Transaction Costs, Legal Regimes and Divorce*

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Abstract

This analysis extends H. Elizabeth Peters' "Marriage and Divorce: Information Constraints and Private Contracting" (1986). Peters presented two models of intramarital bargaining as mutually-exclusive archetypes. 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. Our model implies divorce laws have a marginal effect on the probability of divorce, and the marginal effect's magnitude is a positive function of transaction costs. As an empirical test, we analyze the influence of marriage and family therapists on divorce across legal regimes. Evidence is consistent with the theory. *JEL*: D13, D23, J12, K36

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I. Introduction

In the United States, thirty-one states have adopted some form of unilateral divorce, which permits either spouse to dissolve a marriage without the other's consent. Nineteen states and the District of Columbia have retained some form of mutual consent laws, which limit divorce to situations in which either both spouses agree to divorce, or one spouse bears significant costs to terminate the marriage (such as proving legal fault or adhering to an onerous separation agreement). From 1969 to 1985, twenty-nine states switched from mutual consent to unilateral divorce, which effectively altered the assignment of property rights within the household.

Economists have used this variation test the applicability of the Coase Theorem (1960) to intrahousehold relations. If rates were invariant to the law, then by implication the Coase Theorem applies and households are Pareto efficient. A consensus has emerged that divorce rates increased with the adoption of unilateral divorce, and the magnitude of that increase was approximately 10%.[1] However, Justin Wolfers (2006) has argued that the increase appears to have dissipated over time and divorce rates in unilateral states are approaching pre-1969 levels. In the absence of a coherent theoretical framework, this evidence is confusing. Does the Coase Theorem apply to households, or not?

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[1] The law's effect on divorce rates was first explored by Peters (1986), who concluded that divorce was invariant to the law. Allen (1992) identified a number of issues with her analysis, and he found a positive association between divorce rates and unilateral regimes. Peters (1992) countered, and Friedberg (1998) addressed the dispute by employing a panel of state-level divorce rates. Friedberg reported that the change in the law accounted for approximately 17% of the rise in divorce between 1968 and 1988. Wolfers (2006) extended Friedberg's panel, and he argued that Friedberg's approach confounded preexisting trends. Wolfers concluded that divorce rates did increase, but roughly only half what Friedberg reported. Wolfers' estimate of the initial increase is consistent with those from Gruber (2004).

We develop a theoretical model to interpret existing empirical evidence, and it substantiates the view that intrahousehold transaction costs are significant. Our model implies the law has a marginal effect on divorce, and the magnitude of the marginal effect is a positive function of transaction costs. Transaction costs are expected to vary across couples in a population, so the law's effect also varies. This notion has been suggested in the literature, but principally as an aside. Allen (1998) and Rasul (2006) make this point, and the most dramatic example may come from Stevenson and Wolfers (2006), who found that the adoption of unilateral divorce was followed by significant declines in the rates of domestic violence, female suicide and females murdered by their partners. Our contribution is to explicitly identify what others have implicitly suggested: the marginal effect of the law varies across couples, and this variation is due to transaction costs.

Peters (1986) presented two models of intramarital bargaining as mutually-exclusive archetypes. We demonstrate those models may be viewed as special cases of a single, three-dimensional framework. The third dimension represents the transaction costs of intramarital exchange. We test our theory by examining the relationship between marriage and family therapists (MFT) and divorce rates across legal regimes. Empirically, the presence of MFT on divorce varies across legal regimes. In mutual consent states, they appear to engender divorce. In unilateral states, their effect on divorce outcomes is significantly smaller. Assuming part of a MFT's role is to help couples reach stronger ("more binding") contracts than they could otherwise, then our empirical results are consistent with the theory.

II. Transaction Costs

As used in this study, the phrase "transaction costs" refers to the costs that individuals incur in order to gain at the expense of their partner.[2] To clarify, individuals within a household may devote effort to increase the size of the household's total output, of which they receive a share. If the marital output is not a pure public good, they may also compete against each other in order to increase their share of the output that exists. These strategies are not mutually exclusive, and individuals are expected to pursue them both. Yet if individuals pursue the second strategy, the transaction costs of intrahousehold exchange are positive and the household is not Pareto efficient.

Why do spouses engage in costly intrahousehold competition? Lundberg and Pollak (1993) suggest it may arise from the inability to form binding marriage contracts, which can be illustrated by a simple "prisoner's dilemma." In the absence of a binding contract to avoid competition, iterative dominance causes both spouses to compete. 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 (and by extension, costless transacting) when modeling intrahousehold exchange. However, in reality, acquiring genuinely symmetric information is

[2] For a thorough discussion of this definition, see Barzel (1997). As noted by Allen (1992), a good deal of confusion exists over the Coase Theorem, which he attributed in part to the absence of a definition by Coase. Similarly, the phrase "transaction costs" may also cause confusion since different authors may use different definitions, or the same author may use a different definition in different contexts. In an attempt to avoid that confusion, we define the phrase explicitly to represent the costs of competitive effort within the household.

prohibitively costly.[3] Spouses may have intimate knowledge of each other's 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. Spouses who specialize will devote their time to different activities, often 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.

Macher and Richman (2008) provide a comprehensive overview of the empirical literature on transaction costs, spanning approximately 900 papers across multiple disciplines. They identify only one paper that explicitly considers the influence of transaction costs on intrahousehold behavior (Hamilton (1999)). However, while most empirical studies fail to acknowledge transaction costs, many implicitly incorporate them. Peters (1986) controlled for a woman's age, for example, but she did not articulate why that variable should matter. It may seem obvious that an individual's age will influence his or her outside opportunity. Yet it should also be expected to influence the level of competition within the household, and therefore it may influence divorce through multiple channels.

[3] Consider that most transactions have a large—possibly infinite—number of potential outcomes. Negotiating each of these outcomes *ex ante* may be prohibitively costly to both parties.

III. Model

Following Peters (1986), individuals play a two-period game. In each period, the payoff from marriage is a known, fixed-amount m. This gain is divided between the spouses. An individual receives a fraction γ_I, and the initial ratio of the shares is determined in the marriage market. 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.

Peters employed a version of Figure 1 to illustrate this configuration. Divorce is socially efficient when realizations occur above the solid diagonal ray. In the second period, new outside opportunities are realized. If $\gamma_I m > a_I$, an individual may be willing to make a transfer to the spouse, denoted t_I, if doing so was necessary to remain married. Alternatively, suppose $\gamma_I m < a_I$. If a transfer was necessary to facilitate divorce, he would be willing to engage in that transaction. The transfer is necessarily constrained, and in Peters' framework the constraint is $t_I \leq |\gamma_I m - a_I|$.

Peters' extended the basic configuration to generate two distinct theoretical models. The first assumed symmetric information between spouses. Symmetric information implies it is costless to renegotiate the gains from marriage, and therefore marriage contracts are efficient.[4] 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 assumed bilateral asymmetric information which compelled spouses to enter into "fixed-wage" contracts. In other words, $t_I \equiv 0$. In the asymmetric model, individuals suffer from the possibility of inefficient marriages continuing or inefficient separations occurring, depending on the legal regime.[5] In contrast with the symmetric model, the asymmetric model predicts that unilateral divorce will result in higher divorce rates than mutual consent regimes.

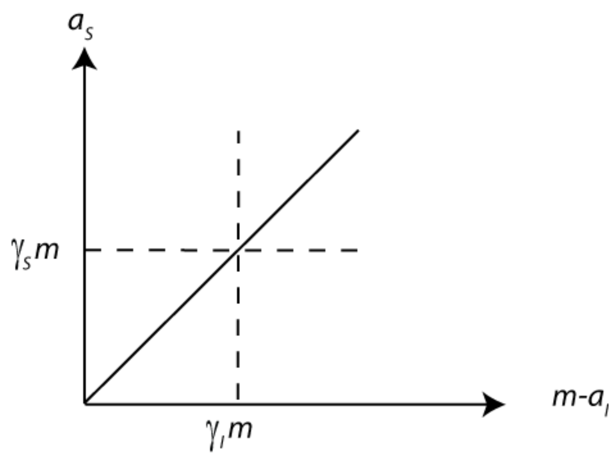

Figure 1: Distribution of outside opportunities at divorce relative to the value of marriage. This two-dimensional figure holds transaction costs constant.

[4] This is frequently referred to as "Coasean bargaining." However, given Coase's personal view on the matter, that eponym seems inappropriate. Coase (1981) writes, "[W]hile consideration of what would happen in a world of zero transaction costs can give us valuable insights, these insights are, in my view, without value except as steps on the way to the analysis of the real world of positive transaction costs. We do not do well to devote ourselves to a detailed study of the world of zero transaction costs, like augurs divining the future by the minute inspection of the entrails of a goose."

[5] This model adopts an extreme view of mutual consent divorce, in that both spouses must consent. In reality, channels exist for one spouse to engender divorce in mutual consent states (principally legal fault and separation). However, those channels are expensive, and we assume they are prohibitively costly in most cases. Therefore, those channels are abstracted from this analysis.

Our model deviates from Peters' in that a couple may transfer wealth between themselves only after bearing a transaction cost $c_{I,S} \geq 0$. The cost may be viewed as the couple's expected loss that results from competing over share. This competition arises because they are unable to form a binding agreement. The cost is assumed to vary for different combinations of individuals, hence the subscripts. The constraint on the transfer can be expressed $t_I \leq |\gamma_I m - a_I| - c_{I,S}$. 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.

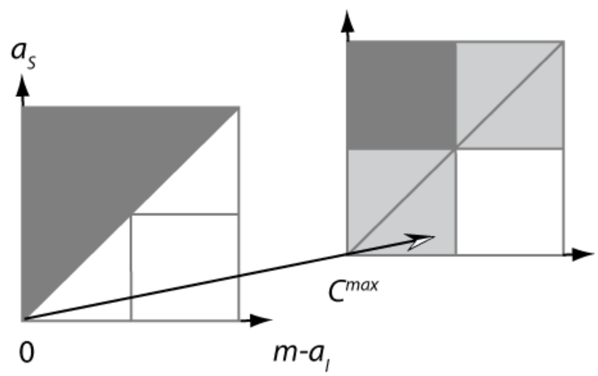

Figure 2: Three-dimensional representation of Peters' two models. Realizations in the darkly-shaded areas will always result in divorce, while those in the lightly-shaded area divorce only in unilateral regimes.

To complete our model, a couple's decisions to remain married or divorce is governed by the rules enumerated in Table 1. As transaction costs increase, the upper bound on the transfer decreases. This 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. Under a unilateral regime, one individual acting alone can terminate a marriage, and the purpose of the transfer is to sustain the marriage. As the ability to transfer wealth decreases, some socially efficient marriages will dissolve.

Denote φ as the "divorce space," which is the two-dimensional area containing realizations of outside opportunities that will result in divorce for a given couple. The divorce space is a function of both transaction costs and the legal regime. The sign of the partial derivative φ_C varies with the law: $\varphi_C < 0$ under mutual consent, and $\varphi_C > 0$ under unilateral divorce.[6] Changes in the law have been unidirectional from mutual consent to unilateral regimes, so the difference between divorce spaces across regimes is $\Delta = (\varphi \mid unilateral) - (\varphi \mid mutual-consent)$. It is evident that $\Delta \geq 0$ ($\Delta = 0$ if and only if $c_{I,S} = 0$). The difference Δ is a function of function of $c_{I,S}$, and the partial derivative Δ_C is strictly positive. In other words, the magnitude of the law's marginal effect is a positive function of transaction costs. Since the level of transaction costs vary across households, changes in the law are expected to affect different individuals differently.

[6] The probability of given couple divorcing may be denoted p, where $p = p\left(m, a_I, a_{S,}\varphi\left(c_{I,S}\right)\right)$. Partial derivatives are signed as follows: $p_M \leq 0$, $p_A \geq 0$ and $p_\varphi \geq 0$. The probability is a function of $c_{I,S}$, and the partial derivative $p_C = p_\varphi \varphi_c$. The sign of this final partial derivative varies with the law: $p_C \leq 0$ under a mutual consent and $p_C \geq 0$ under unilateral divorce.

Consider two populations of married couples. 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 $c_{I,S} \approx 0$, then the marginal effect of changing divorce law may approximate to zero. However, if $c_{I,S} \gg 0$, 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 governing divorce.

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. Discussion

When legal regime changed, the divorce space increased if and only if transaction costs were positive. Figure 3 illustrates the evolution of the divorce space as transaction costs increase from zero. The arrows represent the magnitude of the costs. Under mutual-consent regimes, transaction costs serve to constrict the divorce space, while they expand it under unilateral divorce.

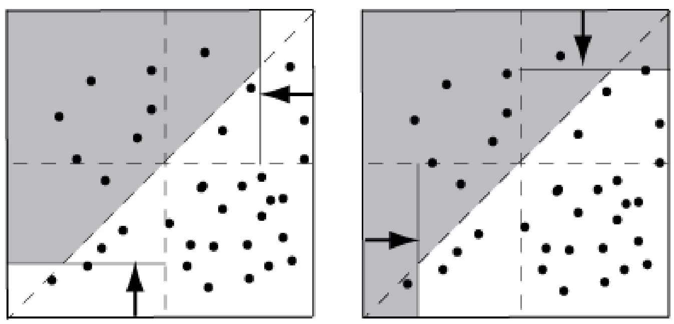

Figure 3: Evolution of the divorce space under mutual consent (left) and unilateral (right) regimes, as the level of transaction costs increase.

Figure 3 highlights how marriage market outcomes may influence the law's marginal effect. First, marriage markets can pair couples in ways that minimize intrahousehold competition ($c_{I,S}$ is low for a given match). This minimizes the marginal divorce space. Second, the marginal divorce space is located in the "northeast" and "southwest" regions. Suppose marriage markets can pair individuals such that outside opportunities are positively correlated or the gains to marriage are maximized. Either result will cluster couples in the southeast quadrant, *ceteris paribus*. In that case, even large changes in the divorce space (due to large transaction costs) may fail to produce significantly higher divorce rates.

How might this model explain Wolfers' findings that the increase in the divorce rate dissipated over time? Suppose couples in both legal regimes reduced the level of intrahousehold competition over time. This would lead to a convergence of divorce rates, which empirically might be interpreted as dissipation (our distinction between those terms is that "convergence" implies both rates move, but in opposite directions, while "dissipation" implies only unilateral rates move downwards). Alternatively, suppose couples responded to unilateral divorce by reducing the level of intrahousehold competition relative to their peers in mutual consent states. If so, then the phenomenon observed by Wolfers is due to dissipation, rather than convergence.

Our model treats transaction costs as an exogenous variable, so this issue is beyond its scope. Hanlon (2009a) models transaction costs as an endogenous variable, and he argues both hypotheses are correct to a degree. In other words, evidence suggests that intrahousehold competition decreased in both mutual consent and unilateral states, but the magnitude of the move was larger in unilateral states.

The notions of convergence and dissipation raise two issues. First, much of the empirical literature has implicitly assumed that transaction costs have been constant over time, and it does not account for differing rates of change across legal regimes.[7] This assumption may hold during the period immediately surrounding the legal transition. However, for extended panels, there may be significant confounding by omitting transaction costs from the analysis. A second issue relates to the position of a population on the transaction-cost axis in Figure 2. If transaction costs decrease, then those populations move towards the origin. If unilateral states are closer to the origin, this implies that unilateral divorce rate may be closer to the socially-efficient rate, *ceteris paribus.* Yet that outcome is not equivalent to socially efficiency, because the notion of a "socially-efficient divorce rate" is only synonymous with "social efficiency" contingent on the legal regime.

Consider a married couple that acquires a marital-specific asset. This acquisition increases the expected gain to marriage, but it may also increase transaction costs within the household.[8] Given transaction costs affect the probability of divorce, would the couple necessarily choose to acquire the asset under both legal regimes? Not necessarily, because under unilateral divorce the marginal benefit of the asset may be trumped by the probability that additional transaction costs will lead to a socially-inefficient divorce. This disparity in the accumulation of marital assets is illustrated by Stevenson (2007), who finds significant differences in the assets acquired by newly-married couples across legal regimes.

[7] The notion that transaction costs might vary over time was suggested by Rasul (2006) and Matouschek and Rasul (2008). They offer an alternate explanation for Wolfers' finding, as does Wickelgren (2009). However, those authors considered transaction costs' affect on behavior in the marriage market, rather than on intrahousehold competition. Since we address different aspects of the issue, our explanations are complementary.

[8] Thanks to Robert Pollak for highlight this point.

It should be expected that the level of intrahousehold competition is positively related to the amount of wealth over which spouses can compete, but couples with the same martial-specific assets do not necessarily engage in the same level of intrahousehold competition, *ceteris paribus*. Competition is not only a function of marital assets, but also the legal regime. The empirical challenge is to avoid confounding these factors.

V. Data and Empirical Identification

Our model has two theoretical predictions: (*i*) higher levels of transaction costs will increase (decrease) the probability of divorce in unilateral (mutual consent) regimes; and (*ii*) unilateral divorce laws will result in more divorce than mutual consent laws, *ceteris paribus*. The ideal econometric test of this theory would measure the gains to marriage, outside opportunities and transaction costs at the household level. Unfortunately, neither that data nor reliable household-level proxies exist. However, we believe a suitable proxy exists at an aggregated level. We cannot measure intrahousehold competition directly, but we can identify evidence that individuals attempt to mitigate it. We proceed on the assumption that more evidence of mitigation equates to lower levels of transaction costs within a jurisdiction, all other factors constant. Individuals may mitigate transaction costs by forming contracts that are stronger than would exist otherwise. Prenuptial and postnuptial contracts may be the most obvious measures of this behavior, but unfortunately no reliable data exists on their use. Consequently, we employ the number of marriage and family therapists (MFT) licensed within a legal jurisdiction.[9]

The notion that therapists play this role is supported by that profession's literature. Hurvitz (1974) first argued that therapists serve as "intermediaries" to assist spouses who "cannot bargain or negotiate because each believes he cannot protect his own interests,"

[9] The use of MFT may reduce transaction costs, but it cannot eliminate them entirely. At a minimum, the cost of employing a MFT is also a transaction cost, albeit one that is borne in order to avoid an even more costly alternative. So while the existence of MFT may imply transaction costs are lower than they would be otherwise, it also implies those costs are necessarily greater than zero.

and “cannot maintain their agreements although they assert the same goal.” Dion (2005) summarizes marriage education programs employed by the profession. She notes that many programs explicitly focus on improving individuals’ “commitment” to their relationships. In economic terms, it seems reasonable to assert that the value of “interpersonal commitment” is to support an informal contract between spouses.[10]

MFT licenses are distinct from other types of counseling and mental health services because their training and clinical experience focuses specifically on working with families and couples. The requirements to obtain MFT licensing are substantial. All jurisdictions require that MFT candidates have either received an accredited master's or doctoral degree in marriage and family therapy, or possess a graduate degree in an associated mental-health along with a post-graduate clinical training. Once the educational requirements are completed, candidates must then pass a board examination. Upon receipt of a graduate degree and successful completion of a board examination, candidates are conferred “associate” status. After completion of a supervised training period, associate MFT are promoted to licensed status (associate MFT are not included in this analysis, as we were unable to reliably collect that data from all jurisdictions).

[10] More formally, players in a “prisoner’s dilemma” can reach the Pareto efficient outcome by forming a coalition to play a cooperative game. This may require a mechanism to maintain the coalition by enforcing cooperative behavior. Our approach is equivalent to assuming that part of a MFT’s role is to serve as that “coalition enforcing” mechanism.

In 2009, forty-seven states and the District of Columbia required MFT to obtain licenses in order to legally practice within the jurisdiction.[11] Via correspondence with each jurisdiction's licensing authority, we collected the number of therapists licensed at the county level. Historical data were generally unavailable, so the data set is a cross-sectional count. We use the MFT data is used to create a variable that represents the total number of active licensees with a practice address located in the licensing jurisdiction.[12] The presence of licensed MFT is an imperfect measure of the actual level of transaction costs within a given legal jurisdiction. However, if MFT serve to reduce intrahousehold competition among spouses, then the frequency of therapists is expected to be negatively correlated with the average level of intrahousehold transaction costs within that same jurisdiction, *ceteris paribus*.

To capitalize on our MFT data, we need cross-sectional divorce data from roughly the same period of time. County-level divorce data is not reported by any national agency. However, many states have an office of vital statistics which publishes an annual statistical abstract. Via those reports (and in some cases, following up for additional information), we were able to match county-level MFT and divorce data for thirty-two

[11] Montana, North Dakota and West Virginia are not part of the MFT sample. In April 2009, both Montana and West Virginia passed legislation requiring MFT to obtain licenses, but licensing procedures are not expected to be fully implemented in the immediate future. North Dakota's legislative requirement was enacted in 2005. However, it appears the state has made no effort to put this law into effect. The first step in implementing licensure is to establish a licensing board. According to Dr. Thomas Carlson at North Dakota State University, as of the spring of 2009 the state had not yet taken that step.

[12] Only 94% of the therapists licensed in the United States list their practice address as being in the same state in which they are licensed. For a detailed discussion of this issue, see Hanlon (2009a). To minimize potential conflicts, we restrict our MFT variable in this study to include only therapists with a practice address in the licensing state.

states and the District of Columbia. These states are listed in Table 2, and this sample represents approximately 71% of the counties within the United States.[13]

The sample over represents mutual-consent states (17 out of the 20 jurisdictions are included while only of 16 of 31 unilateral states are in the sample). Unfortunately, adding states to this sample is prohibitively costly. In most of the remaining jurisdictions, divorce data is managed by a court system which does not aggregate or report divorce statistics, even at the courthouse-level. Court clerks at individual courthouses are generally under no obligation to provide this data. To calculate a count of divorces in Los Angeles County, for example, one would need to contact over a dozen courthouses and convince the clerk in each one to cooperate.[14] To obtain divorce information for a moderately-populated state like Indiana, one would need to contact over 400 individual courthouses across 92 counties. As a practical matter, this data is simply unattainable.

Given that our divorce data is a non-negative integer, it is unsuitable as a dependent variable in a least-squares specification. However, as discussed by Cameron and Trivedi

[13] Most of the divorce data is from 2007. However, at the time this data was collected, not every state had released data for year. Delaware, Ohio, South Carolina and Wyoming data are from 2006. Arkansas and Vermont data are from 2005. We also collected 2000 divorce data for use as an exogenous instrument. Officials from Massachusetts, Michigan and South Dakota could provide data from 2007 but not 2000. The latter series is used as an instrument in this analysis, so those three states were dropped from the sample. We also dropped the small number of counties that have been established since 2000, since a majority of the control variables for those observations required imputation.

[14] We attempted to collect this data from several heavily-populated counties in California, but we were not successful. We were able to collect divorce data from a small number of lightly-populated counties in California and Nevada. In most cases, these counties had a single courthouse, which simplified our task. However, we decided to limit our dataset to only states in which a majority of counties were represented, so those California and Nevada observations are omitted from the sample.

(2005), logarithmic transformations may be considered if the variable's mean is high. Consequently, our dependent variable is the natural logarithm of the divorce count (approximately 1% of the divorce observations were equal to zero, and these were modified to 0.5 prior to transformation). To test our predictions, we use the least-squares model represented by equation (1). *LAW* is an indicator variable, set equal to one for unilateral divorce. *MFT* is the county-level therapist count. The vector *X* contains control variables. The interaction term $LAW * MFT$ isolates the influence of the law across regimes. Our theoretical model predicts $\beta_2 > 0$, $\beta_3 > 0$ and $\beta_4 < 0$.

$$\ln\left(DIVORCE\right) = \beta_1 X + \beta_2 LAW + \beta_3 MFT + \beta_4\left(LAW * MFT\right) + \varepsilon \qquad (1)$$

The relationship between *DIVORCE* and *MFT* is potentially complicated, since therapists may use a jurisdiction's divorce rate to determine where to locate their practice. We test for this simultaneity and determine it exists. So we proceed with both reduced-form TSLS and GMM in which we fit both *MFT* and *LAW*MFT*.[15] Heteroskedasticity is expected with county-level data, and the error terms may be correlated across counties within a given state. Moulton (1990) discusses the potential pitfall of using a state-level policy variable (such as divorce law) on county-level data when disturbances are correlated within a state-level cluster. We cannot rule out heteroskedasticity or within-state correlation, so we report clustered robust standard errors.

[15] It may appear simpler to generate a fitted MFT_{IV} and then incorporate that variable into the regression. However, that approach constitutes the "forbidden regression" described by Wooldridge (2002).

A list of control variables and instruments are reported in Table 3, and descriptive statistics are provided in Table 4. Our principal control is the number of "married-family" households within a jurisdiction. We employ this metric instead of population or an aggregated count of households because only married-family households are eligible to divorce. We also use the number of married-family households with children under the age of eighteen, to account for different types of marriage. This data is from the year 2000, since the U.S. Census Bureau does not report this metric for all counties every year. However, given that households must exist before they can divorce, a lag between the household count and divorce outcomes is appropriate.

A jurisdiction's unemployment rate and the male-female ratio are incorporated to control for unexpected shocks and aspects of marriage-market competition, respectively. Finally, we include both the percentage of the jurisdiction's population who are regular church attendants and the percentage who are Catholic adherents. Conventional wisdom suggests that church attendance generally does not influence divorce, although the Catholic faith's proscription deters the frequency of divorce among its adherents.[16] This latter notion has been espoused by Jacob (1988), among others, who listed Catholic pressure as a factor that mitigated divorce in some jurisdictions. To deal with simultaneity, we employ three instruments. First, we use the lagged divorce rate. For

[16] The Barna Group, a frequently-cited research organization focused on religious issues, has reported that "when evangelicals and non-evangelical born again Christians are combined into an aggregate class of born again adults, their divorce figure is statistically identical to that of non-born again adults," and that "population segments with the lowest likelihood of having been divorced subsequent to marriage are Catholics." See http://www.barna.org/barna-update/article/15-familykids/42-new-marriage-and-divorce-statistics-released, accessed November 2009. A common interpretation of Barna's results (from this report and others) is that Catholics are less likely to divorce, although that is not the only plausible interpretation.

most of the observations, the lag between divorce counts is seven years.[17] The second instrument is a jurisdiction's population density, and the third is the per-capita income within a jurisdiction. Hanlon (2009a) explores MFTs' decision to locate their practices, and finds both of those factors are relevant.

The binary classification of states into mutual consent and unilateral regimes has been a contentious issue in the literature. Peters (1986) identified states with onerous separation requirements as mutual consent states, even if the law permitted unilateral divorce. Allen (1992) questioned the validity of that categorization. We test both Peters' classification, which we denote as our primary classification, and an alternate classification. The alternate classification codes unilateral states with onerous separation agreements as being unilateral, and we include an indicator to distinguish them from unambiguously unilateral states.

Our approach implicitly assumes the quality of marriages is constant across legal regimes. This contrasts with Rasul (2006) and Matouschek and Rasul (2008), who assume that individuals in the marriage market are actively aware of the law governing divorce, and that unilateral divorce produces better matches (in the context of this analysis, "better matches" are ones in which the gains to marriage are higher, the correlation of outside opportunities is higher, or both). To the degree their argument is correct, unilateral divorce should result in higher-quality matches in the marriage market

[17] While not directly relevant to this study, it is interesting that divorce counts decreased from 2000 to 2007, and that the decrease was far more dramatic in mutual-consent states. The number of divorces decreased 9.8% in mutual consent jurisdictions, versus 1.6% in unilateral jurisdictions.

and our point estimate of *LAW* will be downwardly biased. However, given our prediction that the coefficient is positive, this bias does not undermine our approach.

Divorce law is an identifying variable only to the degree that individuals forgo "migratory divorce," where spouses residing in one jurisdiction move to another with fewer restrictions in order to obtain a divorce. While this behavior did occur prior to the widespread adoption of unilateral divorce, it appears to have been infrequent. Carter and Glick (1976) estimate between five and ten percent of all divorces at that time could be characterized as "migratory." Jacob (1988) argues that migratory divorce was available only for very wealthy individuals, suggesting its frequency was even lower. Matouschek and Rasul (2008) address this question with data from the 1970s and conclude it was not a significant factor. Regardless of its historical significance, our study uses contemporary data, and the issue's contemporary significance is limited when individuals have marital assets (particularly children), since that may limit their ability to establish residency in a unilateral state. Also, it should be expected that migratory divorces should be concentrated in a small number of counties, such as the historical "divorce havens" such as Reno, Nevada. The state of Nevada is not part of our sample, and the influence of any contemporary havens that remain should be limited, given the high number of county-level observations.

We assert that individuals do not sort themselves across legal regimes in any way that could explain our results. This assumption is supported by Baker and Emery (1993), who surveyed marriage license applicants and law students on their knowledge of divorce statutes, among other factors. They found that (1) both group's knowledge of the law was largely inaccurate; and (2) both groups expressed "thoroughly idealistic expectations about the longevity of their own marriages." Our interpretation is that divorce laws are not part of the data-generating process individuals use to determine their residence, since that would require both accurate knowledge of the law and *ex-ante* consideration of the possibility of divorce.

We also assert the divorce law is exogenous to the presence of MFT. This is supported on two grounds. First, our MFT data is from 2009, which is over thirty-five years after the height of the "no fault revolution" in divorce laws. It is twenty-four years after the final state to alter its divorce regime did so. Given the recent growth in the MFT profession, it seems reasonable that most therapists have been practicing for less time than twenty-four years. So the divorce law preceded their existence. Second, for the MFT who have been practicing since the 1970s, it seems implausible they influenced the change in divorce laws. McDaniel (1971) summarizes the state of the MFT profession during the early 1970s, when most of the laws changed. He notes the profession had not yet solved "the functional problems of finding a technical base, asserting an exclusive jurisdiction, linking both skills and jurisdiction to standards of training, and convincing the public of the trustworthiness of its services." It seems unlikely that a profession in such an early stage of its development could have exerted significant influence on the laws governing divorce. Rutledge (1973) describes the activities of the American Association for Marriage and Family Therapy (AAMFT) during the 1960s and early 1970s. He argues its focus on was "establishing and maintain high standards in this newly developing field," and he makes no mention of divorce laws. Jacob (1988) documents the transformation in divorce laws during the 1960s and 1970s. He does not mention of marriage and family therapists, and refers to the "mental health community" only briefly and in passing. Rather, he attributes the change in divorce laws to legal practitioners who were frustrated with discrepancies between "black-letter" law and actual practice in the court.

Table 2: States included in the sample, segmented by licensing jurisdiction. The symbol † denotes states that are mutual-consent under both the primary and alternate classifications.

Unilateral	*Mutual Consent*
Alabama	Arkansas †
Arizona	Delaware †
Colorado	District of Columbia
Florida	Illinois
Hawaii	Maryland
Idaho	Mississippi †
Iowa	Missouri
Kansas	New York †
Minnesota	North Carolina
Nebraska	Ohio
New Hampshire	Pennsylvania
Oklahoma	South Carolina
Oregon	Tennessee †
Texas	Utah
Washington	Vermont
Wyoming	Virginia
	Wisconsin

Table 3: Variables and sources

CONTROL VARIABLES	
HH	Number of "married-family" households in the jurisdiction. Source: 2000 Census, U.S. Census Bureau.
HH_CHILD	Number of "married-family" households with "own" children under 18 years of age in the jurisdiction. Source: 2000 Census, U.S. Census Bureau.
UNEMPLOYMENT	Jurisdiction's unemployment rate (2006). Source: U.S. City and County Data Book (2007), U.S. Census Bureau.
MF_RATIO	Number of males per one hundred females. Source: U.S. City and County Data Book (2007), U.S. Census Bureau.
CHURCH	Percentage of the population that attends church. Source: The Association of Religion Data Archives, Religious Congregations and Membership Survey (2000). To generate a percentage, the "adjusted" number of adherents was divided by the jurisdiction's total population. See Finke and Scheitle (2005) for an explanation of the "adjusted" versus the raw count.
CATHOLIC	Percentage of the population that attends a Catholic church. Source: The Association of Religion Data Archives, Religious Congregations and Membership Survey (2000). To generate a percentage, the total number of Catholic adherents was divided by the jurisdiction's total population.
INSTRUMENTS	
DENSITY	Number of people per square land mile. Source: U.S. City and County Data Book (2007), U.S. Census Bureau.
DIVORCE (2000)	Lagged count of divorces that occurred in the jurisdiction.
INCOME	Per-capita income. Source: U.S. City and County Data Book (2007), U.S. Census Bureau.

Table 4: Descriptive statistics for county-level observations, segmented by legal classification. Standard deviations are reported in parentheses. The column *t-test* reports the p-values associated with the null hypothesis that the means across legal regimes are equal.

Legal classification	*PRIMARY*			*ALTERNATE*		
DIVORCE LAW	*MC*	*UNI*	*t-test*	*MC*	*UNI*	*t-test*
Observations	1,109	1,083		317	1,875	
DIVORCE (2007)	333.3 (747.5)	328.8 (1,064.4)	0.909	360.3 (927.1)	326.1 (916.2)	0.540
MFT	4.5 (12.9)	8.3 (34.8)	0.001	4.0 (11.3)	6.8 (27.9)	0.078
HH (thousands)	18.5 (42.7)	14.7 (41.9)	0.036	18.0 (42.8)	16.4 (42.2)	0.533
HH_CHILD	8.3 (20.2)	6.6 (20.1)	0.048	8.1 (20.3)	7.3 (20.1)	0.513
UNEMPLOYMENT	5.3 (1.7)	4.3 (1.4)	0.000	6.1 (1.8)	4.6 (1.5)	0.000
MF_RATIO	97.9 (8.2)	100.3 (10.2)	0.000	97.3 (7.9)	99.4 (9.5)	0.000
CHURCH (pct)	60.9 (19.6)	65.9 (24.9)	0.000	67.7 (21.1)	62.7 (22.6)	0.000
CATHOLIC (pct)	10.4 (13.2)	14.7 (13.9)	0.000	8.1 (13.7)	13.3 (13.6)	0.000
DENSITY	430.5 (2,749.0)	95.8 (283.2)	0.000	663.8 (4,915.7)	197.8 (664.0)	0.000
DIVORCE (2000)	369.5 (831.2)	334.1 (1030.3)	0.376	403.6 (993.7)	343.3 (924.5)	0.288
INCOME (thousands)	17,798.6 (3,948.7)	17,311.4 (3,527.6)	0.002	16,405.1 (3,783.0)	17,752.8 (3,714.4)	0.000

VI. Empirical Analysis

Following Davidson and MacKinnon (1993), we employ a Durbin-Wu-Hausman test to identify simultaneity between divorce and MFT. The results suggest endogeneity exists. For comparison's sake, we also conduct least-squares regressions. The results from OLS, TSLS and GMM regressions are reported in Table 5. As a robustness test, we run each test with a sample restricted to jurisdictions in which at least one MFT is present. The restricted sample contains just over 52% of the counties from the full sample, and those results are reported in Table 6. To interpret the regression coefficients, recall the dependent variable is a logarithmic transformation. For continuous variables, the coefficient multiplied by 100 is equal to the percentage effect of that variable on y. For indicator variables, the percentage change in y is equal to $100\left(e^{\beta}-1\right)$. See Halvorsen and Palmquist (1980) for a more thorough explanation.

To evaluate the validity, relevance and strength of our instruments, we use test statistics suggested by Kleibergen and Paap (2006). Their metrics are appropriate when errors are heteroskedastic. The null hypothesis of the LM test is that the instruments do not identify the endogenous variables. Across both samples and classifications of *LAW*, we consistently reject the null. The Wald F statistic is consistently in the high teens, which exceeds the critical values established by Stock and Yogo (2005) for the maximal IV size at the 10% level. Finally, the Hansen J statistic consistently fails to reject the null hypothesis that the instruments are uncorrelated with the error term. In total, our interpretation is that the instruments are valid, relevant and strong enough to avoid the pitfalls identified by Bound, Jaeger and Baker (1995), among others.

For both the TSLS and GMM regressions, point estimates and levels of significance are consistent across both samples and classifications of *LAW*. As predicted by the theoretical model, the coefficient on *MFT* is positive and the coefficient on *LAW**MFT* is negative. It may seen that the sum of the *MFT* and *LAW**MFT* coefficients should be negative. Yet that is not the case, since *MFT* is jointly determined across legal regimes.[18] Marriage and family therapists may engender divorce in mutual consent states and prevent it in unilateral states, but our model does not imply that the magnitude of those effects is necessarily equal, especially given the unbalanced number of observations from each legal regime.

The coefficient on *LAW* is consistently negative and significant. This is inconsistent with our prediction. However, this result does not imply that intrahousehold competition is zero (even as an approximation). Again, if unilateral divorce produces higher-quality matches as described by Rasul and Matouschek, then our estimate of *LAW* is downwardly biased. In this sense, our results are consistent with their conclusion that individuals in unilateral states exert additional effort in the marriage market to produce better matches.

[18] If the sample is restricted to unilateral jurisdictions only, the validity and relevance of the instruments break down. The coefficient on MFT is slightly positive in those regressions, but given the quality of the instrumentation, we avoid drawing any conclusion from those results.

Two other results merit discussion. First, the coefficient on the male-female ratio is negative and significant at the 1% level. This suggests that higher levels of male competition in the marriage market leads to a decrease in divorce, which implies (on the margin) the decision to divorce is driven by men, or at least men's behavior in response to marriage-market competition. The correlation between unilateral divorce and the male-female ratio is only 0.13 in our sample, but Hanlon (2009b) finds that unilateral states have significantly higher male-female ratios, *ceteris paribus*. Given this evidence, we conclude that including the male-female ratio upwardly biases our point estimate on *LAW*. This is confirmed by omitting it from the regression. A second interesting result is the coefficient on *CHURCH* is negative and significant, but the coefficient on *CATHOLIC* is not. So contrary to the conventional wisdom, once church attendance is controlled for, there is no further effect by controlling for affiliation with the Catholic Church.

In general, the sign and significance of the coefficients of interest are robust to the inclusion of other controls, such as ethnicity, voting patterns and other socioeconomic characteristics. The strength of the instruments is somewhat fragile, in that adding additional variables tends to decrease either their validity or relevance. However, on balance, we believe these results provide a valid and robust defense of our theoretical model.

Table 5: Regression results for the full sample, with ln(*DIVORCE*) as the dependent variable. In each box, point estimates are located in the top left, with clustered robust standard errors in parentheses below. In the top right corner, p-values are reported testing the null that the point estimate differs from zero.

Legal classification	Primary (*16 unilateral, 17 mutual consent*)			Alternate (*28 unilateral 5 mutual consent*)		
Regression	OLS	TSLS	GMM	OLS	TSLS	GMM
Observations	2,192	2,192	2,192	2,192	2,192	2,192
MFT	0.015 *.048* (0.008)	0.114 *.025* (0.051)	0.128 *.008* (0.048)	0.016 *.052* (0.008)	0.112 *0.030* (0.052)	0.131 *.010* (0.051)
*LAW*MFT*	-0.010 *.219* (0.008)	-0.062 *.054* (0.032)	-0.056 *.060* (0.030)	-0.010 *.227* (0.008)	-0.060 *0.064* (0.033)	-0.057 *.071* (0.032)
LAW	-0.436 *.030* (0.192)	-0.436 *.023* (0.191)	-0.497 *.029* (0.228)	-0.713 *.000* (0.171)	-0.735 *0.000* (0.177)	-0.788 *.000* (0.184)
SEPARATION	NA	NA	NA	0.351 *.093* (0.203)	0.353 *0.098* (0.214)	0.433 *.094* (0.259)
HH	0.065 *.000* (0.011)	0.077 *.002* (0.025)	0.083 *.013* (0.034)	0.065 *.000* (0.011)	0.077 *0.002* (0.025)	0.084 *.013* (0.034)
HH_CHILD	-0.104 *.000* (0.021)	-0.187 *.016* (0.078)	-0.221 *.022* (0.097)	-0.104 *.000* (0.021)	-0.187 *0.015* (0.077)	-0.225 *.020* (0.097)
UNEMPLOYMENT	-0.009 *.870* (0.054)	0.042 *.479* (0.059)	0.038 *.524* (0.060)	-0.027 *.623* (0.054)	0.022 *0.710* (0.059)	0.019 *.760* (0.061)
MF_RATIO	-0.017 *.000* (0.003)	-0.016 *.000* (0.004)	-0.016 *.000* (0.004)	-0.017 *.000* (0.003)	-0.016 *0.000* (0.003)	-0.017 *.000* (0.004)
CHURCH	-0.018 *.000* (0.003)	-0.019 *.000* (0.003)	-0.018 *.000* (0.003)	-0.019 *.000* (0.003)	-0.020 *0.000* (0.003)	-0.019 *.000* (0.003)
CATHOLIC	-0.004 *.547* (0.006)	0.001 *.864* (0.007)	0.004 *.591* (0.007)	-0.003 *.645* (0.006)	0.002 *0.761* (0.007)	0.006 *.419* (0.007)
CONSTANT	7.340 *.000* (0.585)	7.020 *.000* (0.650)	7.058 *.000* (0.669)	7.740 *.000* (0.554)	7.450 *0.000* (0.583)	7.493 *.000* (0.608)
Kleibergen-Paap LM statistic	NA	7.695 *.021*		NA	7.695 *0.021*	
Kleibergen-Paap Wald statistic	NA	19.324		NA	18.721	
Overidentification χ^2	NA	1.389 *.239*		NA	1.547 *.214*	

Table 6: Regression results for the restricted sample (observations in which MFT > 0 only). In each box, point estimates are located in the top left, with clustered robust standard errors in parentheses below. In the top right corner, p-values are reported testing the null that the point estimate differs from zero.

Legal classification	Primary (*16 unilateral, 17 mutual consent*)			Alternate (*28 unilateral 5 mutual consent*)		
Regression	OLS	TSLS	GMM	OLS	TSLS	GMM
Observations	1,144	1,144	1,144	1,144	1,144	1,144
MFT	0.014 *.021* (0.006)	0.061 *.001* (0.019)	0.061 *.002* (0.019)	0.014 *.021* (0.006)	0.059 *0.001* (0.019)	0.059 *.001* (0.018)
*LAW*MFT*	-0.009 *.159* (0.007)	-0.030 *.068* (0.017)	-0.031 *.010* (0.012)	-0.010 *.155* (0.007)	-0.029 *0.069* (0.016)	-0.028 *.021* (0.012)
LAW	-0.228 *.255* (0.197)	-0.279 *.267* (0.251)	-0.265 *.175* (0.196)	-0.500 *.005* (0.165)	-0.588 *0.029* (0.270)	-0.611 *.001* (0.181)
SEPARATION	NA	NA	NA	0.134 *.541* (0.217)	0.191 *.470* (0.264)	0.208 *.342* (0.218)
HH	0.045 *.000* (0.007)	0.053 *.001* (0.015)	0.053 *.000* (0.015)	0.045 *.000* (0.007)	0.053 *0.001* (0.015)	0.053 *.001* (0.016)
HH_CHILD	-0.070 *.000* (0.014)	-0.118 *.005* (0.042)	-0.117 *.004* (0.041)	-0.070 *.000* (0.014)	-0.117 *0.004* (0.041)	-0.118 *.004* (0.041)
UNEMPLOYMENT	-0.012 *.798* (0.047)	0.031 *.497* (0.046)	0.031 *.499* (0.046)	-0.029 *.526* (0.046)	0.011 *0.809* (0.046)	0.011 *.814* (0.046)
MF_RATIO	-0.015 *.070* (0.008)	-0.014 *.072* (0.008)	-0.014 *.071* (0.008)	-0.015 *.051* (0.008)	-0.015 *0.049* (0.008)	-0.015 *.049* (0.008)
CHURCH	-0.018 *.000* (0.003)	-0.019 *.000* (0.003)	-0.019 *.000* (0.003)	-0.019 *.000* (0.003)	-0.020 *0.000* (0.003)	-0.020 *.000* (0.003)
CATHOLIC	0.001 *.906* (0.005)	0.007 *.185* (0.005)	0.007 *.162* (0.005)	0.001 *.834* (0.005)	0.007 *0.198* (0.006)	0.008 *.127* (0.005)
CONSTANT	7.561 *.000* (1.056)	7.292 *.000* (1.039)	7.291 *.000* (1.040)	7.998 *.000* (0.940)	7.782 *0.000* (0.888)	7.798 *.000* (0.875)
Kleibergen-Paap LM statistic	NA	8.132 *.017*		NA	8.080 *0.018*	
Kleibergen-Paap Wald statistic	NA	17.301		NA	17.379	
Overidentification χ^2	NA	0.008 *.927*		NA	0.012 *.914*	

VII. Conclusion

We introduce a transaction-cost model of intramarital bargaining. It replicates the predictions from Peters (1986), and it is consistent with empirical analyses inspired by her work. Our theoretical contribution is to show (*i*) the influence of transaction costs on divorce varies across legal regimes; and (*ii*) the magnitude of the law's marginal effect on divorce is a positive function of the transaction costs. We test our theory by examining the relationship between marriage counselors and divorce. We find the influence of marriage counselors on divorce varies across legal regimes. Assuming that part of a marriage counselor's role is to reduce transaction costs within the household, this evidence is consistent with our theory.

In contrast to models of Pareto efficient households, our theoretical model offers a clear interpretation of why this empirical phenomenon occurs. Our results are consistent with the notion from Rasul (2006) and Matouschek and Rasul (2008) that marriage markets produce higher-quality matches in unilateral states. However, our results suggest that marriage markets cannot produce Pareto efficient marriages, even as an approximation.

Our empirical sample is limited, and we do not believe our results are robust enough to pin down the magnitude of factors involved in a household's decision to remain married or divorce. However, the fundamental question explored in the literature on divorce is whether the Coase Theorem applies to martial relations. Unless the observed variation across legal regimes represents perfect compensation for the alteration of property rights, the law affects individuals' welfare. If welfare is conditional on the law, then the Coase Theorem does not apply. Our theoretical model offers a clear interpretation of existing empirical evidence, and it suggests the transaction costs of intrahousehold exchange are significantly greater than zero.

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