody law reform on all children is unknown. I exploit the variation across states in the timing of custody law changes to estimate the long-term implications of growing up in a gender-neutral custody law regime. I find that childhood exposure to gender-neutral custody laws has a negative and significant effect on educational attainment. For example, a man exposed to the new custody law as a child is less likely to graduate from high school by, on average, 2.04 percentage points. Results are similar for women. Moreover, the negative effects are independent of the effects of childhood exposure to unilateral divorce laws.

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† Email: chen.1474@osu.edu; Department of Economics, The Ohio State University, Arps Hall 410, 1945 N. High St., Columbus, OH, 43210-1172.

1 Introduction

Many studies have examined the impact of divorce on child outcomes. The general consensus is that children in divorced families tend to have lower academic performance. They are more likely to drop out of school, are less likely to attend or complete college, and are more likely to be unemployed and on public assistance (see, for example, Krein and Beller, 1988; Amato and Keith, 1991a, 1991b; Zill, Morrison and Coiro, 1993). In recent years, researchers have also been interested in how divorce laws and policies affect all children, including those in intact families. For example, Gruber (2004) examined the implications for children of growing up in states with unilateral divorce laws, which make divorce easier by not requiring the explicit consent of both partners. He found that adults exposed to unilateral divorce laws as children have lower level of education.

While scholars have looked at divorce law changes, few have looked into another important legal reform related to divorce: changes in child custody laws. In the United States, the assignment of child custody is governed by the custody law of each state, just like divorce. Between the 1970s and 1990s, states moved from explicit maternal preference to gender-neutral custody assignment. Prior to this change, the dominating rule in assigning child custody in divorce was the "tender years doctrine" which presumes that the mother is the more suitable custodian for children in the case of parental separation (Klaff, 1982; Buehler and Gerard, 1995; Jones, 1978). According to Jones (1978), courts awarded custody of minor children to mothers in more than 95% of all divorce cases under the tender years doctrine.

It was not until the 1970s that states started to question the validity of maternal preference. In the benchmark case, *Watts v. Watts*, in 1973, Judge Sybil Hart Kooper of the Family Court of New York ruled that any presumptive preference in favor of maternal custody violated the father's right to equal protection under the Fourteenth Amendment of the U.S. Constitution.[1] This invalidated the "tender years doctrine" in New York State. Other states followed the reform either by legislative action or judicial ruling. Since *Watts*, courts have moved to the "best

[1] 77 Misc. 2d 178, 350 N.Y.S.2d 285 (Fam. Ct. 1973).

interests of the child" doctrine, which consists of several criteria to determine which parent is more suitable to be the custodian. The doctrine makes no reference to the gender of the parent, and may include a decision of joint custody. The reform of custody laws was significant and dramatic. By 1990, 39 states had completed the transition to gender-neutral custody laws. At present, only three states still use some sort of maternal preference in custody laws: Idaho, Mississippi and Rhode Island. The movement towards gender-neutral child custody laws coincided with increasing divorce rates and the movement towards unilateral divorce laws (see Figure 1).

Despite the significant changes in custody laws, there has been little empirical analysis of the impact of these changes on children. Since most marriages involve children, the assignment of children in marital dissolution is an overlooked, and potentially important, factor in divorce trends. Custody assignment plays an important part before, during and after the divorce process. Changes in legal doctrine will not only result in increasing father custody after divorce, but more importantly, they also change the relative bargaining power between husbands and wives *ex ante*.

With maternal preference in custody laws, the majority of mothers went through divorce without the concern of losing child custody. The transition to gender-neutral custody assignment increases the possibility that child custody is assigned to fathers or jointly to both parents in contested custody cases.[2] Under the new regime, wives have more to lose in the case of divorce, which might decrease their incentive to divorce. The opposite applies to husbands.

I develop an intra-household Nash bargaining model (McElroy and Horney, 1981) to consider the implications of gender-neutral custody assignment for household bargaining. In the model, a spouse stays married if the utility gain from marriage exceeds the external threat point—each party's best option outside marriage. The comparative statics of the model yield straightforward empirical predictions: the change in the prospect of child custody assignment

[2] Bianchi (1995) used census data and calculated that the percentage of single-father households in all single-parent households has increased by 48.8% between 1970 and 1990. Garasky and Meyer (1996) broke down the single-parent households and found that the largest share of the increase in single-father households came from formerly-married fathers. My own calculations using census data support their calculations.

increases the external threat point for the husband, and decreases that for the wife. As a result, the custody law reform increases the husband's share of resources in the household.

Importantly, the unambiguous increase in fathers' bargaining position applies to all couples with children, whether divorce occurs or not. The aggregate effect on child outcomes, however, is theoretically ambiguous—it depends on the degree of bargaining power change within marriage, whether the couple will divorce, and the effect of marital dissolution on children whose parents divorce. As such, the analysis of divorce and children's development would not be complete without examining the laws governing custody assignment as they have implications for both intact and divorced families. It is possible that some of the effects that we have previously attributed to divorce may, in fact, be due to changes in child custody policies.

One reason for the lack of empirical evidence is the dearth of a comprehensive coding for when each state underwent a transition to gender-neutral custody laws. The major difficulty with constructing a legal coding is the inconsistency between state statutes and actual court practices. While several states had gender neutrality written into their statutes in the 1970s, courts still practiced explicit maternal preference in child custody cases. This is due to the fact that the maternal preference was often in the form of an implicit presumption used in custody cases rather than a formal legal statute. A coding that takes into account one aspect of the legal change without considering the other would be incomplete. Such mechanical coding would not reflect the timing of changes in actual practices.

To accurately measure changes in custody practices, I construct the custody law coding in an innovative way. I define a state to have completed the change in its custody law if the state has met the following criteria: (1) it has added statutes equalizing parental rights in custody assignment, and (2) maternal preference is no longer in use in court practices. This method utilizes information from both state statutes and court practices to determine when each state completed the change from maternal preference to gender-neutral custody assignment. That is, my coding is directly related to the *actual likelihood* that a custody dispute in a given year would be settled in a gender-neutral way.

My legal coding is the first to record each individual state's year of transition into gender-neutral custody assignment in a systematic and transparent way.[3] This coding enables the empirical analysis of the effect of custody law changes. Using the coding, I establish that states' movement to gender-neutral custody laws is independent of the adoption of unilateral divorce laws. I can therefore disentangle the effects of each legal reform in the empirical estimation.

In my empirical analysis, I estimate the long-term implications of growing up in a gender-neutral custody law regime. My empirical strategy exploits the variation across states in the timing of the legal changes. Using data from the U.S. Census and the American Community Survey from 1960 to 2010, I find that childhood exposure to gender-neutral custody laws has a negative and significant effect on educational outcomes. For example, a man exposed to the new custody law as a child is less likely to graduate from high school by, on average, 2.04 percentage points. With a high school graduation rate of 86.86% in the sample, this effect is substantively large.

The negative effect of gender-neutral custody laws increases with the length of exposure in childhood. I find that each additional year of exposure to the new custody law reduces the probability of high school graduation by 0.24 percentage points for men. The negative impact of being exposed before age five is the largest in magnitude. Results are similar for women.

In addition, I estimate a specification that takes into account changes in both custody laws and divorce laws. I incorporate Gruber's (2004) analysis on unilateral divorce laws, controlling for whether one lived under the unilateral divorce law regime as a child. I find that all my results are robust to the addition of the unilateral divorce law controls. That is, child custody law regime has an effect on children that is independent of divorce legislation.

This article proceeds as follows. In Section 2, I give an overview of custody law history and discuss my coding of the custody law transition. Section 3 presents the conceptual framework. In Section 4, I demonstrate the independence between the reforms of custody laws and divorce laws. Section 5 introduces the data. Section 6 and Section 7 present my empirical strategy and

[3] My legal coding of custody law changes is available upon request.

estimation results. Section 8 concludes.

2 Overview of Custody Law Changes

For over a century the dominating rule in assigning child custody in divorce was the tender years doctrine, which was an explicit preference for maternal custody.[4] In most states, the doctrine "establishes a presumption that children of their tender years should be placed in the custody of their mother, because she is best equipped to provide for the physical, emotional, and psychological needs of a young child" (Jones, 1978, p. 696). The doctrine was not always explicitly written in state statutes. It was also an implicit judicial presumption employed in case practices and cited by judges in decisions, whether there was a statute or not (Klaff, 1982).

The movement from maternal preference to gender-neutral custody assignment began in the 1970s. The feminist and the fathers' rights movements started to question the validity of the presumption that mothers are naturally superior to and more suitable for rearing children. In place of the tender years doctrine, courts began to determine children's custody assignment using a new guideline, the "best interests of the child" (BIOC) doctrine. BIOC is the dominant rule followed by most states today. It is usually written in state statutes, and consists of several criteria, such as the emotional ties between children and parents, capacity of parents to meet children's physical, emotional, and educational needs, stability and desirability of environment, and wishes of the child.

Coding Custody Law Transitions

One of the main contributions of this study is the documentation of the transition from maternal preference to gender neutrality in each state. The major difficulty in constructing a legal coding is the inconsistency between state statutes and actual court practice. While quite a few states had gender neutrality written into their statutes in the 1970s, courts continued to practice maternal preference in child custody cases. This is due to the fact that the tender years doctrine was very

[4] This section borrows heavily from Chen (2013).

often in the form of an implicit presumption rather than a legal statute.

The evolution of maternal preference in Utah is a typical example of the inconsistency between state statutes and case law. Maternal preference had long existed in Utah in the form of an explicit statute.[5] In 1977, the statutory presumption that the mother is the preferred custodian was repealed by the legislature.[6] However, the court continued to recognize a *judicial* preference in favor of the mother, all other things being equal. Several custody decisions in the late 1970s and early 1980s cited the tender years presumption or maternal preference and awarded child custody to mothers.[7] It was not until 1986 that the judicial preference towards mothers was explicitly abolished in the case, *Pusey v. Pusey*. After this case, maternal preference ceased to appear in judicial decision making in Utah.[8]

States like Utah present a serious issue of measurement in the coding of custody law changes. Simply relying on the custody statute would not give an accurate description of the case practice. It is entirely likely that a custody dispute decided in Utah would explicitly depend on maternal preference even with a state statute abolishing the practice. In other words, simply relying on the statute would not reflect the reality of custody assignment.[9]

To measure the actual changes in custody cases, I construct custody law transitions in a uniform way. I apply a consistent rubric to determine states' year of transition—I look for the year after which a custody dispute, if contested in court, would be decided on a gender-neutral basis. This way of coding utilizes information from both state statutes and court practices to accurately determine when each state completed the change from maternal preference to

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[5] The statute passed in 1903 required the custody of minor children to be awarded to mother (1903 Utah Laws, ch. 82, § 1). The statute later used the language "the natural presumption that the mother is best suited to care for young children" (1969 Utah Laws, ch. 72, § 7).

[6] See *Jorgensen v. Jorgensen*, 599 P.2d 510, 511 (Utah 1979).

[7] See *Henderson v. Henderson*, Utah, 576 P.2d 1289 (1978); *Smith v. Smith*, Utah, 564 P.2d 307 (1977); *Pennington v. Pennington*, 711 P.2d 254 (Utah 1985)

[8] See *Pusey v. Pusey*, 728 P.2d 117 (Utah 1986)

[9] Different methodologies usually result in different versions of legal codings. One example is the controversies over the legal coding of early legal access to the birth control pill (Bailey, Guldi and Hershbein, 2013; Joyce, 2013).

gender-neutral custody assignment. The rubric assures that my coding of legal transition reflects the actual likelihood of a child custody dispute being resolved on gender-neutral terms.

I carefully read through contested custody cases and state statutes during the transition period to identify the exact year when states completed the transition.[10] I develop consistent criteria and apply them systematically to all states. For example, most states had gender neutrality or best interests of the child written into legal statutes before they were mentioned in custody cases. For case rulings that simply upheld a gender-neutral state statute which was introduced earlier, I code the year when the statute was passed as the year of change. For case rulings that clarified, disavowed, or reinterpreted the previous statutes, on the other hand, the year when the case was decided would be used to code the year of transition.[11] I was able to determine the exact year of custody law changes for 48 states and Washington, D.C.[12]

Compared with the legal summaries prior to this study, my coding makes several contributions. First, I capture a more complete picture of the custody law changes as I look at both aspects of custody laws: statutes and case law. Most of the previous taxonomies only emphasize one aspect, which leads to inaccurate information about the timing of the legal changes. Second, my coding is transparent and conservative. I start from a definition not of policy but of *practice*. My coding is directly related to the actual likelihood that a custody dispute in a given year would be settled in a gender-neutral way. Third, much of the previous literature covers only a short period of time. Most existing studies summarize custody law development in the 1970s, when the removal of maternal preference first brought attention to the issue. My coding updates and extends the legal transition to the present, allowing one to see the full transition.

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[10] I use LexisNexis Academic Dataset for the source of case law. I also compiled information from several secondary sources that summarize the development in custody laws. Information on my sources are available in the legal appendix

[11] Legal appendix records detailed information for how each state's time of transition is determined.

[12] Timing of custody law changes was indeterminate for two states: Maine and Washington. I am not able find sufficient evidence of either maternal preference or gender neutrality in the past few decades.

Joint Custody

One issue worth noting is the introduction and development of joint custody. Joint custody, sometimes called shared custody, is the custodial arrangement where parents share the decision making (joint legal custody) or residential care (joint physical custody) of their children (Melli and Brown, 1995). Like gender-neutral custody assignment, joint custody is also a relatively new development in custody laws. Prior to 1975, joint custody was not an option when parents divorced.[13] Since then, joint custody laws started to be introduced in many states. Currently, joint custody is not available in only two states, Washington and West Virginia (Halla, 2013).

The adoption of joint custody and gender-neutrality are technically two separate aspects of the custody law. These two custody guidelines do not contradict each other in nature or in practice. Joint custody is gender-neutral by design, as the two parents share the decision making and physical care responsibility of their children. Indeed, some family court judges and parents perceive joint custody as an "easy out" solution when parents cannot reach an agreement on custody assignment (Miller, 1979).

Theoretically, the establishment of gender neutrality in custody laws is what alters the bargaining within marriage, not the establishment of joint custody. There are a small number of studies that have analyzed the effect of joint custody on divorce and child outcomes (Brinig and Buckley, 1998a; Leo, 2008; Halla, 2013).[14] What these studies fail to account for, however, is the fact that gender-neutral custody assignment is a necessary legal *precondition* for the establishment of joint custody. Only after maternal preference was destroyed in a state's custody law would joint custody be possibly considered as an option for child custody assignment. Also, there is no standard time lag between the abolition of maternal preference and the introduction of joint custody. The recognition of joint custody after the removal of maternal preference was not

[13] Jacob (1988) recorded that, prior to 1975, only North Carolina had a statutory authorizing joint custody which focused only on situations involving abuse and neglect.

[14] Halla (2013) and Leo (2008) studied the empirical effect of joint custody legislation on divorce rates and child outcomes. Brown and Flinn (2011) developed a continuous time model of marriage, fertility and parenting to study the effect of family policies on outcomes for husbands, wives and children. They included the share of paternal custody after divorce as one of the model parameters.

an automatic movement. On average, the time interval between the two is 5.04 years.[15] States varied in the length of time between the two. The standard deviation of the time interval is 8.24 years.

Joint custody, in essence, is a special case of gender-neutral custody assignment. With maternal preference, it was not possible to assign a portion of custody to each parent. In this study, I focus on the transition from maternal preference to gender-neutral custody laws. The reason is that "gender neutrality" is a more general concept than joint custody. It includes not only joint custody, but also other aspects of equal gender treatment, such as the possibility of the father gaining sole custody if he is qualified as a suitable custodian. Similar reasoning applies to other recent development in custody assignment such as the improvement of the rights of unwed fathers (Weitzman and Dixon, 1979). Gender neutrality is a necessary prerequisite for joint custody assignment and other related legal development, but not the other way around. The passage of gender-neutrality shapes the fathers' rights landscape in state custody laws and leads to changes in bargaining power within marriage before joint custody was introduced.

3 Conceptual Framework of Custody Laws and Child Outcomes

This section develops a simple Nash-bargaining framework to illustrate the effect of custody law changes on households. As with other divorce legislation, the adoption of gender-neutral custody laws has important implications for married parents, apart from the direct impact on divorced parents.[16] With the abrogation of maternal preference in custody cases, mothers lose their absolute advantage in receiving custody in the case of divorce, which decreases their expected valuation of potential divorce. The less desirable outside option alters the mothers' bargaining power position within marriage.

The Nash-bargained household decision framework is developed in the spirit of McElroy and

[15] This calculation is based my legal coding of gender-neutral custody laws and Halla's (2013) coding of joint custody.

[16] There have been a number of previous studies that studied the effect of family law on household bargaining (see, for example, Peters, 1986; Gray, 1998; Voena, 2013).

Horney (1981). In this framework, the married couple bargain in a two-person, non-zero-sum, cooperative game with a Nash solution. Another important aspect of the model is that parents value child custody after divorce. Even if a spouse does not desire child custody, per se, he or she can use the *possibility* of child custody as a threat. Evidence from custody cases suggest that each party may threaten or pretend to want custody in order to gain advantage in the bargaining process (Weitzman and Dixon, 1979; Mnookin and Kornhauser, 1979). What matters in the model is the effect of the custody regime on each spouse' bargaining position.

3.1 Model setup

Assume there is a household with only two people: male and female. Their objects of choice are $x = (x_m, x_f)'$, with given market prices $p = (p_m, p_f)'$, where x_m is the consumption by the husband, and x_f is the consumption by the wife. For simplicity, I do not define public good consumption. I also ignore the consumption of leisure. Neither of the above would alter predictions of the model. Also, in this model, children are not needed to reach the theoretical predictions.[17]

If the two individuals were not married, each of them would maximize a twice continuously differentiable, nondecreasing, quasiconcave utility function, subject to that individual's budget constraint, $p_k = I_k$, for k = m, f. I_k is the nonwage income. Each person has a well-defined continuous, strictly quasiconvex indirect utility function which gives the maximum attainable utility level as a function of prices and nonwage income:

$$V_0^k = V_0^k(p_k, I_k), \qquad \text{for k = m, f.}$$

Following McElroy and Horney's (1981) model, I assume that, when married, the utility of each spouse depends not only on own consumption, but also on that of the other spouse. Their individual utility functions are defined as:

[17] For simplicity, children are not directly modeled into the household utility function and budget constraint. The existence of children creates a channel through which the change in custody laws influences spouses' threat points.

$$U^k = U^k(x), \qquad \text{for } k = m, f.$$

Therefore, the gain from being married compared to single is:

$$U^k(x) - V_0^k(p_k, I_k), \qquad \text{for } k = m, f.$$

This term is assumed to be non-negative for the marriage to exist. When married, two spouses pool resources together, so the budget constraint for marriage is:

$$p_m x_m + p_f x_f = I_m + I_f.$$

I assume that bargaining over the allocation of x achieves the Nash solution that satisfies the following axioms: (1) invariance to linear transformations of individual utility functions, (2) Pareto efficiency, (3) independence of irrelevant alternatives, and (4) symmetry with respect to the roles of the players.

The couple chooses x to maximize the utility-gain product function subject to (4), what McElroy and Horney called the "utility-gain product function," which is a special case of the Nash product function,

$$N = [U^m(x) - V_0^m(p'_m, I_m;\ \alpha_m)]\left[U^f(x) - V_0^f(p'_f, I_f;\ \alpha_f)\right].$$

Each term in the bracket is the gain from marriage compared to being single. McElroy and Horney interpreted V_0^k as the threat point of each spouse, which is the best each could expect to get if the couple divorced. As outside situations change, the threat points may shift. The "relevant shift parameter" α_k represents the possible changes to the opportunities outside of the marriage. In the comparative statics of the model, I will discuss how the custody law reform affects the equilibrium by influencingα_k.

Setting the first partial derivatives of the Lagrangian function equal to zero gives the necessary conditions for maximization:

(1) $$N_k \equiv U_k^m\left(U^f - V_0^f\right) + (U^m - V_0^m)U_k^f = \lambda p_k, \qquad k = m, f,$$

(2) $$-N_\lambda \equiv p_m x_m + p_f x_f - I_m - I_f = 0,$$

where $N_k \equiv \frac{\partial N}{\partial x_k}$, $U_k^j = \frac{\partial U^j}{\partial x_k}$ for $j = m, f$ and $k = m, f$; $N_\lambda = \frac{\partial N}{\partial \lambda}$, and λ is the Lagrange multiplier. Solving (6) and (7) gives the optimal solutions:

$$x_k = h_k(\mathrm{p}', I_m, I_f) \qquad k = m, f,$$
$$\lambda = h_\lambda(\mathrm{p}', I_m, I_f)$$

3.2 Comparative Statics

To analyze the comparative statics of the effect of custody law changes, I use McElroy and Horney's definition of "iso-gain product curves" and "family rate of substitution". An iso-gain product curve (IGPC) is defined by the locus of bundles, x, for which the utility-gain product function, N, is constant. IGPCs have the standard properties of neoclassical indifference curves: (a) two curves never intersect; (b) an IGPC located further northeast has a higher value; (c) there is an IGPC going through every point in the space where the x is nonnegative; (d) IGPCs are convex.

The family rate of substitution (FRS) of x_f for x_m is defined as the negative slope of the IGPC at x:

$$FRS_{fm} = -\frac{\mathrm{d}x_m}{\mathrm{d}x_f}\bigg|_{\bar{N}} = \frac{N_f}{N_m}$$

Equilibrium is reached at the point at which the family rate of substitution equals the slope of the family budget constraint, as illustrated in Figure 2. At *x*, the FRS_{fm} depends on *p*, I_m, I_f, α_m and α_f. Custody law reform enters the model by altering the levels of α_m and α_f, which changes the equilibrium resource allocation by influencing FRS_{fm}.

In my analysis, the changes in the parameter α_k come from custody law changes. With the new custody assignment rule, females have a much higher risk of losing custody if divorce occurs, which lowers the expected valuation of divorce, i.e. the threat point. Custody law reform increasesα_m, the relative shift parameter for males, and decreasesα_f, the shift parameter for females. Later in the analysis, I will show how the change in α_k alters the resource allocation in equilibrium.

I evaluate the impact of custody law changes on FRS_{fm} by taking partial derivatives of

FRS_{fm} with respect to α_m and α_f[18]:

$$(3) \qquad \frac{\partial \mathrm{FRS_{fm}}}{\partial \alpha_\mathrm{m}} = \frac{\mathrm{U_m^m U_m^f}}{\mathrm{N_m^2}} \left(\frac{\mathrm{U_f^f}}{\mathrm{U_m^f}} - \frac{\mathrm{U_f^m}}{\mathrm{U_m^m}} \right) \left[\frac{\partial \mathrm{V_0^f}}{\partial \alpha_\mathrm{m}} (\mathrm{U^m} - \mathrm{V_0^m}) - \frac{\partial \mathrm{V_0^m}}{\partial \alpha_\mathrm{m}} \left(\mathrm{U^f} - \mathrm{V_0^f} \right) \right]$$

$$\equiv \frac{U_m^m U_m^f}{N_m^2} \left(\Delta MRS_{fm} \right) [W]$$

The first term in (3), $\frac{U_m^m U_m^f}{N_m^2}$, is positive. The second term, ΔMRS_{fm}, is the difference between female and male in the individual marginal rate of substitution of x_f for x_m. I assume individuals to be selfish, which implies that the female places a higher relative value on her own consumption than her husband does. Hence $\Delta MRS_{fm} > 0$.

The sign of the third term, W, is determined by the signs of $\frac{\partial V_0^f}{\partial \alpha_m}$ and $\frac{\partial V_0^m}{\partial \alpha_m}$, since $(\mathrm{U^m} - \mathrm{V_0^m})$ and $\left(\mathrm{U^f} - \mathrm{V_0^f} \right)$ are assumed to be nonnegative. With the change in custody laws, α_m increases. The value of being single decreases for the female, and increases for the male. Therefore, $\frac{\partial \mathrm{V_0^f}}{\partial \alpha_\mathrm{m}} < 0$, $\frac{\partial \mathrm{V_0^m}}{\partial \alpha_\mathrm{m}} > 0$, and W is negative. Therefore, $\frac{\partial \mathrm{FRS_{fm}}}{\partial \alpha_\mathrm{m}}$ is negative. The increase in α_m decreases FRS_{fm}. By symmetry, $\frac{\partial \mathrm{FRS_{fm}}}{\partial \alpha_\mathrm{f}} > 0$.

The adoption of gender-neutral custody laws raises α_m and decreases α_f, both of which will lead to a decrease in the family rate of substitution of x_f for x_m. With the family budget constraint unchanged, this will alter the allocation of resources in the household. As illurated by the example in Figure 3, when FRS_{fm}, or the slope of the IGPC decreases, the optimal consumption bundle moves from $\left(x_f^0, x_m^0 \right)$ to $\left(x_f^1, x_m^1 \right)$. The wife's consumption decreases and the husband's consumption increases due to the change in bargaining power. Note that, when defining consumption x, I do not differentiate between physical goods consumption and leisure consumption. This reallocation within the couple might have implications for both physical resources, time input into home production, and other types of comsumption.

[18] $\frac{\partial FRS_{fm}}{\partial \alpha_m} = \frac{\partial \left(\frac{N_f}{N_m} \right)}{\partial \propto_m} = \frac{N_m \frac{\partial N_f}{\partial \propto_m} - N_f \frac{\partial N_m}{\partial \propto_m}}{N_m^2}$. Substituting in (1) and $\frac{\partial N_k}{\partial \alpha_m} = -\frac{\partial V_0^m}{\partial \alpha_m} U_k^f - \frac{\partial V_0^f}{\partial \alpha_m} U_k^m$ gives (3).

3.3 Implications for Child Outcomes

The theoretical framework above predicts that movement to gender-neutral custody laws unambiguously improves the husbands' bargaining position within marriage. As a result, husbands enjoy a higher share of resources within marriage. Previous studies have shown that an increase in the women's share of resources in households leads to better child outcomes and higher consumption by children (see, for example, Thomas, 1990; Duflo, 2003; Lundberg, Pollak and Wales, 1997; Bobonis, 2009). Since the model predicts that fathers are unambiguously more empowered by the new custody law, we would expect that children living in *intact families* would have fewer resources devoted to them under the new gender-neutral custody regime. One problem with the empirical evidence on women's share of resources and child outcomes is that most of them are from developing countries. It is unclear whether this finding holds for U.S. families. An advantage of the analysis of child custody laws is that the legal reform leads to an unambiguous increase in the husband's relative bargaining position. As such, we can investigate whether that is related to child outcomes.

Apart from the impact on married families, changes in custody laws could also affect children whose parents divorce. One the one hand, custody law changes will affect the divorce rate. Chen (2013) finds that state divorce rates increase approximately seven years after states' adoption of the new custody law and persist thereafter. It is well known that divorce has negative impacts on child outcomes in many dimensions. For example, Amato and Keith (1991) found that children of divorce have more difficulties adjusting both socially and psychologically. Guidubaldi, Cleminshaw, Perry, and Mcloughlin (1983) found that adolescents whose parents have divorced are more likely to have low academic performance and to drop out of school even after one controls for socioeconomic status. Overall, divorce is often found to be correlated with adverse outcomes for children; although it is unclear whether children would be worse off had their parents remain married. On the other hand, the gender-neutral custody law is directly related to children's living arrangement after divorce. Its adoption can lead to a higher probability that children are assigned to a custodian (or joint custody of two parents) that is more

conducive to their best interests, compared to the case where maternal presumption dominates. Therefore, it is possible that some children are negatively affected by the additional divorces if their states have adopted gender-neutral custody laws, but other children could be better off.

Overall, the aggregate effect on child outcome is theoretically ambiguous, as it depends on the degree of bargaining power change within marriage and the effect of divorce on children whose parents divorce. Only investigating the outcomes for children with divorced parents would be misleading. Indeed, children from intact families are still the majority of the population. According to the statistics from the Survey of Income and Program Participation (SIPP), in 2009, 60% of the children live with both biological parents. In my empirical estimation, I will test the net effect of custody laws on all children.

I focus on educational attainment since it is an outcome directly related to parental investment in children and is an important factor determining well-being later in life. In particular, I concentrate on the extensive margin of high school graduation and the intensive margin of years of education. A substantial body of the literature on family structure and children's educational outcomes has also focused on these two outcomes. Studies have shown that high school completion depends crucially on the financial and time input from one's family (see, for example, Shaw, 1982; Krein and Beller, 1988; Astone and McLanahan, 1991).

4 Independence of Custody Law Reform and Unilateral Divorce Reform

Before analyzing the effect of custody law changes, it is important to establish that the movement away from the "tender years" doctrine to the gender-neutral custody laws and the movement to unilateral divorce laws are uncorrelated with each other.[19] If the two movements are related they could have been jointly determined, and disentangling the effects would be difficult. In this section, I will divide the states into two groups: states that have completed both legal reforms and states that have not changed one of the laws. For the former group, I test the

[19] This section reproduces the findings in Chen (2013).

correlation between the time of custody law reform and divorce law reform. For the latter group, I show that there is no particular pattern in when or whether states changed one law while they have not changed the other law.

Figure 4 plots the states that *have* completed both transitions. There are 31 such states. The vertical axis shows when each state adopted unilateral divorce[20] and the horizontal axis plots when each state changed its child custody law. If there is a strong positive or negative correlation between the timing of the two legal transitions, we would expect to see states clustered in a linear pattern in the graph. However, there is no such pattern. Indeed, when a line is fit to the relationship, the slope of the fitted line is 0.054 (std. err. = 0.08). This suggests that the correlation between the times of changes in two laws is very weak. Knowing *when* a state adopted the unilateral divorce law does not tell us *when* it changed its child custody law.

Similarly, knowing *whether* a state has adopted unilateral divorce does not tell us *when* or *whether* the state adopted the gender-neutral custody law either. Panel A of Table 1 lists the states that have not yet adopted unilateral divorce laws. If there is a strong positive correlation between the two legal changes, we should expect to see that most of these states would have custody laws unchanged. However, this is not the case. There is no pattern of states clustering in a certain year. In fact, states are quite spread out with respect to when they changed custody laws. Panel B of Table 1 lists states that have not yet changed custody laws. Similarly, there is no pattern in when these states adopted unilateral divorce laws.

In fact, if one takes a closer look at each individual state's family law history in the past few decades, it is easy to reach the conclusion that divorce regulations and custody laws indeed developed independently. A typical example is the state of New York. As discussed earlier, it was the first state to remove the tender years doctrine and establish the gender-neutral custody law in the United States. On the other hand, New York was the last state to adopt no-fault divorce. In August 2010, New York's Domestic Relations Law §170 permits divorce where "the relationship between husband and wife has broken down irretrievably for a period of at least six months,

[20] I use Gruber's (2004) coding for states' adoption of unilateral divorce laws.

provided that one party has so stated under oath." However, New York still lacks unilateral divorce regulations in its divorce law.

Another example is the state of Oklahoma. Contrary to New York, Oklahoma was one of the first states that adopted unilateral divorce. It passed a unilateral divorce bill in 1953. However, Oklahoma was behind most states in the transition to the gender-neutral custody laws. In 1980, when more than half of the states had outlawed the use of the tenders year doctrine, the Oklahoma court, in the case, *Boyle vs. Boyle*, stated that the Oklahoma's tender years-maternal preference statute is to be used as a "tie-breaker".[21] It was not until 1986, in the case, *Manhart v. Manhart*, that Oklahoma moved away from maternal preferences in determining custody.[22]

5 Data

The data for this analysis come from the U.S. Census and the American Community Survey (ACS). The ACS is conducted every year since 2000, and collects information previously contained only in the long form of the U.S. Census. For years before 2000, I use the decennial sample from the 1960 to 2000.[23]

To examine children's adult educational outcomes, I use individuals age 20 to 50. I collapse the data into cells by year, age, state of birth, state of residence and sex. Cells are used since variation in individuals' exposure to laws exists only at this level. Previous studies of unilateral divorce laws (Gruber, 2004; Wolfers, 2006) use a similar methodology. The data are also divided by sex as child custody law reform may have a different impact on boys and girls.[24] Regressions are run at the cell level, weighted by cell size.

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[21] See *Boyle vs. Boyle*, 615 P.2d 301 (Oklahoma 1980).

[22] See *Manhart v. Manhart*, 725 P.2d 1234 (Oklahoma 1986).

[23] I use the 1% sample from 1960 and 1970 census, the 5% sample from 1980 to 2000 census, the 0.4% sample from 2001 to 2004 ACS, the 1% sample from 2005-2010 ACS.

[24] Empirical evidence shows that fathers are more likely to contest a custody case if there is higher ratio of sons in the family (Weitzman and Dixon, 1979). Divorced fathers are much more likely to obtain custody of sons compared to daughters (Moretti and Dahl, 2008). Also, maternal preference was often stronger for daughters in the determination of custody cases. As a result, the removal of maternal preference may differ systematically in how it affects boys and girls.

The data report each person's age and place of birth, which I use to match to the time of legal changes in each state. This yields the key variables in the analysis: exposure to the gender-neutral custody law before age 18. I am able to compute both extensive and intensive measures of childhood exposure to gender-neutral custody laws. The details of the computation of these variables are discussed below.

Since the majority of states reformed their divorce laws and custody laws between the 1970s and 1990s, the data I use cover the period before and after both legal reforms. Individuals' exposure to both gender-neutral custody laws and unilateral divorce laws in the sample ranges from zero years to 18 years. Table 2 presents the sample means.[25] On average, around 45% of my sample has been exposed to gender-neutral custody laws before they were 18. Furthermore, I separate the sample by the length of exposure in the summary statistics. There is a relatively even distribution of individuals across different lengths of exposure. The sample provides sufficient variation in the main explanatory variable—the childhood exposure to gender-neutral custody laws—to study its impact on educational outcomes as adult.

6 Empirical Strategy

To assess the impact of being exposed to gender-neutral custody laws as a youth on adult outcomes, I employ an estimation strategy similar to that of Gruber (2004). The baseline specification is of the form:

$$Outcome_{sbta} = \beta_1 ChildCustody_{abt} + \sum_{s} State\ of\ residence_s + \sum_{b} State\ of\ birth_b + \sum_{r} Race_{sbta} + \sum_{t} \text{Age}_a * Time_t + \varepsilon_{sbta}$$

where, *a* indexes age, *s* indexes state, *b* indexes state of birth and *t* indexes year. *Outcome* is a measure of educational attainment. *Race* consists of two variables for the percentage of white

[25] The data I use in the empirical analysis excludes Maine and Washington, as their years of custody law changes are indeterminate in my legal coding.

and percentage of black in the cell. The regression also contains indicators for state of residence, state of birth, year, and interaction of age and year, in order to control for other factors that might be correlated with children's educational outcomes. In all specifications, standard errors are clustered by state of residence to correct for possible serial correlation within a state over time.

One potential concern with this specification is that there might be trends in child outcomes that are specific to each state. Moreover, these trends might be correlated with the legal reforms. I address this concern by estimating regressions which control for state-specific trends for both state of birth and state of residence in addition to state fixed effects. The state specific time trends control for time-varying factors within states that are correlated with legal reforms or changes in educational attainment by cohorts within a state. For example, if one state sees an increase in its school quality that does not occur in other states, state-specific time trends will be able to control for the effect of such increase on educational attainment. Similar reasoning applies to any state-specific trends that might influence states' adoption of gender-neutral custody laws. This allows me to identify the changes in educational outcomes that are only due to different exposure to custody laws.

$ChildCustody_{abt}$ is the measure of the exposure to the gender-neutral custody law before 18, which is the main variable of interest in this analysis. In my analysis, I use two different measures of the exposure in different specifications, a dichotomous and continuous measure. The two measures focus on different aspects of the influence of legal exposure. The dichotomous measure equals one if the individual has *ever* been exposed to gender-neutral custody law regime before he or she was 18, and equals zero otherwise. This variable measures the exposure to the new custody law at the extensive margin.

The second measure is a continuous variable, which measures the length of exposure to the gender-neutral custody law before 18. For example, if a child's birth state adopted the gender-neutral custody law when he or she was 6 years old, his or her length of exposure is 12 years. This measure looks at the exposure to the new custody law on the intensive margin.

In census and ACS surveys, there is information about individuals' birth place and state of

current residence, but they do not ask where people have lived during their youth. In my analysis, I assume that individuals stay in the same state from birth to 18 years old.[26] This might potentially cause bias in the estimation results. In Section 6.4, I address this problem and check the robustness of my results to using sample with the same state of birth and state of residence.

7 Estimation Results

7.1 Baseline Specification

In the baseline specification, I estimate the effect of childhood exposure to gender-neutral custody laws. Estimation results for men and women are presented in Table 3A and Table 3B, respectively. Each column reports results from a separate regression.

I estimate the effect of childhood legal exposure on two educational outcomes: the likelihood of graduating from high school and the average years of education. For the outcome of high school graduation, the coefficients are multiplied by 100 for easy interpretation as percentage changes. Within each outcome, the two measures of exposure to the law are used in separate regressions. In both tables, column (1), (2), (5), and (6) report the estimated coefficients on the dichotomous measure: being ever exposed to the new law before 18. Columns (3), (4), (7) and (8) report coefficients on the continuous measure: years of exposure before 18. As discussed in the empirical strategy section, I also control for state-specific trends for both state of birth and state of residence in a separate set of regressions. In each table, columns (2), (4), (6), (8) report results from specifications with trends.

In both Tables 3A and 3B all estimated coefficients are negative and statistically significant. Overall, exposure to the gender-neutral custody law before 18 has a negative impact on educational attainment, for both men and women. For example, being exposed to the new child custody law decreases the probability of graduating from high school by 2.04 percentage point for males. This result is *substantively* large: it is about 2.34% of the sample mean, and one sixth

[26] Similar assumptions are made in other studies with similar data constraints, see, for example, Gruber (2004) and Goldin and Katz (2002).

of the standard deviation. Being exposed to the gender-neutral custody law also decreases men's years of education by 0.123 years. The results are of similar magnitude for women. Exposure to the gender-neutral custody law as a child lowers women's likelihood of high school completion by 1.85 percentage points.

Not only does exposure to the new custody law matter for educational outcomes, the length of exposure determines the magnitude of the effect, as well. For men, one additional year of exposure to the gender-neutral custody law decreases the likelihood of high school completion by 0.239 percentage points, and reduces the years of education by 0.013 years. Again, I find similar results for women.

The negative coefficient on years of exposure confirms the negative impact of the gender-neutral custody law on the intensive margin. While the magnitude of one year's impact is rather small, the accumulation effects are large, given that the average length of exposure in the sample is 4.72 years for men and 4.68 years for women.

One should note, however, that the negative effect associated with years of exposure can have two potential implications. One is that the younger the child is first exposed to the gender-neutral custody law, the worse his or her educational outcome would be, holding other things equal. The other is that longer amount of time the child is exposed to the law negatively affects his or her outcome. Because the years of exposure is constructed using both year of birth and state's legal reform status in each year, it is impossible to disentangle these two effects. The negative coefficients could be due to younger age of first exposure, longer length of exposure, or a combination of the two.

Another finding from the Tables 3A and 3B, is that estimation results are quite robust to the inclusion of state-specific trends. Controlling for state-specific trends decreases the estimated coefficients in all regressions. However, the differences between coefficients with trends and without trends are quite modest. Since state-specific trends control for any time-varying factors within states that might be correlated with custody law reforms or influence educational attainment, the robustness to the inclusion of trends indicates that such factors do not

significantly bias the estimates in the original estimation.

My estimation results show that the aggregate effect of the gender-neutral custody law on children's educational attainment is negative. Exposure to the gender-neutral custody law as a child is associated with lower likelihood of graduating from high school and reduced years of education on average.[27] The negative effect increases with additional years of exposure. In my conceptual framework, I illustrate that gender-neutral custody laws increase the bargaining power of fathers in marriages, which has negative impacts on children in married households. However, theoretically, the net effect for all children is ambiguous. These results suggest that the net effect on all children is negative and large.

7.2 Results by Length of Exposure

In the baseline specification, one of the measures of exposure as a child that I examine is the years of exposure to the gender-neutral custody law. By regressing educational outcomes on years of exposure, I find that an additional year of exposure to the new custody law has a negative impact on educational outcomes for both men and women. One issue with the specification is that the amount of exposure may not necessarily affect children in a linear way. Longer length of exposure implies that the child is first exposed at a younger age, which can have a systematically different impact on the child compared to those who are exposed at a much later stage of development. Besides the effect of age first exposed, the amount of exposure itself can affect children in a nonlinear pattern. Chen (2013) illustrates the possibility of a dynamic linear response to changes in custody laws. State divorce rates do not begin to increase until the seven years after the adoption of gender-neutral custody laws in states.

To allow the effect of exposure to vary non-linearly, I estimate a more flexible specification. I define three dummy variables for one to five years of exposure, six to thirteen years of exposure,

[27] In Appendix Tables 1A and 1B, I present the estimation results for two additional educational outcomes: obtaining some college education and being a college graduate. The results are mostly negative for these two outcomes.

and fourteen to eighteen years of exposure.[28] The cutoffs are chosen so that individuals are divided into three groups by the age first exposed: first exposed after age thirteen, first exposed between age five and age thirteen, and first exposed before age five. These three age groups are often deemed by courts as development stages of children that are distinctly different (Klaff, 1982).

Instead of regressing educational outcomes on the single measure of years of exposure, I now estimate the three dummy variables in the specification. Table 4 reports the estimation results. Overall, the results are consistent with those from the baseline specifications. The coefficients on the exposure variables in most regressions are negative and statistically significant. Being exposed for 14 years and above have the largest negative impact on educational outcomes in all but one regression. For example, men exposed to gender-neutral custody laws for more than 14 years as children are 4.3 percentage points less likely to graduate from high school. The differences between six to thirteen years of exposure and one to five years of exposure are relatively small.

Overall, the reduction in educational outcomes is greater for longer exposure time. While one to five years and six to thirteen years of exposure have similar effects, being exposed for 14 years and above has the largest negative impact on outcomes. This also implies that first being exposed to gender-neutral custody laws before school age is strongly related to negative educational outcomes.[29]

7.3 Unilateral Divorce Laws

I carry out additional checks to verify the results from the baseline specification. I investigate whether the effect of exposure to gender-neutral custody laws holds when I control for the

[28] Gurber (2004) and Johnson and Mazingo (2000) used a similar specification to estimate the effect of childhood exposure to unilateral divorce laws on outcomes.

[29] Appendix Tables 2A and 2B report results from alternative specifications where I estimate the effect of ever been exposed before age 5 and exposed before age 13, individually. Coefficients on the two variables are also negative and significant.

exposure to unilateral divorce laws. Gruber (2004) and Johnson and Mazingo (2000) find that childhood exposure to unilateral divorce laws negatively affects adults' educational outcomes. Chen (2013) examines the correlation between the divorce law reform and child custody law reform, and shows that states' movement to gender-neutral custody laws is independent of the adoption of unilateral divorce laws. Therefore, I am able to estimate the childhood exposure to both unilateral divorce laws and gender-neutral custody laws in one specification.

To verify that I am not capturing any negative effect from unilateral divorce laws, I include an additional variable $ChildUnilateral_{abt}$ in the specification, indicating whether the child has ever been exposed to unilateral divorce laws before 18. The transitions in unilateral divorce laws are based on Gruber's (2004) coding, where he incorporates and updates Friedberg's (1998) coding using both primary and secondary sources. This is the most up-to-date coding of unilateral divorce law changes in the literature.[30]

Tables 5A and 5B report the results from specifications controlling for childhood exposure to unilateral divorce laws for men and women respectively. For simplicity, I show only regressions without state-specific trends. Appendix Tables 3A and 3B present results with state-specific trends. Results are similar.

For each educational outcome and exposure measure in Tables 5A and 5B, I present results from three regressions. The first columns reproduce the results from Table 3A and 3B, which estimates only the effect of exposure to child custody laws. The second columns present results from estimating only the effect of exposure to unilateral divorce laws, which is similar to the strategy used by Gruber (2004). The third regression simultaneously considers both unilateral divorce laws and child custody laws. Comparing the coefficients on exposure to custody laws in the first columns and the third columns, I find that they are very similar. When the control for unilateral divorce is added, the coefficient on custody laws is decreased by only a modest amount. The negative impact of gender-neutral custody laws is robust to the control of unilateral divorce laws.

[30] Other versions of the coding include, for example, Friedberg (1998), Brinig and Buckley (1998b), Nakonezny, Shull and Rodgers (1995), Ellman and Lohr (1998), and Johnson and Mazingo (2000).

I also obtain similar findings when comparing the second columns with the third columns. My estimation results for exposure to unilateral divorce laws are similar to those found by Gruber (2004). Moreover, they are robust to the control of child custody laws. Overall, the results suggest that the negative impact of the gender-neutral child custody laws is independent of the effect of unilateral divorce laws. Both legal changes have negative impacts on educational outcomes.

7.4 Robustness Check for Migration

As discussed in the empirical strategy section, I assume one's state of birth to be the state of residence between age zero and eighteen when constructing measures of childhood exposure to the gender-neutral custody law. Unless no one in the sample had ever migrated across states before turning 18, this measure will not be able to capture the true treatment status for everyone in the sample. This is a common problem with repeated cross-section data. For example, in the study on childhood exposure to unilateral divorce laws, Gruber (2004) assumes that the state of birth is the state of residence as youth. Similarly, Goldin and Katz (2002) use state of birth as a proxy for state of residence between 18 and 21 years old.

If children's migration was random, i.e., independent of states' custody law changes or their future educational outcomes, the estimated coefficients I obtained above would potentially be an underestimation of the true impact of gender-neutral custody laws. Some sample in the treatment group is mistakenly classified as a control group, and vice versa, which introduces standard measurement error. Furthermore, there is a concern that migration might be correlated with state custody laws or states' average educational attainment, which would further bias the estimation.

As a robustness check, I restrict the sample to individuals with the same state of birth and the current state of residence. For such individuals, it is still possible that they moved to other states during their childhood before moving back to the state of birth, but the possibility of migration is much smaller compared to those who are living in a different state from their state of birth. Table 6 presents the summary statistics of the restricted sample. Out of the original sample, 64.42% of

the individuals have the same state of birth and state of residence. The sample means of all variables are similar to those in Table 2.

Tables 7A and 7B report the results from regressions using the restricted sample. Instead of collapsing into cells by year, age, state of residence, state of birth and sex, I now collapse the sample by year, age, state of birth and sex. Compared with Tables 3A and 3B, the estimation results using the restricted sample are very close to those in regressions using the original sample. Most coefficients are slightly larger than those in the original estimation, which is consistent with my earlier prediction that migration might bias coefficients toward zero. The results from the robustness check suggest that potential migration does not cause serious issues in estimations using the whole sample.

8 Conclusion

Between the 1970s and 1990s, state custody laws moved from maternal preference, also known as the “tender years doctrine”, to the “best interests of the child” doctrine which gives fathers and mothers equal treatment in child custody cases. While the custody law reform intends to improve the welfare of the children who suffer from parental separation, it also induces changes in household bargaining for couples with children. This article attempts to evaluate the legal changes in custody assignment. I examine the implications of childhood exposure to gender-neutral custody laws for future educational outcomes as adults.

To empirically analyze the impact of the custody law reform, I created the first comprehensive coding of when each state changed its custody laws. I established that the pattern of custody law changes across states is independent of the movement towards unilateral divorce laws. The independence made it possible to estimate the two laws in one equation. I could compare my results with previous studies that only estimated the effect of unilateral divorce laws.

In my conceptual framework, I developed an intra-household Nash bargaining model which predicts that the new, gender-neutral custody regulations give fathers greater bargaining power.

Importantly, the change in the custody law alters the bargaining power for all married couples with children. The aggregate effect on child outcomes, however, is theoretically ambiguous, as it depends on the degree of bargaining power change within marriage and the effect of marital dissolution on children whose parents divorce.

Using the Census and the American Community Survey data, I estimated the effect of growing up in a gender-neutral custody law regime for children. My source of identification was the time difference in adopting the new custody laws across states. I found that being exposed to the gender-neutral custody law as a child reduces the likelihood of high school graduation and decreases average years of education. The negative effects have a larger magnitude when the years of exposure are longer, i.e., when the child is first exposed to the new custody law at a younger age. My results were robust to various checks, which includes controlling for childhood exposure to unilateral divorce laws, implying that changes in child custody laws have an effect on children that is independent of divorce legislation.

My empirical findings provide evidence that the increased bargaining power of fathers due to the adoption of gender-neutral custody laws negatively affects child outcomes. Living under a gender-neutral custody law regime has negative implications, on average, for children, which seems to contradict with the welfare-improving goal of the legal reform. The custody law reform not only affects children in divorced families, but has an impact on those in intact families as well. Overall, changes in child custody laws have negative effects that were previously unnoticed on the educational outcomes of all children.

The findings in this paper have notable implications for future research. Changes in child custody laws play an important and overlooked role in child outcomes. The externalities of the legal reform are large and negative. Future analysis of trends in divorce, marriage and household economics should account for the role of changes in child custody assignment.

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Table 1 - States that Have Reformed Only One Law

	Year of Custody Law Reform	No. of States
A. States without unilateral divorce law	1973	1
	1976	1
	1977	2
	1978	2
	1979	2
	1982	1
	1983	1
	1985	2
	1994	1
	1995	1
	1997	1
B. States without custody law changes	Year of Adopting Unilateral Divorce	No. of States
	1971	1
	1975	1

Note: I use Gruber's (2004) coding for states' adoption of unilateral divorce laws.

Table 2 - Sample Means

	Men	Women
Exposed to new custody law before age 18 (%)	45.93	45.75
1 to 5 years of exposure (%)	13.27	13.31
6 to 13 years of exposure (%)	16.10	16.10
14 + years of exposure (%)	16.56	16.34
Years of exposure under 18	4.72	4.68
Exposed to Unilateral divorce before age 18 (%)	36.09	35.96
Age	34.86	34.96
White (%)	86.50	86.35
Black (%)	8.84	8.84
High school graduate (%)	86.86	88.46
Years of education	13.00	13.08
No. of cells	607,653	613,753
No. of samples (before collapsing)	11,393,687	11,837,667

Notes: Data collapsed into cells by year, age, state of birth, state of residence and sex. IPUMS data from 1960 to 2010: 1960 1-percent state samples, 1970 1-percent state sample, 1980-2000 5-percent state samples, 2001-2004 0.4-percent ACS sample, 2005-2010 1-percent ACS sample. Restricted to population age 20-50. Excludes Maine and Washington. Sample means are weighted by number of observations in each cell.

Table 3A - The Effect of Custody Laws on Educational Outcomes, Men

Independent Variables	Dependent variable: High school graduate Mean = 86.86%				Dependent variable: Years of education Mean = 13.00 years			
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
	No trend	Trend	No trend	Trend	No trend	Trend	No trend	Trend
Ever exposed	-2.036**	-1.678**			-0.148**	-0.123**		
	(0.328)	(0.240)			(0.036)	(0.031)		
Years of exposure			-0.239**	-0.181**			-0.013**	-0.007*
			(0.051)	(0.032)			(0.004)	(0.003)
Controls								
Year effects	√	√	√	√	√	√	√	√
State of residence effects	√	√	√	√	√	√	√	√
State of birth effects	√	√	√	√	√	√	√	√
State of residence trends		√		√		√		√
State of birth trends		√		√		√		√
Age & cohort effects	√	√	√	√	√	√	√	√
% of black & white	√	√	√	√	√	√	√	√
Adjusted R^2	0.617	0.643	0.617	0.642	0.455	0.473	0.454	0.472
No. of cells	607,653	607,653	607,653	607,653	607,653	607,653	607,653	607,653

** p<0.01, * p<0.05, † p<0.10

Notes: With state population weights. Robust standard errors adjusted for heteroskedasticity are in parentheses. Coefficients on high school graduate are multiplied by 100. Restricted to population age 20-50. Excludes Maine and Washington.

Table 3B - The Effect of Custody Laws on Educational Outcomes, Women

Independent Variables	Dependent variable: High school graduate Mean = 88.46%				Dependent variable: Years of education Mean = 13.08 years			
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
	No trend	Trend	No trend	Trend	No trend	Trend	No trend	Trend
Ever exposed	-1.849**	-1.412**			-0.068*	-0.051*		
	(0.360)	(0.217)			(0.029)	(0.024)		
Years of exposure			-0.241**	-0.186**			-0.005	-0.001
			(0.053)	(0.025)			(0.004)	(0.003)
Controls								
Year effects	√	√	√	√	√	√	√	√
State of residence effects	√	√	√	√	√	√	√	√
State of birth effects	√	√	√	√	√	√	√	√
State of residence trends		√		√		√		√
State of birth trends		√		√		√		√
Age & cohort effects	√	√	√	√	√	√	√	√
% of black & white	√	√	√	√	√	√	√	√
Adjusted R^2	0.651	0.686	0.652	0.686	0.550	0.563	0.550	0.563
No. of cells	613,753	613,753	613,753	613,753	613,753	613,753	613,753	613,753

** p<0.01, * p<0.05, † p<0.10

Notes: With state population weights. Robust standard errors adjusted for heteroskedasticity are in parentheses. Coefficients on high school graduate are multiplied by 100. Restricted to population age 20-50. Excludes Maine and Washington.

Table 4 - Amount of Exposure to Custody Laws on Educational Outcomes

Dependent variable:	Men				Women			
	High school graduate Mean = 86.86%		Years of education Mean = 13.00 years		High school graduate Mean = 88.46%		Years of education Mean = 13.08 years	
Independent variables:	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
Length of exposure	No trend	Trend	No trend	Trend	No trend	Trend	No trend	Trend
1 to 5 years	-2.200**	-1.984**	-0.195**	-0.177**	-1.861**	-0.094**	-0.108**	-0.094**
	(0.288)	(0.222)	(0.033)	(0.028)	(0.320)	(0.021)	(0.026)	(0.021)
6 to 13 years	-2.509**	-1.998**	-0.111*	-0.067†	-2.583**	0.011	-0.019	0.011
	(0.520)	(0.389)	(0.048)	(0.038)	(0.533)	(0.031)	(0.039)	(0.031)
14 + years	-4.347**	-3.506**	-0.281**	-0.207**	-4.193**	-0.078†	-0.127*	-0.078†
	(0.830)	(0.503)	(0.070)	(0.047)	(0.883)	(0.041)	(0.057)	(0.041)
Controls								
Year effects	√	√	√	√	√	√	√	√
State of residence effects	√	√	√	√	√	√	√	√
State of birth effects	√	√	√	√	√	√	√	√
State of residence trends		√		√		√		√
State of birth trends		√		√		√		√
Age & cohort effects	√	√	√	√	√	√	√	√
% of black & white	√	√	√	√	√	√	√	√
Adjusted R^2	0.618	0.643	0.456	0.474	0.652	0.686	0.551	0.564
No. of cells	607,653	607,653	607,653	607,653	613,753	613,753	613,753	613,753

** p<0.01, * p<0.05, † p<0.10

Notes: With state population weights. Robust standard errors adjusted for heteroskedasticity are in parentheses. Coefficients on high school graduate are multiplied by 100. Restricted to population age 20-50. Excludes Maine and Washington.

Table 5A - The Effect of Custody Laws and Unilateral Divorce on Educational Outcomes, Men

Independent Variables	Dependent variable: High school graduate Mean = 86.86%						Dependent variable: Years of education Mean = 13.00 years					
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)	(11)	(12)
Ever exposed	-2.036**		-1.703**				-0.148**		-0.114**			
(Custody laws)	(0.328)		(0.307)				(0.036)		(0.035)			
Ever exposed		-2.201**	-1.854**					-0.212**	-0.189**			
(Unilateral divorce law)		(0.690)	(0.679)					(0.060)	(0.059)			
Years of exposure				-0.239**		-0.223**				-0.013**		-0.012**
(Custody laws)				(0.051)		(0.048)				(0.004)		(0.004)
Ever exposed					-2.201**	-2.037**					-0.212**	-0.203**
(Unilateral divorce law)					(0.690)	(0.645)					(0.060)	(0.057)
Controls												
Year effects	√	√	√	√	√	√	√	√	√	√	√	√
State of residence effects	√	√	√	√	√	√	√	√	√	√	√	√
State of birth effects	√	√	√	√	√	√	√	√	√	√	√	√
State of residence trends												
State of birth trends												
Age & cohort effects	√	√	√	√	√	√	√	√	√	√	√	√
% of black & white	√	√	√	√	√	√	√	√	√	√	√	√
Adjusted R^2	0.617	0.617	0.618	0.617	0.617	0.619	0.455	0.456	0.456	0.454	0.456	0.456
No. of cells	607,653	607,653	607,653	607,653	607,653	607,653	607,653	607,653	607,653	607,653	607,653	607,653

** p<0.01, * p<0.05, † p<0.10

Notes: With state population weights. Robust standard errors adjusted for heteroskedasticity are in parentheses. Coefficients on high school graduate are multiplied by 100. Restricted to population age 20-50. Excludes Maine and Washington. Specifications do not control for state of residence trends or state of birth trends. Results from specifications with trends are reported in Appendix Table 2A.

Table 5B - The Effect of Custody Laws and Unilateral Divorce on Educational Outcomes, Women

Independent Variables	Dependent variable: High school graduate Mean = 88.46%						Dependent variable: Years of education Mean = 13.08 years					
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)	(11)	(12)
Ever exposed	-1.849**		-1.604**				-0.069*		-0.046†			
(Custody laws)	(0.360)		(0.323)				(0.029)		(0.027)			
Ever exposed		-1.688*	-1.359†					-0.131**	-0.121**			
(Unilateral divorce law)		(0.824)	(0.810)					(0.046)	(0.045)			
Years of exposure				-0.241**		-0.230**				-0.005		-0.004
(Custody laws)				(0.053)		(0.050)				(0.004)		(0.003)
Ever exposed					-1.688*	-1.515†					-0.131**	-0.128**
(Unilateral divorce law)					(0.824)	(0.783)					(0.046)	(0.045)
Controls												
Year effects	√	√	√	√	√	√	√	√	√	√	√	√
State of residence effects	√	√	√	√	√	√	√	√	√	√	√	√
State of birth effects	√	√	√	√	√	√	√	√	√	√	√	√
State of residence trends												
State of birth trends												
Age & cohort effects	√	√	√	√	√	√	√	√	√	√	√	√
% of black & white	√	√	√	√	√	√	√	√	√	√	√	√
Adjusted R^2	0.651	0.651	0.652	0.652	0.651	0.653	0.550	0.551	0.551	0.550	0.551	0.551
No. of cells	613,753	613,753	613,753	613,753	613,753	613,753	613,753	613,753	613,753	613,753	613,753	613,753

** p<0.01, * p<0.05, † p<0.10

Notes: With state population weights. Robust standard errors adjusted for heteroskedasticity are in parentheses. Coefficients on high school graduate are multiplied by 100. Restricted to population age 20-50. Excludes Maine and Washington. Specifications do not control for state of residence trends or state of birth trends. Results from specifications with trends are reported in Appendix Table 2B.

Table 6 - Sample Means for Individuals with the Same State of Birth and State of Residence
A direct comparison with Table 2

	Men	Women
Exposed to new custody law before age 18 (%)	47.22887	47.01159
Exposed to Unilateral divorce before age 18 (%)	37.18503	36.87589
Age	34.4213	34.51598
White (%)	10.90243	12.79763
Black (%)	84.35505	82.45939
Years of education	12.69659	12.85333
High school graduate (%)	85.29866	87.39243
No. of cells	22781	22772
No. of samples (before collapsing)	7341530	7625136

Note: Sample only includes individuals with the same state of birth and current state of residence. Data collapsed into cells by year, age, state of birth and sex. IPUMS data from 1960 to 2010: 1960 1-percent state samples, 1970 1-percent state sample, 1980-2000 5-percent state samples, 2001-2010 ACS state samples. Restricted to population age 20-50. Excludes Maine and Washington.

Table 7A - Robustness Check for Migration. The Effect of Custody Laws on Educational Outcomes, Men Restricted to individuals with the same state of bith and state of residence. A direct comparison with Table 3A.

Independent Variables	Dependent variable: High school graduate Mean = 85.30%				Dependent variable: Years of education Mean = 12.70 years			
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
	No trend	Trend	No trend	Trend	No trend	Trend	No trend	Trend
Ever exposed	-2.393**	-2.025**			-0.182**	-0.155*		
	(0.616)	(0.466)			(0.065)	(0.059)		
Years of exposure			-0.257**	-0.225**			-0.012*	-0.008†
			(0.079)	(0.054)			(0.006)	(0.004)
Controls								
Year effects	√	√	√	√	√	√	√	√
State of birth effects	√	√	√	√	√	√	√	√
State of birth trends		√		√		√		√
Age & cohort effects	√	√	√	√	√	√	√	√
% of black & white	√	√	√	√	√	√	√	√
Adjusted R^2	0.887	0.915	0.887	0.915	0.779	0.811	0.777	0.809
No. of cells	22,781	22,781	22,781	22,781	22,781	22,781	22,781	22,781

** p<0.01, * p<0.05, † p<0.10

Notes: With state population weights. Robust standard errors adjusted for heteroskedasticity are in parentheses. Coefficients on high school graduate are multiplied by 100. Restricted to population age 20-50. Excludes Maine and Washington.

Table 7B - Robustness Check for Migration. The Effect of Custody Laws on Educational Outcomes, Women
Restricted to individuals with the same state of bith and state of residence. A direct comparison with Table 3B.

Independent Variables	Dependent variable: High school graduate Mean = 87.39%				Dependent variable: Years of education Mean = 12.85 years			
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
	No trend	Trend	No trend	Trend	No trend	Trend	No trend	Trend
Ever exposed	-2.010**	-1.567**			-0.080	-0.062		
	(0.618)	(0.411)			(0.049)	(0.044)		
Years of exposure			-0.260**	-0.223**			-0.005	-0.002
			(0.085)	(0.047)			(0.005)	(0.004)
Controls								
Year effects	√	√	√	√	√	√	√	√
State of birth effects	√	√	√	√	√	√	√	√
State of birth trends		√		√		√		√
Age & cohort effects	√	√	√	√	√	√	√	√
% of black & white	√	√	√	√	√	√	√	√
Adjusted R^2	0.889	0.930	0.890	0.930	0.837	0.856	0836	0.856
No. of cells	22,772	22,772	22,772	22,772	22,772	22,772	22,772	22,772

** p<0.01, * p<0.05, † p<0.10

Notes: With state population weights. Robust standard errors adjusted for heteroskedasticity are in parentheses. Coefficients on high school graduate are multiplied by 100. Restricted to population age 20-50. Excludes Maine and Washington.

Appendix Table 1A - The Effect of Custody Laws on Other Educational Outcomes, Men

Independent Variables	Dependent variable: Some college Mean = 52.84%				Dependent variable: Collge and above Mean = 23.19%			
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
	No trend	Trend	No trend	Trend	No trend	Trend	No trend	Trend
Ever exposed	-1.903**	-2.042**			-0.445	-0.587		
	(0.544)	(0.559)			(0.417)	(0.424)		
Years of exposure			-0.004	0.027			0.026	0.018
			(0.065)	(0.069)			(0.048)	(0.064)
Controls								
Year effects	√	√	√	√	√	√	√	√
State of residence effects	√	√	√	√	√	√	√	√
State of birth effects	√	√	√	√	√	√	√	√
State of residence trends		√		√		√		√
State of birth trends		√		√		√		√
Age & cohort effects	√	√	√	√	√	√	√	√
% of black & white	√	√	√	√	√	√	√	√
Adjusted R^2	0.411	0.417	0.410	0.416	0.276	0.284	0.276	0.284
No. of cells	607,653	607,653	607,653	607,653	607,653	607,653	607,653	607,653

** p<0.01, * p<0.05, † p<0.10

Notes: With state population weights. Robust standard errors adjusted for heteroskedasticity are in parentheses.All coefficients are multiplied by 100. Restricted to population age 20-50. Excludes Maine and Washington.

Appendix Table 1B - The Effect of Custody Laws on Other Educational Outcomes, Women

Independent Variables	Dependent variable: Some college Mean = 55.31%				Dependent variable: Collge and above Mean = 23.04%			
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
	No trend	Trend	No trend	Trend	No trend	Trend	No trend	Trend
Ever exposed	-0.645†	-0.785*			0.376	0.146		
	(0.360)	(0.339)			(0.425)	(0.397)		
Years of exposure			-0.020	-0.016			0.133*	0.121†
			(0.067)	(0.064)			(0.063)	(0.071)
Controls								
Year effects	√	√	√	√	√	√	√	√
State of residence effects	√	√	√	√	√	√	√	√
State of birth effects	√	√	√	√	√	√	√	√
State of residence trends		√		√		√		√
State of birth trends		√		√		√		√
Age & cohort effects	√	√	√	√	√	√	√	√
% of black & white	√	√	√	√	√	√	√	√
Adjusted R^2	0.597	0.601	0.597	0.601	0.321	0.334	0.322	0.335
No. of cells	613,753	613,753	613,753	613,753	613,753	613,753	613,753	613,753

** p<0.01, * p<0.05, † p<0.10

Notes: With state population weights. Robust standard errors adjusted for heteroskedasticity are in parentheses.All coefficients are multiplied by 100. Restricted to population age 20-50. Excludes Maine and Washington.

Appendix Table 2A - The Effect of Exposure Before Age 13 and Age 5, Men

Independent Variables	Dependent variable: High school graduate Mean = 86.86%				Dependent variable: Years of education Mean = 13.00 years			
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
	No trend	Trend	No trend	Trend	No trend	Trend	No trend	Trend
Ever exposed before 13	-1.134**	-0.529†			0.010	0.064*		
	(0.410)	(0.275)			(0.039)	(0.029)		
Ever exposed before 5			-1.400**	-0.873**			-0.121**	-0.084**
			(0.361)	(0.163)			(0.034)	(0.028)
Controls								
Year effects	√	√	√	√	√	√	√	√
State of residence effects	√	√	√	√	√	√	√	√
State of birth effects	√	√	√	√	√	√	√	√
State of residence trends		√		√		√		√
State of birth trends		√		√		√		√
Age & cohort effects	√	√	√	√	√	√	√	√
% of black & white	√	√	√	√	√	√	√	√
Adjusted R^2	0.615	0.641	0.615	0.642	0.453	0.472	0.454	0.472
No. of cells	607,653	607,653	607,653	607,653	607,653	607,653	607,653	607,653

** p<0.01, * p<0.05, † p<0.10

Notes: With state population weights. Robust standard errors adjusted for heteroskedasticity are in parentheses. Coefficients on high school graduate are multiplied by 100. Restricted to population age 20-50. Excludes Maine and Washington.

Appendix Table 2B - The Effect of Exposure Before Age 13 and Age 5, Women

Independent Variables	Dependent variable: High school graduate Mean = 88.46%				Dependent variable: Years of education Mean = 13.08 years			
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
	No trend	Trend	No trend	Trend	No trend	Trend	No trend	Trend
Ever exposed before 13	-1.424**	-0.849**			0.047	0.081**		
	(0.388)	(0.193)			(0.031)	(0.024)		
Ever exposed before 5			-1.283**	-0.798**			-0.077*	-0.060†
			(0.368)	(0.124)			(0.031)	(0.030)
Controls								
Year effects	√	√	√	√	√	√	√	√
State of residence effects	√	√	√	√	√	√	√	√
State of birth effects	√	√	√	√	√	√	√	√
State of residence trends		√		√		√		√
State of birth trends		√		√		√		√
Age & cohort effects	√	√	√	√	√	√	√	√
% of black & white	√	√	√	√	√	√	√	√
Adjusted R^2	0.650	0.685	0.650	0.685	0.550	0.563	0.550	0.563
No. of cells	613,753	613,753	613,753	613,753	613,753	613,753	613,753	613,753

** p<0.01, * p<0.05, † p<0.10

Notes: With state population weights. Robust standard errors adjusted for heteroskedasticity are in parentheses. Coefficients on high school graduate are multiplied by 100. Restricted to population age 20-50. Excludes Maine and Washington.

Appendix Table 3A - The Effect of Custody Laws and Unilateral Divorce on Educational Outcomes, Men
All specifications include trends for state of residence and state of birth. A direct comparison to Table 5A, which has no trend.

Independent Variables	Dependent variable: High school graduate Mean = 86.86%						Dependent variable: Years of education Mean = 13.00 years					
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)	(11)	(12)
Ever exposed	-1.678**		-1.353**				-0.123**		-0.087**			
(Custody laws)	(0.240)		(0.225)				(0.031)		(0.029)			
Ever exposed		-2.284**	-1.937**					-0.236**	-0.214**			
(Unilateral divorce law)		(0.378)	(0.375)					(0.043)	(0.043)			
Years of exposure				-0.181**		-0.176**				-0.007*		-0.007*
(Custody laws)				(0.032)		(0.031)				(0.003)		(0.003)
Ever exposed					-2.284**	-2.248**					-0.236**	-0.235**
(Unilateral divorce law)					(0.378)	(0.356)					(0.043)	(0.042)
Controls												
Year effects	√	√	√	√	√	√	√	√	√	√	√	√
State of residence effects	√	√	√	√	√	√	√	√	√	√	√	√
State of birth effects	√	√	√	√	√	√	√	√	√	√	√	√
State of residence trends	√	√	√	√	√	√	√	√	√	√	√	√
State of birth trends	√	√	√	√	√	√	√	√	√	√	√	√
Age & cohort effects	√	√	√	√	√	√	√	√	√	√	√	√
% of black & white	√	√	√	√	√	√	√	√	√	√	√	√
Adjusted R^2	0.643	0.643	0.644	0.642	0.643	0.644	0.473	0.474	0.474	0.472	0.474	0.474
No. of cells	607,653	607,653	607,653	607,653	607,653	607,653	607,653	607,653	607,653	607,653	607,653	607,653

** p<0.01, * p<0.05, † p<0.10

Notes: With state population weights. Robust standard errors adjusted for heteroskedasticity are in parentheses. Coefficients on high school graduate are multiplied by 100. Restricted to population age 20-50. Excludes Maine and Washington.

Appendix Table 3B - The Effect of Custody Laws and Unilateral Divorce on Educational Outcomes, Women
All specifications include trends for state of residence and state of birth. A direct comparison to Table 5B, which has no trend.

Independent Variables	Dependent variable: High school graduate Mean = 88.46%						Dependent variable: Years of education Mean = 13.08 years					
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)	(11)	(12)
Ever exposed	-1.412**		-1.207**				-0.0512*		-0.034			
(Custody laws)	(0.217)		(0.195)				(0.024)		(0.024)			
Ever exposed		-1.538**	-1.227**					-0.112**	-0.103**			
(Unilateral divorce law)		(0.435)	(0.425)					(0.035)	(0.035)			
Years of exposure				-0.186**		-0.183**				-0.001		-0.001
(Custody laws)				(0.025)		(0.024)				(0.003)		(0.003)
Ever exposed					-1.538**	-1.498**					-0.112**	-0.112**
(Unilateral divorce law)					(0.435)	(0.412)					(0.035)	(0.034)
Controls												
Year effects	√	√	√	√	√	√	√	√	√	√	√	√
State of residence effects	√	√	√	√	√	√	√	√	√	√	√	√
State of birth effects	√	√	√	√	√	√	√	√	√	√	√	√
State of residence trends	√	√	√	√	√	√	√	√	√	√	√	√
State of birth trends	√	√	√	√	√	√	√	√	√	√	√	√
Age & cohort effects	√	√	√	√	√	√	√	√	√	√	√	√
% of black & white	√	√	√	√	√	√	√	√	√	√	√	√
Adjusted R^2	0.686	0.685	0.686	0.686	0.685	0.686	0.563	0.563	0.563	0.563	0.563	0.563
No. of cells	613,753	613,753	613,753	613,753	613,753	613,753	613,753	613,753	613,753	613,753	613,753	613,753

** p<0.01, * p<0.05, † p<0.10

Notes: With state population weights. Robust standard errors adjusted for heteroskedasticity are in parentheses. Coefficients on high school graduate are multiplied by 100. Restricted to population age 20-50. Excludes Maine and Washington.

Figure 1 – Custody law reform, divorce law reform and national divorce rate

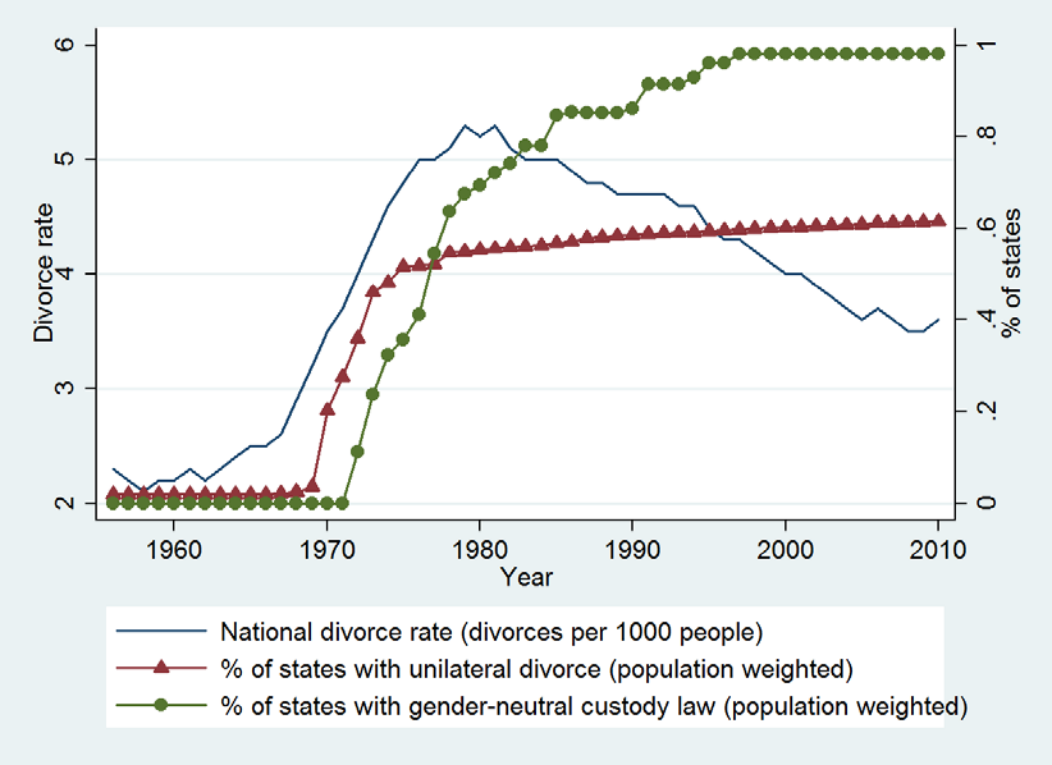

Figure 2 – Equilibrium: tangent point between IGPC and family budget constraint

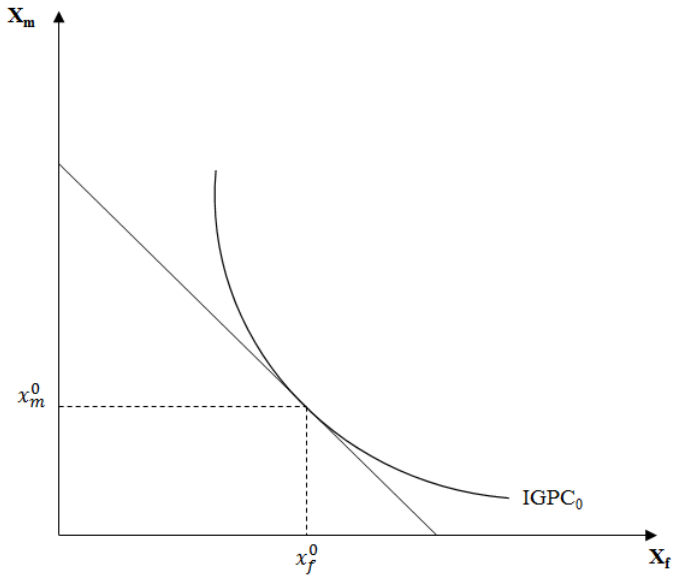

Figure 3 – Effect of FRS change on equilibrium

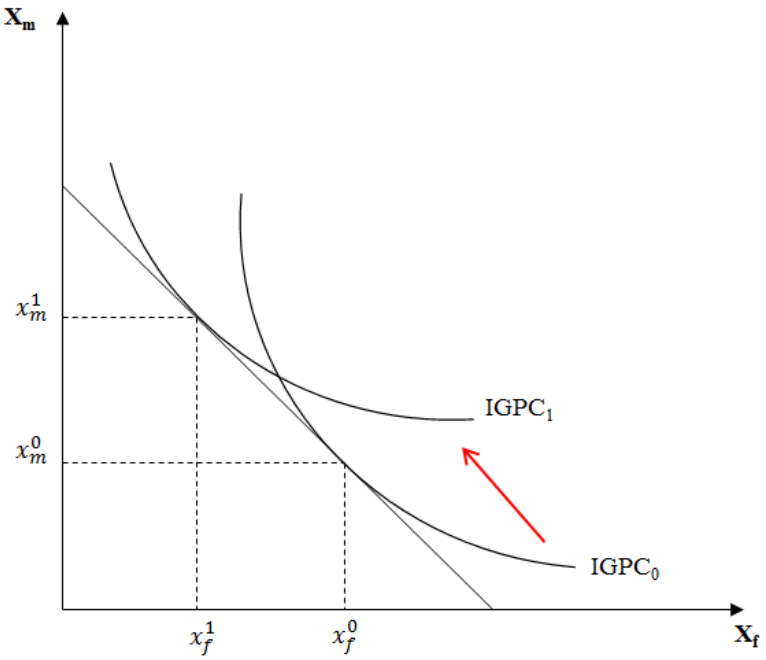

Figure 4 - Years when states changed their laws