The Influence of Transaction Costs and Legal Regimes on Divorce

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Abstract

This analysis extends H. Elizabeth Peters' "Marriage and Divorce: Information Constraints and Private Contracting" (1986). Peters introduced two models of marital bargaining. We argue those models may be viewed as special cases of a single, three-dimensional framework. The third dimension represents the transaction costs of intramarital exchange. The transaction-cost model illustrates that changes in divorce law have a marginal effect on economic outcomes, and the magnitude of the effect increases along with transaction costs. Transaction costs vary across a population. Therefore, on the margin, divorce law will influence which type of couples divorce.

Using divorce data from 1968 to 1985, we find significant differences in the characteristics of divorcing couples across legal regimes. These differences are consistent with our beliefs about the transaction costs impeding intramarital exchange. We find these differences are generally consistent across definitions of unilateral divorce, not representative of household sorting prior to divorce, and persist beyond the transition period between legal regimes. All of these tests support the hypothesis that transaction costs influence divorce outcomes, and that the influence varies across legal regimes.

I. Introduction

Does the Coase Theorem apply to martial relations? H. Elizabeth Peters explored this question in her pioneering paper "Marriage and Divorce: Informational Constraints and Private Contracting" (1986). She concluded that legal regimes did not affect aggregate divorce rates, and intramarital negotiation was best described by Coasean bargaining. In other words, the divorce rate was socially efficient regardless of the assignment of property rights.

Peters' result provoked a series of empirical analyses. Two noteworthy examples are works by Leora Friedberg (1998) and Justin Wolfers (2006). Friedberg concluded that legal regimes had an economically significant effect on divorce rates. Wolfers employed a panel of data that extended Friedberg's sample. He found that the divorce rate increased during the transition between legal regimes, but questioned whether that effect persisted over time. A key conclusion from both of these studies, among others, is that changing from a mutual-consent to a unilateral legal regime positively affected divorce rates, at least initially.

The Coase Theorem comes from Ronald Coase's revolutionary analysis of social costs (1960), and it applies to marital relations if and only if the transaction costs of intramarital exchange are zero. This study develops a theoretical model to interpret existing empirical evidence. It further substantiates the view that transaction costs are positive within the household, and therefore argues the Coase Theorem should not be

applied to marital relations. Moreover, it illustrates that the magnitude of the law's marginal effect is a positive function of transaction costs. The transaction costs impeding intramarital exchange are expected to vary from couple to couple, implying that the law governing divorce will influence which type of couples divorce, *ceteris paribus*.

The notion that the law's marginal effect varies across individuals has been suggested in the literature, but principally as an aside. Peters implies this result, and Douglas W. Allen (1998), among others, reached this general conclusion. The most dramatic example may come from Betsy Stevenson and Wolfers (2006), who found that shifts to unilateral divorce were followed by significant declines in the rates of domestic violence, female suicide and females murdered by their partners. The principal contribution of this model is to explicitly identify what others have implicitly suggested: the marginal effect of the law varies across individuals, and this variation is due to transaction costs.

In her study, Peters considered two distinct, two-dimensional models of intramarital bargaining. This study demonstrates that those models may be viewed as special cases of a single, three-dimensional framework. The third dimension represents the transaction costs of intramarital exchange.

What precisely comprises transaction costs? As used herein, the term refers to the costs of engaging in an economic exchange, or delineating ownership of an economic asset. In marriage, individuals collect information about outside opportunities, and engage in negotiation to divide the gains from marriage. This behavior—collecting information and

negotiation—is an attempt by individuals to increase their economic rights, and more efficiently allocate tasks between them. It is done not just at the beginning of the relationship, but throughout the entire relationship. This behavior is costly, and it plainly fits the definition of transaction costs above.

Browning and Chiappori (1998), among many others, note that married individuals are engaged in a repeated game, and suggest spouses have intimate knowledge of each other's actions and preferences. This argument is frequently employed to rationalize the use of symmetric information when modeling intrahousehold exchange. However, in reality, acquiring genuinely symmetric information is prohibitively costly.[1] Spouses may have intimate knowledge of each others preferences and actions, but intimate knowledge is not necessarily complete. Even if symmetric information was attainable, assuming its existence within marriage may be inconsistent with the assumption that spouses specialize in household and market production in order to gain from trade. Consider that spouses who specialize will devote their time to different activities, typically in different locations. Many will spend only a few waking hours per day in each other's company, which is hardly conducive to acquiring perfect information.

Transaction costs are often overlooked as an economic concept because they are difficult to empirically identify. An empirical literature exploring their effects exist, but it is relatively small. Macher and Richman (2008) provide a comprehensive overview,

[1] Consider that most transactions have a large—possibly infinite—number of potential outcomes. Negotiating each of these outcomes *ex ante* is prohibitively costly to both parties. See Barzel (1997) for an extended discussion.

spanning approximately 900 papers from multiple disciplines. They identify only one paper—Hamilton (1999)—that explicitly considers the influence of transaction costs on intrahousehold behavior.

While most empirical studies fail to acknowledge transaction costs, many implicitly address their effects. Within the divorce literature, consider that Peters and Friedberg both controlled for the presence of children and religious affiliation. Those variables influence the gains from marriage and an individual's outside opportunities, but they may also affect the costs of intramarital exchange. In this sense, part of this study's contribution is to explicitly recognize transaction costs within the household, and to illustrate that their influence on observed household behavior is economically significant.

The theoretical aspects of this study are a direct extension of Peters' work, and philosophically consistent with the transaction-cost approach articulated by Robert A. Pollak (1985). In section II, we review the relevant literature on divorce, which is a foundation for the transaction-cost model. In section III, we develop the model, which illustrates how the magnitude of the law's marginal effect will vary along with the transaction costs of intramarital exchange. In section IV, we use this model to interpret Wolfers' results. Our empirical approach to testing the model is detailed in section V, and the results of our tests are reported in section VI.

The model developed herein is consistent with the claim from Becker, Landes and Michael (1977) that "the majority of divorces results from uncertainty and unfavorable

outcomes." However, it contradicts their conclusion that "a couple dissolves their marriage if, and only if, their combined wealth when dissolved exceeds their combined married-wealth." This contradiction arises in this model because transferring wealth is not costless.

As articulated by Shelly Lundberg and Pollak (1996), "the prevalence of destructive or wasteful phenomena such as domestic violence and child abuse, as well as the demand for marriage counseling and family therapy, suggests we consider the possibility that family behavior is sometimes inefficient." This study is consistent with that perspective. Marriage may be "more efficient" than many other economic institutions, but only if the transaction costs of intramarital bargaining are relatively low. Yet those costs are never zero, so marriage is not perfectly efficient.

II. Extant Literature

The model established in section III is an extension of Peter's theoretical framework. For convenience, a generalized adaptation is presented herein. There are two periods, and in each period the payoff from marriage is a known, fixed-amount m. This gain is divided between the spouses, and an individual receives a fraction γ_I. This division implies the payoff is rivalrous, and process of dividing it may be subject to dissipating losses. However, dissipation is abstracted in this framework. So $\gamma_I + \gamma_S = 1$ where I and S denote the individual and spouse, respectively. The initial ratio of the shares is determined in the marriage market, but that process is abstracted herein.

An individual's outside opportunity is denoted as a_I. Outside opportunities are stochastic, and the first-period realization serves as the expectation for the second period. Individuals choose to marry in the first period if $\gamma_I m \geq a_I$, and remain single otherwise.

In the second period, new outside opportunities are realized. Suppose an individual's gain from marriage exceeds the outside opportunity. The individual may be willing to make a transfer to the spouse, denoted t_I, if doing so was necessary to remain married. However, the transfer would be constrained such that $t_I < \gamma_I m - a_I$. Alternatively, suppose $\gamma_I m < a_I$. If a transfer was necessary to engender divorce, the individual would be willing to do so, subject to the constraint $t_I \leq a_I - \gamma_I m$.

Peters employed Figure 1 to illustrate this configuration. Divorce is socially efficient when realizations occur above the solid diagonal ray. The outcome to remain married or divorce depends on the realization of outside opportunities, and potentially on the governing legal regime. Peters considered two regimes: mutual-consent and unilateral divorce. Mutual-consent laws grant divorces only when both spouses agree to it. This legal regime was prevalent in the United States prior to 1969. During the "no-fault" revolution from 1969 to 1985, twenty-nine states switched to unilateral regimes, which permit either spouse to dissolve the marriage without the other's consent.[2]

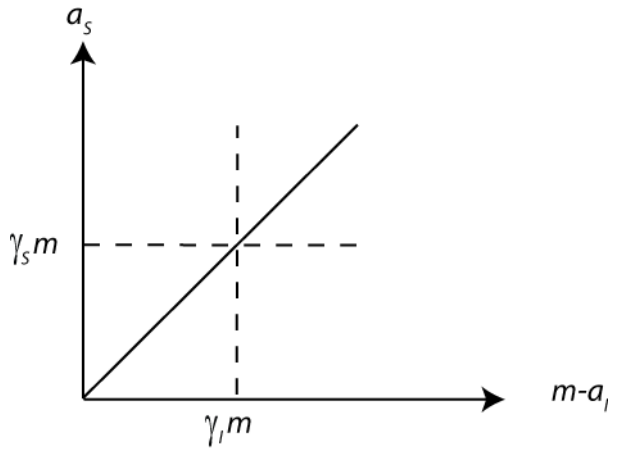

Figure 1: Distribution of outside opportunities at divorce relative to the value of marriage

Peters' extended the basic configuration to generate two distinct, theoretical models. The first assumed symmetric information between spouses. Symmetric information implies it

[2] The terms "unilateral" and "no-fault" are often used interchangeably, but they are not synonymous. There is a difference between consenting to a divorce and committing legal fault. This analysis explores the effect of changing the assignment of property rights, so it focuses on mutual-consent versus unilateral regimes, without regard to fault. Legal fault may have empirical ramifications, however, and that is addressed in sections V and VI.

is costless to renegotiate the gains from marriage, and therefore marriage contracts are efficient. In the symmetric model, both spouses will always agree on the outcome of either to divorce or remain married, so the probability of divorce is invariant to the law.

Her second model assumes bilateral asymmetric information, and that spouses respond to this asymmetry by entering into "fixed-wage" contracts that prohibit *ex post* bargaining. In other words, $t_1 \equiv 0$ regardless of the realization of outside opportunities. In the asymmetric model, individuals suffer from the possibility of inefficient marriages continuing or inefficient separations occurring, depending on which legal regime governs divorce. It predicts that unilateral divorce will result in higher divorce rates than mutual-consent regimes.

Peters' empirical analysis supported the symmetric model. Allen (1992) challenged Peters' treatment of states that changed legal regimes during the sample period. Peters (1992) countered, and Friedberg (1998) addressed the dispute by employing a panel of state-level divorce rates. Friedberg controlled for state and year-specific effects, and found that changes in legal regimes accounted for approximately one-sixth of the rise in divorce between 1968 and 1988.

Measured in divorces per thousand people, the divorce rate increased during this period from 2.2 to 4.8, or 118%. So Friedberg's result suggests the legal regime accounted for approximately a 20% increase in divorce rates. Wolfers (2006) argues that Friedberg's approach confounded preexisting trends in the divorce rate with the dynamic effects from

the policy shock, and concludes this increase dissipated over time. Yet he still finds that the divorce rate rose sharply immediately following the adoption of unilateral divorce.

Are Peters' theoretical models consistent with these empirical results? Our interpretation is that the symmetric model has been plainly refuted. The asymmetric model is not necessarily refuted, but the notion that spouses fail to negotiate is at odds with observed behavior. As noted by both Friedberg and Wolfers, the evidence suggests that reality lies between these two extreme cases, albeit closer to the symmetric case.

III. Transaction-Cost Model

Peters' two models are distinguished by the cost of renegotiating in the second period. Consider that these models may be connected by a third dimension, which represents the transaction costs impeding intramarital exchange. In Peters' symmetric model, the value along the third dimension is zero, while it assumes a prohibitively high $c^{\max}$ value for the asymmetric model. This notion is illustrated in Figure 2. Realizations occurring in the darkly-shaded regions of Figure 2 will end in divorce. Those in lightly-shaded areas will result in divorce only under unilateral regimes.

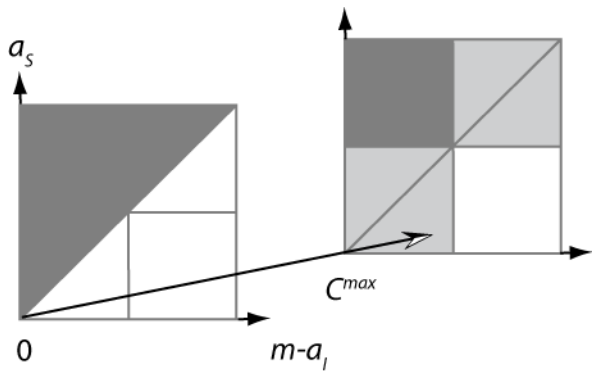

Figure 2: Three-dimensional representation of Peters' two models

This model deviates from Peters in its assumption that intramartial exchange during the second period may be costly. Specifically, a couple may transfer wealth between themselves only after bearing a transaction cost $c_{I,S} \geq 0$. The cost is assumed to vary for different combinations of individuals, hence the subscripts. Graphically, $c_{I,S}$ is represented as the third dimension in Figure 2. Recall that spouses may choose to

transfer wealth during the second period, and the constraint on the transfer can be expressed $t_I \leq |\gamma_I m - a_I| - c_{I,S}$. To complete the model, couples adhere to the outcomes enumerated in Table 1, where the decision to remain married or divorce is conditional on the realization of outside opportunities, transaction costs and the legal regime**.**

As transaction costs increase, the upper bound on the transfer decreases. This result holds under both legal regimes, but its effect on behavior varies with the law. Under mutual-consent regimes, one individual acting alone can sustain the marriage, so the purpose of transfers is to engender divorce. If the ability to transfer wealth is impaired, then some number of socially inefficient marriages may persist. Meanwhile, under a unilateral regime, one individual acting alone can terminate a marriage. Therefore, the purpose of the transfer is to sustain the marriage. As the ability to transfer wealth decreases, some socially efficient marriages will dissolve.

In reference to Figure 2, denote φ as the "divorce space," which is the area containing divorce realizations for a given level of transaction costs. Graphically, φ is a two-dimensional "slice" of the three-dimensional space. It is evident that φ is a function of $c_{I,S}$ and the legal regime. The sign of the partial derivative φ_C varies with the legal regime: $\varphi_C < 0$ under a mutual-consent regime, and $\varphi_C > 0$ under unilateral divorce.

The probability of divorce may be denoted as p, where $p = p\left(m, a_I, a_{S,}\varphi\left(c_{I,S}\right)\right)$. Partial derivatives are signed as follows: (*i*) $p_M \leq 0$; (*ii*) $p_A \geq 0$ for both the individual's and

spouse's outside opportunities; and (*iii*) p_{φ} must be nonnegative. The probability of divorce can be expressed as a function of $c_{I,S}$, where the partial derivative $p_C = p_{\varphi}\varphi_c$. Therefore, $p_C \leq 0$ under a mutual-consent and $p_C \geq 0$ under unilateral divorce. In other words, this model predicts that the probability of divorce will be negatively associated with transaction costs under mutual-consent. Yet it will be positively associated with transaction costs under unilateral regimes.

What occurs when the legal regime changes from mutual-consent to unilateral divorce? As a convention, assume changes in the law are unidirectional from mutual-consent to unilateral regimes. Denote the difference between the divorce spaces as $\Delta = (\varphi \mid unilateral) - (\varphi \mid mutual - consent)$. It is evident that $\Delta \geq 0$, and $\Delta = 0$ if and only if $c_{I,S} = 0$. Note that Δ is a function of function of $c_{I,S}$. Given the signs of φ_C under each legal regime, the partial derivative Δ_C is strictly positive.

The result that $\Delta_C > 0$ highlights that changes in the law are expected to affect individuals differently. When the legal regime changes, couples facing higher transaction costs will experience a larger increase in their divorce space, and therefore they are expected to face a larger increase in the probability of divorce. In other words, the marginal effect of the law increases along with transaction costs.

As an analytical exercise, consider two populations of married individuals. The first faces transaction costs of $\underline{c}$, the second faces $\overline{c}$, and $0 < \underline{c} < \overline{c} < c^{\max}$. Under a mutual-

consent regime, the low-cost group is expected to have a higher divorce rate, *ceteris paribus*. The opposite result is expected under unilateral divorce. When the legal regime switches from mutual-consent to unilateral divorce, the divorce rate is expected to increase for both populations, but the magnitude of the increase will be greater for the high-cost group. In other words, changes in the law will not affect all individuals equally. If $c \approx 0$, then the marginal effect of changing divorce law may approximate to zero. However, if $c \gg 0$, then the marginal effect of the change may be dramatic.

Table 1: Outcome of the second-period, given outside opportunities, transaction costs and the legal regime

Individual	Spouse	Legal regime	Outcome
$\gamma_I m \geq a_I$	$\gamma_S m \geq a_S$	Either	$t_I = t_S = 0$, remain married
$\gamma_I m < a_I$	$\gamma_S m < a_S$	Either	$t_I = 0$, divorce
$\gamma_I m > a_I$	$\gamma_S m < a_S$	Mutual-consent	If $a_S - \gamma_S m - c > \gamma_I m - a_I$ Then $t_S = \gamma_I m - a_I$ and divorce Otherwise, remain married
$\gamma_I m < a_I$	$\gamma_S m > a_S$	Mutual-consent	If $a_I - \gamma_I m - c > \gamma_S m - a_S$ Then $t_I = \gamma_S m - a_S$ and divorce Otherwise, remain married
$\gamma_I m > a_I$	$\gamma_S m < a_S$	Unilateral	If $\gamma_I m - a_I \geq a_S - \gamma_S m - c$ Then $t_I = a_S - \gamma_S m$ and remain married Otherwise, divorce
$\gamma_I m < a_I$	$\gamma_S m > a_S$	Unilateral	If $\gamma_S m - a_S \geq a_I - \gamma_I m - c$ Then $t_S = a_I - \gamma_I m$ and remain married Otherwise, divorce

IV. Interpretation of Existing Empirical Results

When the legal regime changed unilateral divorce, the divorce space increased if and only if transaction costs were positive. Friedberg concludes that 20% more realizations occurred in the larger unilateral divorce space. Given any reasonable assumption of the distribution of outside opportunities, this model's interpretation of Friedberg's finding is that the transaction costs of intramarital exchange are significantly greater than zero.

How might this model explain Wolfers' findings that the increase in the divorce rate dissipated over time? The transaction-cost framework developed herein is static, and it is not designed to address the dynamic evolution of behavior. However, if transaction costs significantly influence economic outcomes, then we should expect individuals to be aware of them, at least implicitly. This has implications for behavior both in the marriage market and during marriage.

In his study, Wolfers identified four possible explanations for the dissipation of the increase in divorce rates: (*i*) unilateral divorce caused "bad matches" to dissolve more quickly than they would under mutual-consent; (*ii*) fewer people married under unilateral regimes, which downwardly influenced the number of divorces per thousand people.[3] Meanwhile, those who did marry may have invested more resources prior to marriage into producing better matches; (*iii*) social norms around divorce evolved over time to

[3] Wolfers argues that "divorces per 1,000 persons" may an inappropriate metric, and analysis should focus on "divorces per 1,000 married persons."

produce the observed pattern; and (*iv*) perhaps the data simply regressed to the mean. This study suggests a fifth possibility: individuals responded to unilateral divorce by modifying behavior in ways that reduced the transaction costs of intramarital exchange.

Assuming most marriages are socially efficient, then most individuals have little incentive to mitigate transaction costs under mutual-consent regimes. The opposite holds under unilateral regimes. Over time, it may be that individuals' response to unilateral divorce was to alter behavior in ways that produced fewer transaction costs, which resulted in a lower divorce rate. This behavior may drive the average level of costs in the population towards the origin of Figure 2, which implies the resulting divorce rate may be closer to the socially efficient rate, *ceteris paribus.*

However, it is noteworthy that a "socially efficient divorce rate" is only synonymous with "social efficiency" contingent on the legal regime. This contingency is because decisions that produce marital gains also produce transaction costs. Suppose the presence of a child produces both a gain to marriage and transaction costs. Under mutual-consent, the couple is expected to have the child. Under unilateral divorce, the marginal benefit of the child may be trumped by the probability that additional transaction costs will lead to a socially inefficient divorce. If so, the couple will not have the child, even though the net gain to the marital payoff is positive. This illustrates that when the law changes, analyzing social efficiency becomes extremely muddled because few factors are held constant.

V. Empirical Methodology and Data

The key proposition of this model is that the magnitude of the law's marginal effect will vary across individuals. As a result, the law will influence which types of couples divorce. Given the available data, we focus testing the latter outcome, rather than the magnitude of the marginal effect directly. Therefore, our strategy is to identify systematic differences in the type of couples that divorce under each legal regime.

To isolate the effect of the law on divorce decisions, a commonly employed strategy has been to model the measure driving divorce as a latent variable, which determines the binary decision of whether to divorce. Equation (1) is a representative data-generating process (DGP) for this approach. Equation (2) represents the binary choice.

$$y^{*}_{i,s,t} = \alpha + X_{i,t}\beta + I^{Unilateral}_{i,s,t}\gamma + \varepsilon_{i,s,t} \qquad (1)$$

$$I^{Divorce}_{i,s,t} = 1 \text{ if } y^{*}_{i,s,t} \geq \overline{y} \qquad (2)$$

The transaction costs of intramarital exchange are not readily observable, but this analysis assumes they are correlated with observable variables. Our theoretical model predicts that transaction costs tip the likelihood of divorce in opposing directions, depending on the legal regime. As a result, equation (1) incorrectly specifies the true DGP, because the legal regime enters as a fixed intercept. Rather, the explanatory variables interacted with a state's unilateral status, as in equation (3).

$$y^{*}_{i,s,t} = \alpha + X_{i,t}\beta_1 + I_{s,t}^{Unilateral} X_{i,t}\beta_2 + \varepsilon_{i,s,t} \quad (3)$$

Coasean bargaining implies $\beta_1 = \beta_2$, and positive transaction costs would cause $\beta_1 \neq \beta_2$. Yet this straightforward test is problematic, because observable information that provides variation between couples is certainly endogenous, and strong instruments are difficult to identify. Consequently, we propose an alternative approach.

Our model predicts divergent transaction costs among divorcing couples across legal regimes. Therefore, inasmuch as transaction costs are correlated with observable traits, we expect that identifiable characteristics of divorcing couples in a unilateral regime will systematically vary from couples divorcing under a mutual consent regime. Specifically, divorcing couples in unilateral states will exhibit higher levels of transaction costs relative to those in mutual-consent states, *ceteris paribus*. Assuming couples neither marry nor invest in marriage with the intention to ultimately divorce, we treat all matching or marital variables as predetermined prior to the decision to divorce. This removes any potential endogeneity in our estimation.

Our theoretical model assumes individuals vary in the transaction costs they face, but are homogenous in their ability to bear those costs. In reality, individuals vary in both the costs they face and their ability to bear them. There is a difference between the "gross" and "net" levels of transaction costs within a given household. Across households, those with the same gross level do not necessarily have the same net level. This complexity acknowledged, this study proceeds on two assumptions: (*i*) net transaction costs within a

household are positively correlated with the number of children present in the household; and (*ii*) net transaction costs are negatively correlated with the duration of the marriage.

These assumptions merit further justification. Regarding children, sundry evidence exists that their presence positively increases the costs of intramarital bargaining. Strauss (1978) explicitly recognizes that the presence of children influence the bargaining process. Madden-Bulman (1981) interviewed married women about the causes of marital disputes, and reported that "discipline of children" and "division of household labor and child care" were two of the top four responses. A vast literature in sociology suggests that spouses engage in repeated negotiation over the allocation of household responsibilities. Since children increase those responsibilities, it is reasonable to assume their presence serves to increase the costs of intramarital bargaining.

In regards to the duration of a marriage, we assume that marital negotiation can be described as a "learning-by-doing" process. As the duration increases, the parties become more efficient negotiators. *Ceteris paribus*, this implies the costs that individuals devote to negotiation will decrease as the duration increases.

Given these assumptions, we expect to be able to predict under which legal regime a divorce occurred, given the vector of observable characteristics. The prediction mechanism is represented by the probit model in equation (4), which represents the likelihood that a specified marriage dissolved under a unilateral regime.

$$\Pr(\text{Unilateral State}) = \alpha + X_{i,t}\beta + \nu_t + \varepsilon_{i,t} \qquad (4)$$

The vector of observable variables and a year fixed effect are denoted $X_{i,t}$ and ν_t, respectively. Our prediction is that characteristics associated with higher levels of transaction costs (more children and shorter marriages) will be positively associated with unilateral divorce. Note that our ability to predict the prevailing legal regime is possible only if the DGP specified in equation (3) is correct. Equation (1) does not predict systematic variation in the characteristics of divorcing couples, but rather a significant difference in the number of divorces only.

Our empirical test employs public-use divorce data collected from the National Center for Health Statistics (NCHS), spanning the years 1968 to 1985. In Table 3, we present the descriptive means for couples dissolving their marriages in our data, segmented by the legal regime at the time of divorce (estimated standard deviations in parentheses). This dataset captures the duration of the "no-fault revolution," and the data are representative samples of all divorce certificates decreed during the year for each participating state.[4] The Divorce Registration Area state membership varied over this time period (states both entered and exited the system), beginning with 26 states and expanding to 31 states by 1985. Data is available on a total of 32 unique states over the period: 13 states are classified as "mutual consent" for all observations, 3 states are "unilateral" for all

[4] See the NCHS website at http://www.cdc.gov/nchs/products/elec_prods/subject/divorce.htm for supporting documentation on this data. Additional documentation, and the data itself, are available through the National Bureau of Economic Research at http://www.nber.org/data/marrdivo.html.

observations, and 16 transitioned from "mutual consent" to "unilateral" during the period of observation.[5]

Given the age of the data and the certificate-based process of sampling for data collection, missing observations pose a problem on variables with no method to reliably impute the values. To minimize the loss of observations while obtaining the most informative data, we employ three different samples for the study: (*i*) the full data sample; (*ii*) all observations that have all pertinent demographic controls; and (*iii*) all observations with both demographic and separation variables.[6]

The empirical literature on divorce illustrates that results can be sensitive to a number of issues. To demonstrate the robustness of our findings, we conduct three additional tests of our hypothesis. First, we consider differing definitions of "unilateral" divorce. Second, we control for household sorting among the underlying population of married couples. Third, we confirm the results persist beyond the transition period between legal regimes. Each of those issues is discussed in detail below.

Friedberg (1998) classified legal regimes as "unilateral" under three different rules: by decree, by decree after separation requirements are met, and by incorporating fault in divorce settlements. Unless specified, the unilateral classification employed in this analysis is unilateral by decree, which is the most commonly held classification.

[5] The U.S. Virgin Islands are included in the original NCHS data, but are omitted from our analysis.
[6] Information on the interval length between separation and divorce was not recorded until 1970; thus, the third sample omits the years 1968 and 1969 from the sample.

However, Allen (1992) illustrated that the coefficient associated with the legal regime is sensitive to the definition of "unilateral divorce." To address this concern, we order the legal regimes according to the degree of restriction on a spouse initiating divorce. From most to least restrictive: (*i*) mutual-consent; (*ii*) unilateral with separation requirements; (*iii*) unilateral without separation; and finally (*iv*) unilateral without fault entering into settlement decisions.

We classify states into these four categories, and employ an ordered probit model to predict the average probability of a couple divorcing in each of these progressively liberal legal regimes. We assume that transaction costs monotonically increase with the restrictiveness of the divorce law, since more bargaining (or more intense bargaining) is required. Therefore, we expect that characteristics correlated with transaction costs to produce estimates consistent with earlier binomial tests.

Beyond the definition of unilateral divorce, another concern may be that whatever pattern observed in the data is simply due to regional differences. A given state's unilateral status is correlated with surrounding states, and intact couples may not be distributed uniformly across legal regimes. Therefore, systematic differences between legal regimes may simply reflect differences in the underlying population.

Consider equations (5) and (6), where ϕ is a sorting parameter on intact couples that determines the state in which they choose to reside, and δ determines divorce using the same variables. If equations (5) and (6) were the true DGP and we failed to address this,

we would attribute the magnitude of ϕ to our estimate of β (the estimated parameter from equation (4)).

$$\Pr(\text{Reside in Unilateral State}) = \alpha_1 + X_{i,t}\phi + \nu_{1,t} + \varepsilon_{1,i,t} \qquad (5)$$

$$\Pr(\text{Divorce}) = \alpha_2 + X_{i,t}\delta + \nu_{2,t} + \varepsilon_{2,i,t} \qquad (6)$$

To address this issue, we duplicate the prediction model on representative household data. The data for this test is obtained from the 1980 United States Census, where we isolate all head-of-household married couples in the same states participating in the original NCHS data during 1980. The census is a 1-in-100 representative sample of households across the United States, and it provides the means to determine whether the variation across legal regimes is one that arises due to transaction costs influencing a couple's decision to divorce or whether this is due to sorting across states alone.

Finally, it is possible that any observed systematic variation is due to the dynamic transition from one legal regime to the other. To eliminate this possibility, we replicate the analysis on NCHS divorce data spanning the years 1991-1995.[7] Testing for the presence of transaction costs in these data provide another important check on our model's predictions.

[7] NCHS stopped collecting micro-level divorce data after 1995. However, the data from these latter years of the program are more than five years removed from the transition period ending in 1985.

Table 3: Descriptive Statistics of Divorcing Couples, by Legal Regime

Legal regime governing divorce	Mutual Consent	Unilateral
Number of children [a]	1.019 (1.194)	1.059 (1.192)
Marriage duration [a]	9.256 (8.279)	8.481 (8.004)
Demographic Variables [b]		
Husband's age at marriage	25.484 (8.178)	25.745 (8.627)
Wife's age at marriage	22.869 (7.431)	23.078 (7.804)
Age difference (husband - wife)	2.619 (4.990)	2.671 (5.055)
Husband married previously	0.225 (0.418)	0.279 (0.449)
Wife married previously	0.219 (0.414)	0.277 (0.447)
Husband white	0.860 (0.347)	0.907 (0.290)
Husband African American	0.128 (0.334)	0.053 (0.224)
Wife white	0.861 (0.345)	0.905 (0.293)
Wife African American	0.125 (0.331)	0.049 (0.215)
Mixed-race couple	0.016 (0.124)	0.034 (0.182)
Separation Variable [c]		
Interval from separation to divorce (in months)	24.230 (37.178)	15.047 (24.257)
Observations in unrestricted sample	2,531,727	867,859
Observations with demographic variables	2,120,252	613,714
Observations with demographic and separation variables	1,640,967	451,053

a: Measured in unrestricted sample for all non-missing observations.
b: Measured among observations where all variables are non-missing.
c: Measured among observations where all separation and demographic variables are non-missing.

Source: NCHS data

VI. Empirical Analysis

The results of our prediction model are reported in Table 4. The dependent variable assumes the value of one if the divorce occurred in a unilateral state. Each cell reports the marginal effect associated with the change in the explanatory variables at the sample mean. In both the restricted and unrestricted samples, we see that the presence of children is positively associated with divorce in unilateral states, and that "marriage duration" is negatively associated with a unilateral divorce regime. Given our assumptions, these results are consistent with the transaction-costs model.

As mentioned previously, a state's unilateral status is correlated with surrounding states. Thus, adding regional fixed effects to detect within-region differences is an important robustness check on our findings. Both the unrestricted and restricted samples are robust to the inclusion of these regional effects. The association between children and unilateral divorce peaks around two to three children, and actually turns negative in the presence of five or more children. This challenges our assumption that the net level of transaction costs monotonically increase with respect to children, at least when the number of children in the household becomes exceptionally high. Our interpretation is that individuals who choose to have five or more children likely are willing and able to bear a high level of transaction costs (thus resulting in a relatively low net level).

The marginal effects reported in Table 4 are small in magnitude. While the results are statistically significant, their economic significance may not be entirely clear. Recall that

our model is a prediction of under which legal regime a divorce occurred, given the vector of observable control variables. Transaction costs are economically significant if and only if they have a high degree of predictive power. Therefore, the fundamental issue is how the presence of transaction costs influences our ability to predict the legal regime.

To address this question, reconsider the populations facing transaction costs of $\underline{c}$ and $\overline{c}$, where $0 < \underline{c} < \overline{c} < c^{\max}$. In regards to children, Table 4 reports that the marginal prediction peaks in the presence of two to three children. Assume couples with two children represent the high-cost group, and those without children constitute the low-cost group. For marriage duration, the marginal prediction plateaus near ten years of marriage. Along the dimension of duration, assume this represents the low-cost group, and marriages lasting less than one year represent the high-cost group.

Given those assumptions, we can calculate the percentage difference in conditional probability along both dimensions, as expressed in equation (7). In Table 5, we report those results from 1970 to 1985. They are calculated on an annual basis because the original estimates were calculated using year fixed effects. The magnitude of the results is sensitive to the proportion of unilateral divorces occurring in the sample, which increased over time as more states adopted unilateral divorce (hence the downward trend). However, it is evident that both the presence of children and the duration of marriage had an economically significant effect, in that they strongly influence our ability to predict the legal regime.

$$\frac{\Pr(Unilateral \mid \overline{c}) - \Pr(Unilateral \mid \underline{c})}{\Pr(Unilateral \mid \underline{c})} \qquad (7)$$

As discussed in section V, divorce analysis can be sensitive to the definition of "unilateral." To confirm our results, we test whether these findings hold across the alternative classifications. In Table 6, we explicitly address the subtle gradations inherent in the unilateral divorce classifications through the use of an ordered probit model. Higher values in the dependent variable here indicate less restrictive divorce laws: "mutual-consent" = 0, while "unilateral by decree" = 3. Note that since the type of unilateral divorce is strongly correlated with other states in the region, we do not employ regional fixed effects in this specification.

In regards to children, the results in Table 5 are consistent with those from Table 4. However, marriage duration shows a positive relationship with unilateral divorce. At a minimum, this inconsistency casts doubt on this variable's correlation across alternate categorizations of unilateral divorce. Given the myriad of possible explanations, we are hesitant to draw any strong conclusion. Yet this result suggests the need for further investigation.

The presence of systematic differences does not necessarily imply that transaction costs are the cause. Consider that these differences may reflect predetermined sorting across the states. Table 7 reports the results of a test designed to address this concern. We

stacked 1980 census data with the 1980 divorce data, and interact the explanatory variables with the data source, such that the resulting X matrix was a block matrix. Therefore, differences on prior household sorting are estimated separately with the census data while differences in divorcing couples are estimated with the divorce data. These results lead us to reject the notion that the systematic variation is due to any type of "geographic sorting" behavior, or any observable differences in the underlying population.

The results in Table 7 clearly support the predictions of the transaction-cost model. If the differences were predetermined in the population, the resulting parameter estimates from the Census data should show little difference from those of the NCHS data. Yet in the NCHS data, the presence of children is significantly correlated with divorce in unilateral regimes. In the census data, that correlation is negative. Coupled with the strong and consistent evidence that the "number of children" is associated with couples divorced under unilateral regimes, these results suggest that these differences in behavior arose due to the divorce law.

Our final test of the model uses more current divorce data, spanning 1991-1995. This time period post-dates the "no-fault revolution" by over five years. Our model does not suggest that observed differences were a result of the transition between legal regimes, but rather due to the difference in the static costs of divorce under a given legal regime. Table 8 presents the results of similar predictions using this data. In both the unrestricted and the restricted samples, the current data exhibits a relationship generally consistent in

sign and magnitude with results reported earlier in Table 4. This holds for both the number of children and the duration of marriage, and these results are robust to the inclusion of demographic controls and the inclusion of regional fixed effects.

In summary, a key prediction of our theoretical model is that transaction costs will tip the likelihood of divorce in opposing directions, depending on the legal regime. Assuming that the presence of children is positively correlated with transaction costs, the evidence is consistent with our prediction. We illustrate these results cannot be attributable to pre-existing differences within the underlying population, and that these correlations have persisted well beyond the "no-fault" transition period.

Table 4: Estimated Marginal Effects Predicting Unilateral Divorce Regime

	Unrestricted Sample		Restricted Sample		
Demographic Controls	--	--	--	√	√
Year Fixed Effects	√	√	√	√	√
Region Fixed Effects	--	√	--	--	√
1 child	0.004**	0.006**	0.012**	0.020**	0.018**
	(0.001)	(0.001)	(0.001)	(0.001)	(0.001)
2 children	0.023**	0.027**	0.033**	0.041**	0.041**
	(0.001)	(0.001)	(0.001)	(0.002)	(0.001)
3 children	0.019**	0.028**	0.026**	0.033**	0.041**
	(0.002)	(0.002)	(0.002)	(0.002)	(0.002)
4 children	0.002	0.020**	0.009**	0.016**	0.032**
	(0.003)	(0.003)	(0.003)	(0.003)	(0.003)
5-7 children	-0.027**	-0.005	-0.022**	-0.015**	0.008
	(0.004)	(0.004)	(0.004)	(0.005)	(0.004)
8 or more children	-0.059**	-0.022	-0.078**	-0.068**	-0.019
	(0.015)	(0.016)	(0.017)	(0.018)	(0.018)
Married 1 year	-0.034**	-0.031**	-0.044**	-0.031**	-0.039**
	(0.002)	(0.002)	(0.003)	(0.003)	(0.003)
Married 2 years	-0.047**	-0.040**	-0.070**	-0.051**	-0.062**
	(0.002)	(0.002)	(0.003)	(0.003)	(0.003)
Married 3-5 years	-0.060**	-0.050**	-0.092**	-0.067**	-0.083**
	(0.002)	(0.002)	(0.002)	(0.002)	(0.002)
Married 6-9 years	-0.072**	-0.058**	-0.111**	-0.082**	-0.099**
	(0.002)	(0.002)	(0.002)	(0.002)	(0.002)
Married 10-19 years	-0.085**	-0.064**	-0.127**	-0.092**	-0.106**
	(0.002)	(0.002)	(0.002)	(0.002)	(0.002)
Married 20-29 years	-0.090**	-0.058**	-0.127**	-0.086**	-0.091**
	(0.002)	(0.002)	(0.002)	(0.003)	(0.002)
Married 30-39 years	-0.087**	-0.050**	-0.122**	-0.073**	-0.078**
	(0.003)	(0.003)	(0.003)	(0.004)	(0.003)
Married 40 or more years	-0.090**	-0.045**	-0.134**	-0.081**	-0.077**
	(0.006)	(0.006)	(0.006)	(0.007)	(0.007)
Observations	3,399,586	3,399,586	2,733,966	2,733,966	2,733,966

* significant at 5%; ** significant at 1%. Robust standard errors in parentheses.

Note: Indicators for missing variables were created and used for both the number of children and marriage duration in the unrestricted sample. Observations with missing information are dropped in the restricted sample. Demographic controls include race, age at marriage, and previously married indicator variables for both spouses, and the couple's age difference.

Table 5: Percentage difference in conditional probability between “high-cost” and “low-cost” populations

Year	Marriage Duration	Number of Children
1970	88.6%	23.4%
1971	90.2%	23.8%
1972	85.4%	22.7%
1973	68.9%	18.7%
1974	67.0%	18.2%
1975	60.4%	16.6%
1976	59.0%	16.2%
1977	58.7%	16.1%
1978	61.4%	16.8%
1979	60.6%	16.6%
1980	59.7%	16.4%
1981	60.5%	16.6%
1982	62.0%	17.0%
1983	62.5%	17.1%
1984	62.3%	17.0%
1985	62.2%	17.0%

Table 6: Ordered Probit Predicting Increasingly Liberal Unilateral Divorce Regimes

	Restricted Sample	
Demographic Controls	√	√
Separation Controls	--	√
Year Fixed Effects	√	√
Region Fixed Effects	--	--
1 child	0.051**	0.045**
	(0.004)	(0.004)
2 children	0.086**	0.064**
	(0.004)	(0.004)
3 children	0.058**	0.033**
	(0.006)	(0.006)
4 children	0.01	-0.014
	(0.009)	(0.009)
5-7 children	-0.082**	-0.098**
	(0.014)	(0.014)
8 or more children	-0.312**	-0.334**
	(0.060)	(0.060)
Married 1 year	0.093**	0.002
	(0.008)	(0.008)
Married 2 years	0.116**	0.037**
	(0.008)	(0.008)
Married 3-5 years	0.121**	0.061**
	(0.007)	(0.007)
Married 6-9 years	0.114**	0.079**
	(0.007)	(0.008)
Married 10-19 years	0.091**	0.072**
	(0.008)	(0.008)
Married 20-29 years	0.102**	0.089**
	(0.008)	(0.008)
Married 30-39 years	0.134**	0.142**
	(0.011)	(0.011)
Married 40 or more years	0.080**	0.137**
	(0.023)	(0.023)
Observations	2,092,020	2,092,020

* significant at 5%; ** significant at 1%. Robust standard errors in parentheses.

Note: Demographic controls include race, age at marriage, and previously married indicator variables for both spouses, and the couple's age difference.

Table 7: Estimates Predicting Unilateral Regime from Stacked Divorce and Census Data, 1980 only

	NCHS estimates	Census estimates	F-test for equality	NCHS estimates	Census estimates	F-test for equality
Demographic Controls	√			√		
Region Fixed Effects	--			√		
1 child	0.027**	-0.113**	164.15	0.023**	-0.083**	85.7
	(0.005)	(0.010)	0.000	(0.005)	(0.010)	0.000
2 children	0.058**	-0.117**	229.78	0.056**	-0.088**	142.23
	(0.005)	(0.010)	0.000	(0.005)	(0.011)	0.000
3 children	0.066**	-0.114**	194.79	0.064**	-0.084**	122.33
	(0.008)	(0.010)	0.000	(0.008)	(0.011)	0.000
4 children	0.043**	-0.101**	79.48	0.041**	-0.072**	46.09
	(0.013)	(0.010)	0.000	(0.013)	(0.011)	0.000
5-7 children	0.061**	-0.092**	44.89	0.059**	-0.066**	29.42
	(0.021)	(0.010)	0.000	(0.021)	(0.011)	0.000
8+ children	0.033	-0.074**	1.19	0.033	-0.056**	0.84
	(0.103)	(0.011)	0.257	(0.102)	(0.012)	0.361
Married 1 year	-0.076**	-0.003	35.58	-0.070**	0.001	30.95
	(0.009)	(0.008)	0.000	(0.009)	(0.009)	0.000
Married 2 years	-0.109**	0.006	91.83	-0.101**	0.008	76.05
	(0.008)	(0.008)	0.000	(0.009)	(0.009)	0.000
Married 3-5 years	-0.131**	-0.012	136.32	-0.117**	-0.005	107.88
	(0.007)	(0.007)	0.000	(0.008)	(0.007)	0.000
Married 6-9 years	-0.162**	-0.01	231.56	-0.147**	0.002	199.01
	(0.007)	(0.007)	0.000	(0.007)	(0.007)	0.000
Married 10-19 years	-0.176**	-0.016*	264.87	-0.158**	-0.002	222.89
	(0.006)	(0.007)	0.000	(0.007)	(0.007)	0.000
Married 20-29 years	-0.163**	-0.025**	169.54	-0.141**	-0.006	144.43
	(0.007)	(0.007)	0.000	(0.008)	(0.007)	0.000
Married 30-39 years	-0.154**	-0.024**	91.96	-0.131**	-0.005	78.16
	(0.010)	(0.007)	0.000	(0.011)	(0.007)	0.000
Married 40+ years	-0.120**	-0.009	16.06	-0.112**	0.007	18.03
	(0.024)	(0.007)	0.000	(0.024)	(0.007)	0.000
Observations	485,072			485,072		

* significant at 5%; ** significant at 1%. Robust standard errors in parentheses.

Note: Demographic controls include race, age at marriage, and previously married indicator variables for both spouses, and the couple's age difference.

Table 8: Predicting Unilateral Regime Outside of Transition Years, NCHS Data, 1991-1995

	Unrestricted Sample	Restricted Sample	
Demographic Controls	--	√	√
Year Fixed Effects	√	√	√
Region Fixed Effects	--	--	√
1 child	0.010**	0.020**	0.016**
	(0.002)	(0.002)	(0.002)
2 children	0.043**	0.053**	0.047**
	(0.002)	(0.002)	(0.002)
3 children	0.053**	0.063**	0.054**
	(0.003)	(0.004)	(0.004)
4 children	0.072**	0.085**	0.059**
	(0.007)	(0.007)	(0.007)
5-7 children	0.055**	0.063**	0.009
	(0.012)	(0.013)	(0.012)
8 or more children	0.125**	-0.133**	-0.109**
	(0.048)	(0.005)	(0.005)
Married 1 year	-0.038**	-0.034**	-0.027**
	(0.005)	(0.005)	(0.005)
Married 2 years	-0.068**	-0.064**	-0.050**
	(0.004)	(0.005)	(0.005)
Married 3-5 years	-0.105**	-0.101**	-0.075**
	(0.004)	(0.004)	(0.004)
Married 6-9 years	-0.129**	-0.126**	-0.091**
	(0.004)	(0.004)	(0.004)
Married 10-19 years	-0.141**	-0.139**	-0.100**
	(0.004)	(0.004)	(0.004)
Married 20-29 years	-0.142**	-0.139**	-0.098**
	(0.004)	(0.004)	(0.004)
Married 30-39 years	-0.138**	-0.132**	-0.090**
	(0.005)	(0.006)	(0.006)
Married 40 or more years	-0.128**	-0.118**	-0.082**
	(0.009)	(0.010)	(0.010)
Observations	641,164	594,034	594,034

* significant at 5%; ** significant at 1%. Robust standard errors in parentheses.

Note: Demographic controls include race, age at marriage, and previously married indicator variables for both spouses, and the couple's age difference.

VII. Conclusion

This study introduces a transaction-cost framework of intramarital bargaining. In accordance with the correspondence principle, it replicates the predictions from Peters (1986), and it is consistent with empirical analyses inspired by her work. The fundamental question explored in this literature is whether the Coase Theorem should be applied to martial relations. The analysis herein suggests it should not.

The principal contribution of this study is to demonstrate that the marginal effect of the law is positively associated with the transaction costs of intramarital exchange. Since those costs vary across a population, so does the marginal effect of the legal regime governing divorce. As a result, the law governing divorce will influence which type of couples divorce, *ceteris paribus*.

The model predicts that divorcing couples in unilateral states should exhibit higher levels of transaction costs compared to those in mutual-consent states. This analysis employs a representative data set of divorces in the United States from 1968 to 1985. It tests this implication via a probit specification to predict under which legal regime a given divorce occurred. The empirical results strongly support the transaction cost model, finding that the characteristics of divorcing couples are influenced by the prevailing legal regime.

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