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TESTING NASH-BARGAINING HOUSEHOLD MODELS WITH TIME-SERIES DATA

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ABSTRACT

This paper uses a “natural experiment” in Canadian divorce law reform to discriminate empirically between unitary and Nash-bargained models of the household. Using time-series data from three Canadian provinces, it demonstrates that following landmark divorce law reforms in the 1970s—reforms that led to improvements in women's expected settlement upon divorce in Ontario and British Columbia, suicide rates for older, married women in these provinces registered a sharp decline. Similar declines were not registered for younger, unmarried women or men in Ontario and British Columbia, nor for older, married women in Quebec, where the legal basis for divorce did not change. These results are consistent with Nash-bargained models of the household but not with the unitary model.

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1. INTRODUCTION

> If the joint family utility framework is to be replaced by a less parsimonious model of intra-family allocation, the increase in complexity should be explicitly demonstrated to have empirically distinguishable predictions (Rosenzweig and Schultz 1984).

Economists analyzing household decisionmaking allocation have traditionally assumed that the household acts as a single unit. It is assumed that there exists a single decisionmaker whose preferences form the basis of a household welfare function and that all household resources are effectively pooled.[1] This is termed the “unitary model” (Alderman et al. 1995), the “common preference model” (Lundberg and Pollak 1996), or the “joint family utility model” (Rosenzweig and Schultz 1984).

More recently, the assumptions behind this approach have been questioned. An alternative, the “collective model,” takes as its starting point the possibility that different household members have different preference orderings and that the resolution of these differences is a nontrivial problem. One approach taken by Pierre-Andre Chiappori and his collaborators starts with individual utility functions and assumes only that the reconciliation of differing preferences is Pareto efficient (Browning et al. 1994). A second

[1] The theoretical issues associated with these assumptions, and with alternative models that also generate a single household welfare function, are discussed in Bergstrom (1997) and Alderman et al. (1995).

approach is taken by Marjorie McElroy (1990, 1997), building on earlier collaborative work with Mary-Jean Horney (McElroy and Horney 1981) as well as a related paper by Marilyn Manser and Murray Brown (1980). Here, individuals use a cooperative Nash-bargaining solution as the mechanism by which differences in preference rankings are resolved. A somewhat different approach is taken by David Ulph (1988), Lundberg and Pollak (1994), and Michael Carter and Elizabeth Katz (1997), who use noncooperative game theory. Instead of resolving differences through preferences, individual household members maximize their own utility functions, taking as given the maximizing actions of other household members. Finally, Lundberg and Pollak (1993) have suggested a "separate spheres model" in which household members base their joint decisions on a cooperative bargaining solution, with a noncooperative outcome being the fallback position of each party.

Despite the plethora of theoretical models that have been developed in the last 15 years, reviews by Jere Behrman (1996); Hoddinott, Alderman, and Haddad (1997); Lundberg and Pollak (1996); John Strauss and Duncan Thomas (1995); and John Strauss and Kathleen Beegle (1995) emphasize that attempts to differentiate unitary from collective models suffer from several failings. One is the problem of "observational equivalence." Certain collective models yield predictions identical to those derived from a unitary model. A second is that regressors that are claimed to reflect bargaining power, such as an individual's income, can plausibly be regarded as endogenous, rendering their estimated parameters biased. Finally, even if a "bargaining variable" is truly exogenous, it may be correlated with some unobserved characteristic of the individual, again producing biased parameter estimates. Alderman et al. (1995) and Lundberg and Pollak (1996) argue that on the whole, the evidence supports a shifting of the burden of proof onto those who would maintain the unitary model as a basis of understanding intrahousehold allocations. Others, such as Behrman and Strauss, argue that these econometric deficiencies mean that studies claiming to differentiate unitary from collective models are fatally flawed.

This paper provides a somewhat stark empirical test that discriminates between unitary and one variant of collective models, the Nash bargaining model. At the core of this model is the argument that a change in women's fallback position outside of the household improves their position within the household. The empirical challenge is to find such a change that is not subject to the critiques noted above.

We argue that changes in provincial divorce law in Canada provide precisely such an example. Matrimonial property law in Canada comes under provincial jurisdiction. Under common law "the husband and wife became one person" by marriage, and common law added the rider that "the husband became that person."[2] The husband acquired control over all property owned by his wife at the time of marriage or acquired by her during marriage, except for minor, personal items such as jewelry and clothing. In the late 19th century, the enactment of the "Married Woman's Property Act" in the United Kingdom, and its subsequent replication in all Canadian provinces except Quebec, gave

[2] This discussion draws heavily on McLeod and Mamo (1993, amended 1995).

women the right to hold property in their own name, including property acquired both before and during marriage. However, the acquisition and ownership of property was tied to direct financial contribution. If a husband was the household breadwinner while his spouse remained at home, the wife had no legal claim to assets acquired in her husband's name.

A case before the Supreme Court of Canada in 1973, *Murdoch v. Murdoch,* was the catalyst for change in this aspect of family law. The case involved a contested divorce between a Mr. and Mrs. Murdoch. For the first four years of their marriage, the Murdochs worked on ranches as a hired couple, with their pay being given to Mr. Murdoch. These funds were used, in part, to purchase a ranch and homestead. Over the next 20 years, Mrs. Murdoch made a substantial contribution to the operation and management of the ranch. When the marriage broke down, she sought a judicial separation and claimed she was entitled to a one-half share, not only in the homestead, but also in the ranch. However, in the absence of a direct, financial contribution or an extraordinary financial contribution, the court held that Mrs. Murdoch's actions were "just about what the ordinary rancher's wife does." As there was no explicit agreement linking Mrs. Murdoch's labor efforts to an entitlement to a share of the ranch, she was deemed to have no interest in the ranch. She thus lost her case and faced financial destitution.

Almost immediately, remedial legislative action was put into place. The Government of Ontario passed the Family Law Reform Act, 1975. It provided that "no spouse should be dis-entitled to an interest in property simply because his or her efforts were no more than might be expected from a reasonable spouse of the kind" (McLeod and Mamo 1995, O–4). Nonmonetary contributions to the household, such as child care, gave the contributory spouse an interest in assets acquired during marriage. It is important to note that the law applied to all marriages in existence at that time. The act was subsequently amended by the Family Law Act, 1978, which went on to "provide that upon marriage breakdown each spouse was entitled to an equal interest in the 'family assets' regardless of who owned the assets."

Other provinces followed suit, with the exception of Quebec. As a French colony until 1763, Quebec was subject to civil law as embodied in the Custom of Paris. This body of laws was retained when Quebec became a British colony and was eventually codified in 1866 and protected in the bill that provided Canada with its independence from Britain in 1867. "The Custom of Paris treated matrimonial property as community property but permitted the parties to enter into a marriage contract and, thereby, choose to be in separation of property" (McLeod and Mamo 1995, Q-1). Although this law has been subject to some change over time, the basic nature of this arrangement—the fact that both parties have entitlement to property acquired over the course of the marriage—has remained unchanged.

We test the Nash cooperative bargaining model by examining the impact of the legislative change on the suicide rates of adult women. Although suicide rates might at first appear a rather extreme indicator of well-being, we believe it is well suited for the purposes of our study. Not only is the decision to take one's own life the ultimate indicator of an absence of well-being, we can draw on the formal economic analysis of suicide presented by Daniel Hamermesh and Hugo Soss (1974), and the recent extensions proposed by Avinash Dixit and Robert Pindyck (1994), to derive testable implications for the theory of household models.[3]

The paper begins with a short theoretical discussion that provides a direct link between suicide and a change in the bargaining position of women within the household. After briefly outlining our data sources and our modeling strategy, we consider the impact of this legislative change on suicide rates of adult women and men in Ontario. We then extend our analysis to Quebec and British Columbia.

2. THEORY

In this section, we synthesize McElroy and Horney's (1981) and McElroy's (1990, 1997) cooperative Nash-bargained model of household behavior with Hamermesh and Soss's (1974) economic theory of suicide. Doing so satisfies the criterion laid out by Rosenzweig and Schultz (1984) at the beginning of this paper: that our alternative to the unitary model has an empirically distinguishable prediction.

We assume that a household consists of two individuals, m and f. If they are unmarried, their utility functions are given by $U_o^m(x^m, l^m)$ and $U_o^f(x^f, l^f)$, where U is the utility, the x's are consumption goods consumed by m and f, respectively, and the l's refer

[3] Moreover, an explicit link between suicide and unhappy marriages is not an unusual concept, forming as it does the basis of novels such as Tolstoy's *Anna Karenina* and Fassbinder's film, *Effi Briest*.

to the consumption of leisure time. These utility functions are maximized subject to individual full-income constraints. Denoting p's as prices, w's as wages, T as time endowment, and I's as nonlabor incomes, these constraints can be written as

$$p^i \cdot x^i + w^i \cdot l^i = I^i + w^i \cdot T \text{ for } i = m,f.$$

This constrained maximization generates indirect utility functions, $V_o^i(p^i, w^i, I^i; \alpha^i)$. Note that these include the parameters α^m and α^f. McElroy (1990, 563) terms these "extra-household environmental parameters" or EEPs. These are variables that shift the maximum value of utility attainable outside of marriage. "EEPs include wealth or permanent income and productivity outside of marriage . . . They would also include parameterizations of variations in the rules for property settlements and of the rules governing marriage and divorce" (McElroy 1997, 58).

If m and f marry, V_o^m and V_o^f serve as threat points. m and f are assumed to have utility functions defined over their own and their spouse's consumption of goods and leisure, $x = (x^m, x^f, l^m, l^f)$. This gives $U^m(x)$ and $U^f(x)$. The Nash-bargained solution requires that m and f choose x so as to maximize the product of the gains from marriage, that is,

$$N \equiv [U^m(x) - V_o^m(p^m, w^m, I^m; \alpha^m)] \cdot [U^f(x) - V_o^f(p^f, w^f, I^f; \alpha^f)].$$

Maximizing the gains from marriage subject to the sum of the individual full-income constraints yields demand functions for goods and leisure of the following form:

$$x^i = x^i\,(p^m, p^f, w^m, w^f, I^m, I^f; \alpha^m, \alpha^f) \qquad i = m, f$$

$$l^i = l^i\,(p^m, p^f, w^m, w^f, I^m, I^f; \alpha^m, \alpha^f) \qquad i = m, f\,.$$

Note that $(\partial x^i / \partial \alpha^i) \geq 0$ and $(\partial l^i / \partial \alpha^i) \geq 0$ for $i = m, f$. An improvement in the extra-household environmental parameter for person i generates an increase in his or her level of consumption and thus in his or her well-being. It is important to note, as McElroy demonstrates, that these EEPs do not appear in the standard, unitary model of the household.

The next step borrows heavily from Hamermesh and Soss's (1974) economic theory of suicide. We begin by denoting a vector, Q^i, whose elements consist of x^i and l^i. For an individual aged n, the present value of their lifetime utility at age a can be written as

$$Z_i(a,Q) \;=\; \int_a^{\omega} e^{-r(n-a)}\, U_n^i\, P(n)d(n)\,,$$

where r is the private discount rate, ω is the highest attainable age, Ui is i's utility function as defined above, and $P(n)$ is the probability of surviving to age n given that that person has survived to age a. Z is an increasing function of Q, that is, $\partial Z/\partial Q > 0$. Following Hamermesh and Soss, we define $b^i \sim N(0, \sigma^2)$ as i's distaste for suicide. A person commits suicide if

$$Z_i\,(a, \boldsymbol{Q}) + b^i = 0.$$

The fraction of individuals within a given cohort who commit suicide at age a is

$$S(a) = f\,[\,-\,Z(a, Q)],$$

where $f(\cdot)$ is the density function for b^i. As Hamermesh and Soss note (1974, 85), the suicide rate is the proportion of individuals within a cohort for whom $Z(a, Q)$ equals b at age a. Totally differentiating this yields

$$dS = -\,f'\;(\,(\partial Z\,/\,\partial a)\cdot da + (\partial Z/\partial Q)\cdot dQ)\;.$$

Since $\partial Z/\partial Q > 0$ and $\partial Q\,/\,\partial \alpha^i > 0$, an improvement in the EEP leads to an increase in Z.

Three empirical implications follow from this. First, the change in divorce settlement law is clearly an example of a change in an EEP. We expect that it should reduce suicides only among married women: there should be no effect on suicides of unmarried women. Second, the legislative change described in the introduction should have a larger effect on the suicide rates of married, older women. For life-cycle reasons, these women are more likely to be in households that have accumulated significant levels of wealth. By contrast, we would expect the legislative change to have had a lesser effect on newly married women whose households had not yet begun to accumulate. Also, if older women find it more difficult to re-enter the labor force, access to household assets on divorce would be an especially important source of income. Third, the impact of this legislative change is asymmetric. In this model, suicide occurs when the present value of lifetime utility falls to zero. Ensuring that women have the ability to avoid destitution after divorce can be seen as lifting their post-divorce lifetime utilities above zero. Although the legislative change reduces men's utilities, these do not fall to zero. Men's utilities would fall to zero only if, following divorce, women were given all assets and men were left destitute.

3. DATA AND METHODS

The basic hypothesis we wish to test is that the change in Ontarian law should have led to a structural break in the trend suicide rate among married women in Ontario after 1975. Before turning to our results, we discuss several aspects of the data and our methods.

Our data on suicides are annual rates for the period 1960–1994. Although the legal reforms were enacted in all Anglophone provinces between 1975 and 1979, we focus our attention on the impact of this legal change in Ontario. There are two reasons for this. First, the accuracy of suicide data is often contested. Some authors claim that up to a third of Canadian coroners are reluctant to certify a death as suicide (Syer-Solursh and Wyndowe 1981). In Ontario, all coroners must be medically trained and all must attend provincial programs on death certification. As a result, there is greater consistency in the reporting of suicides in Ontario than in other provinces (Health and Welfare Canada 1989). Further, because suicide is a comparatively rare event, measured in cases per

100,000 individuals, even small variations in reporting from year to year will have a marked effect on the reported rate in much less populated provinces. For example, the reported rates for suicides among women aged 40–44 in the province of Saskatchewan for the six years between 1970 and 1975 were 12.0, 0.0, 8.4, 0.0, 4.4, and 18.0. These data were for a period in which there were 3, 0, 2, 0, 1, and 6 suicides, respectively, for this age group.

Second, data on suicides are reported by age and sex, not by marital status.[4] This is somewhat problematic for us as the strongest test of our model involves comparing the impact of this legislative change on married and unmarried women. Our solution is to divide our data into four age cohorts: 15–19, 20–24, 25–44, and 45–64. The proportion of women who are married in the younger two cohorts around the time of this legislative change was about 8 percent and 50 percent, respectively. By contrast, approximately 90 percent of women in the two older cohorts were married. There is a second reason for dividing the date by age cohort, one that follows from three interrelated factors: a relatively high incidence of suicide throughout Canada during the Great Depression, a relatively low rate in the postwar period, and the established fact that individuals, and especially children, who are bereaved by a suicide are themselves much more likely to commit suicide themselves (Cain 1966; Giffin and Felsenthal 1983). Suppose we ignored these factors and presented results based on data for the age group 20–65. This combined

[4] Our data are drawn from the *Vital Statistics* series 804-202, -204, -205, and -206 published each year by Statistics Canada and, prior to 1971, the Dominion Bureau of Statistics.

age cohort would contain individuals who enter into our sample with significantly less exposure to suicide at precisely the point in time at which we hypothesize that there should exist a structural break in the suicide trend for a completely different reason. To elaborate, a 25-year-old born in Ontario in 1950 who has relatively little exposure to suicide in her parents' generation enters the sample in 1975. By separating the sample into four cohorts, we can test whether this factor (an age cohort that is not very familiar with suicide) is affecting our results.

We deliberately exclude children in the work reported below. For young people, other factors, such as difficulties with schooling, broken homes and/or parental divorce, child abuse, and sexual assault, are held to be much more important with regard to suicide (Garfinkel and Golombek 1977). We exclude consideration of individuals over the age of 65 because here, too, the principal cause of suicide lies in factors completely unrelated to distribution of resources within marriage. In particular, deterioration of both physical and mental health are major determinants of suicide in elderly Canadians (Lepine 1982).

The next feature we note is that the period spanned by these data is one that also witnessed major changes in labor market conditions and, in particular, a significant increase in the long-run or natural rate of unemployment (Bean, Layard, and Nickell 1986). Female unemployment in Ontario rose from 4.7 percent in the period from 1966 to 1974 to 8.0 percent in the period 1975–1994, and over the same period, the rate for men rose from 3.6 to 7.1 percent. This shift away from near full-employment occurred around the same time as the hypothesized break in the trend rate, raising the possibility that results for a univariate time-series analysis of suicide rates may be confounded by changes in labor market conditions. We therefore examine the evolution of suicide rates in a multivariate setting where we control for changes in unemployment.[5]

Finally, we include a trend variable in the model to capture other changes in all other variables that might affect suicide rates during this period. In preliminary work, we experimented with the inclusion of women's labor market participation rates, levels of real per capita income, and the ratio of male-female wages. However, because all three trend upwards over this period, their effects are captured via the inclusion of a deterministic trend. Our hypothesis, therefore, is that there is a significant break in the trend suicide rate following the change in marriage law and this is written in the following form:

$$S_{ijt} = \alpha^{0}_{ij} + \beta^{0}_{ij}t + \alpha^{1}_{ij}DU_{j} + \beta^{1}_{ij}DT_{tj} + \gamma_{j}\Delta UN_{jt} + \epsilon_{ijt}, \qquad \textbf{(1)}$$

where S is the suicide rate defined as suicides per 100,000 people, ΔUN is the change in the unemployment rate, i denotes the age-sex cohort, j the province, and t is time. The dummy variables are defined relative to the hypothesized "break-point," Tb, thus

$$DU_{j} = 1 \qquad \text{if } t > Tb\text{, 0 otherwise,}$$

[5] We use the aggregate provincial unemployment rates instead of gender-specific rates. This allows us to utilize a longer time series on unemployment. However, for the subsample where we do have separate data on male and female unemployment rates, there is no evidence of any substantial deviations between the two series other than in their mean levels.

$$DT_j = t - Tb \qquad \text{if } t > Tb\text{, 0 otherwise.}$$[6]

Finally, a word of caution is in order. In addition to the issues noted above, we are also faced with the problem of drawing inferences from a relatively short time series. With only 35 observations, the precision of parameter estimates will be relatively low and tests of the significance and timing of structural breaks will exhibit low power (Favero and Hendry 1990; Adam 1993).

4. RESULTS

FEMALE SUICIDES IN ONTARIO BY AGE COHORT

We begin by testing the hypothesis that the revision to the Ontario law of settlement in light of *Murdoch v. Murdoch* leads to a change in the trend suicide rate among women older than 25 (our proxy for married women) and that the magnitude of this change is greatest for the oldest cohort. We start with a description of the data and then proceed to test, within the confines of our simple model, whether there is evidence of a structural break.

[6] Note that our data are consistent with at least two processes. The data could be the outcome of a random walk or difference-stationary process or, as discussed by Pierre Perron (1989), they could be generated by a trend-stationary process where the deterministic trend has experienced a break. Following the method proposed by Perron (1989) we conclude that the cohort-specific suicide rates are best characterized as stationary processes around "broken" deterministic trends. To save space, these statistics are not presented here, but are available on request.

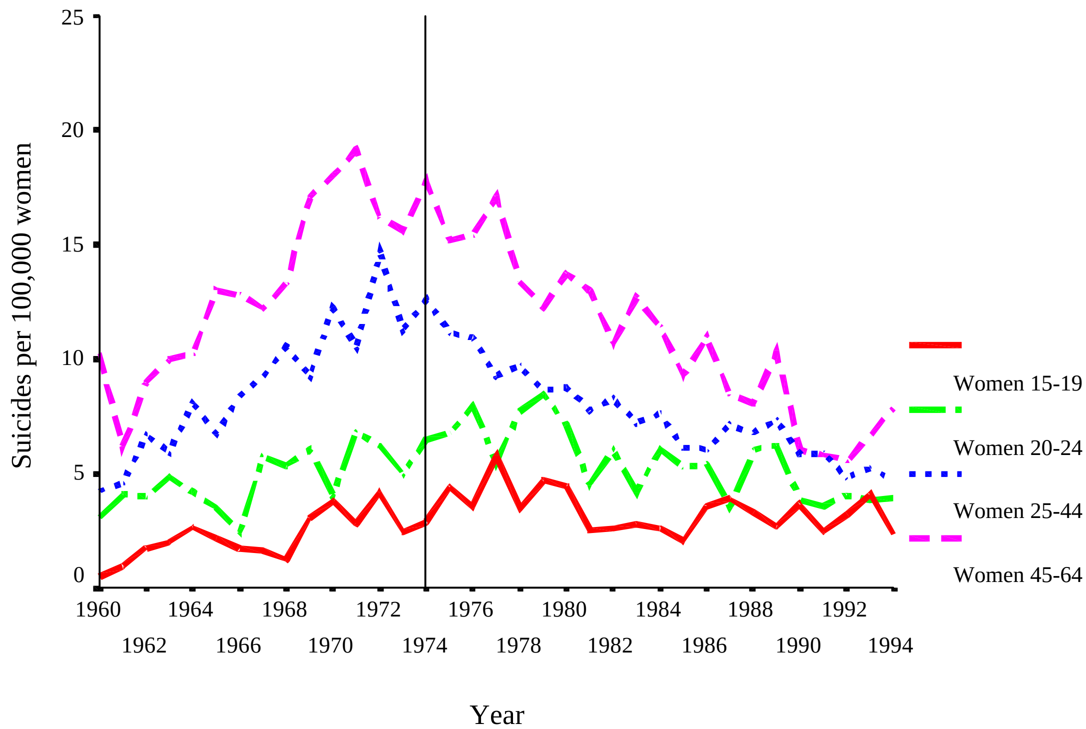
Figure 1 illustrates the evolution of the age-specific suicide rates in Ontario. Two features are immediately apparent. The first is that the marked decline in female suicide rates in Ontario for the age cohorts 25–44 and 45–64 from the mid-1970s onwards. By contrast, there is no obvious change in the rates for the younger cohorts. The second feature, which is borne out by our subsequent analysis, is that the two older age cohorts follow broadly similar paths. To the extent that there are cohort effects, these are fully captured in the mean value and the corresponding shift in the mean of the series. In other words, the timing and sign of the structural break—where it exists—is common across these age cohorts.

Table 1 reports the results of estimating equation (1) for the four age cohorts.[7] The hypothesis that there is a significant structural break is tested formally by examining the validity of the restriction in equation (1):

$$H_0\colon\ \alpha^1_{ij} = \beta^1_{ij} = 0 \,. \tag{2}$$

[7] Note that, in these regressions, the dependent variable is the provincial suicide rate. An alternative basis for the test would be to normalize Ontario and British Columbia rates on those of Quebec. Doing so does not alter the results presented here.

Figure 1   Female suicide rates: Ontario



The *prima facie* evidence of a change in the trend rate of female suicide in Ontario following legal reform is borne out in the results presented in Table 1. First, the 1975 dummy is negative and significant for the two older age cohorts, but is not negative for the younger cohorts. Second, the F-tests strongly reject the null hypothesis of no structural break for women aged 25–44 and 45–64. Third, note that the magnitude of the change in trend is highest for the 45–64 age cohort. After 1975, the trend rate of suicide per 100,000 women in this age cohort falls by –0.539 per year (0.847 – 1.386), twice as

Table 1 Determinants of female suicide rates in Ontario, by cohort

Cohort	15–19	20–24	25–44	45–64
Constant	1.131	3.528	4.145	6.398
	(2.612)	(4.657)*	(7.234)**	(6.755)**
Trend	0.0145	0.199	0.612	0.847
	(1.615)	(3.173)	(10.062)**	(8.432)**
Change in unemployment rate	0.244	–0.183	0.127	–0.168
	(0.833)	(0.659)	(0.891)	(0.717)
1975 dummy (DU)	0.851	1.56	–2.792	–2.718
	(1.367)	(2.032)	(4.469)*	(2.629)
Trend break (DT)	–0.021	–0.383	–0.907	–1.386
	(1.490)	(5.022)*	(12.853)**	(11.864)**
R-square	0.38	0.55	0.88	0.87
n	35	35	35	35
Mean suicide rates				
1960–1994	2.91	5.19	8.12	11.85
1960–1974	2.24	4.79	8.99	13.39
1975–1994	3.40	5.50	7.46	10.70
F tests on structural break coefficients (Hypothesized structural break is 1975)				
Intercept (a1 = 0)	1.474	0.69	3.36	0.25
Trend (b1 = 0)	3.83	18.92**	165.20**	140.74**
Joint (a1 = b1 = 0)	2.055	9.81**	83.72**	70.38**
Cox encompassing tests for nonnested models				
Null hypothesis: Model 1 encompasses Model 2	–2.582**	–1.092	0.391	0.591
Null hypothesis: Model 2 encompasses Model 1	–3.487**	–0.764	–17.721**	–12.048**

Notes: t-statistics are in parentheses. ** denotes significance at 1 percent; * denotes significance at 5 percent. Critical values for regressions are 1 percent, 5.99; 5 percent, 3.40. In Model 1, the structural break occurs in 1975. In Model 2, the structural break occurs in 1980. Cox test has standard normal distribution under the null hypothesis that Model 1 encompasses Model 2.

fast as the fall in the rate of women 25–44 and three times as fast as the fall in the rate for women 20–24.

There are two obvious criticisms of the interpretation of these results. First, an observer examining Figure 1 might argue that the break occurs before 1975. Second, although the 1975 intercept dummy variable is not significantly negative for the youngest cohort, the trend break is negative and significant. We address these concerns in turn. The short span of the sample data means that standard approaches to determining breaks in series when the timing of the break-point is unknown, such as those discussed by Donald Andrews (1993), will tend to have low discriminating power. However, it is possible to select informally the most likely break-point by comparing the explanatory power of the model estimated over all possible break-points within a range, here taken to be the period 1970 to 1984. This process confirms that the "most likely" break-point for suicide rates for women aged 25–44 and 45–64 in Ontario was around 1974 or 1975, but certainly no earlier.[8]

The second puzzle is that although the time-series analysis rejects the null hypothesis that there is no break in the deterministic trend for the 20–24 cohort, visual inspection of Figure 1 suggests that a break occurs around 1980. Whether 1975 or 1980 is the most likely break-point can be tested, using Cox's (1961) nonnested encompassing test. The Cox test is performed under two specifications of the null hypothesis: (1) that the model specified with the structural break occurring in 1975 (Model 1) encompasses

[8] These results are available on request.

the same model specified with the structural break occurring in 1980 (Model 2), and (2) vice versa.[9] The results of the Cox tests are presented at the bottom of Table 1. They provide decisive evidence that the "early-breaking" model variance-dominates the "late-breaking" one for the 25–44 and 45–64 age cohorts, an outcome that is consistent with our argument. For the 20–24 cohort, there is no decisive evidence that either 1975 or 1980 represents a decisive break in the series. For the 15–19 cohort, neither break-point encompasses the other. This reflects the fact that there is no trend at all. Including a break simply creates noise and consumes degrees of freedom.

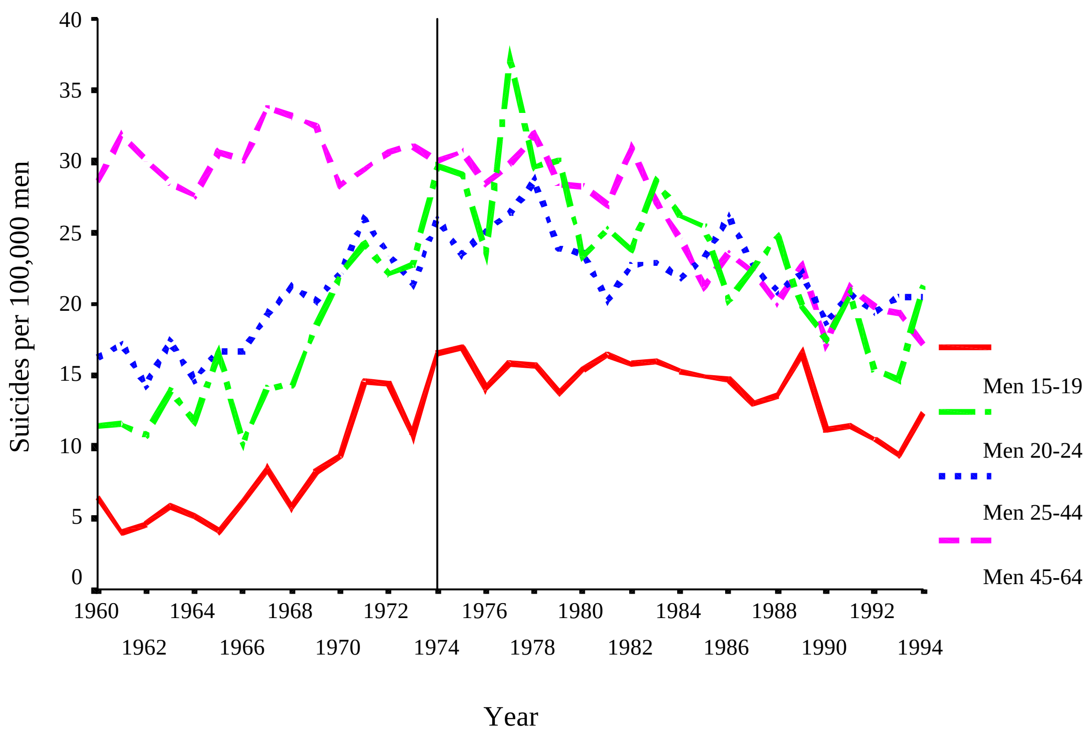
MALE SUICIDE RATES IN ONTARIO BY AGE COHORT

We now examine whether there exists a structural break in the trend suicide rate for men in Ontario. The null hypothesis is that there should be no break associated with a change in the divorce law. Rejection of this null hypothesis would cast doubt on the claim made thus far as it would suggest that there was some other factor, common to both men and women, that led to a fall in suicide rates in the mid-1970s.

Figure 2 provides age-specific male suicide rates for men living in Ontario. Visual inspection of the data suggests that, if anything, male suicide rates rise in the immediate post-1975 period. However, the data also show a fall after 1978. With these data we

[9] The Cox test is one of a class of variance-encompassing tests. Two outcomes are possible. The first, decisive outcome is that there is one-way encompassing, so that Model 1 encompasses Model 2 and Model 2 does not encompass Model 1. The second is where neither model encompasses the other, indicating that neither model is a dominant specification of the data.

Figure 2 Male suicide rates: Ontario



estimate equation (1) for our three male age cohorts and test the null hypothesis of no structural break. As the F tests reported in the middle of Table 2 indicate, the null hypothesis is rejected for all three cohorts. In other words, the data seem to suggest that the structural break observed in the female suicide rates in the aftermath of the Family Law Reform Act, 1975 is also observed among men for whom the legal reform should, according to our model, have had no impact.

However, closer examination of the characteristics of the time series on male suicide rates suggests a possible explanation for the apparent break in trend. First, as

Table 2 Determinants of male suicide rates in Ontario, by cohort

Cohort	15–19	20–24	25–44	45–64
Constant	1.455	5.935	13.811	31.283
	(1.066)	(2.578)	(14.512)**	(31.395)**
Trend	0.772	1.118	0.678	–0.073
	(6.821)**	(6.200)**	(8.718)**	(0.988)
Change in unemployment rate	0.175	–0.139	0.055	0.691
	(0.078)	(0.287)	(0.193)	(2.348)
1980 dummy (DU)	2.861	5.273	3.862	–0.661
	(2.068)	(2.306)	(3.004)	(2.629)
Trend break (DT)	-1.061	–1.939	–0.899	–0.734
	(7.716)**	(8.363)**	(6.705)**	(5.226)**
R-square	0.82	0.79	0.74	0.87
n	35	35	35	35
Mean suicide rates				
1960–1994	11.66	20.95	21.34	27.09
1960–1974	8.32	16.94	19.53	30.42
1975–1994	14.17	23.95	22.70	24.59
F tests on structural break coefficients (Hypothesized structural break is 1975)				
Intercept (a1 = 0)	4.09*	0.138	0.44	0.83
Trend (b1 = 0)	0.023	19.53**	59.18**	21.99**
Joint (a1 = b1 = 0)	7.35*	10.68**	32.26**	13.14**
Cox encompassing tests for nonnested models				
Null hypothesis: Model 1 encompasses Model 2	–0.615	–1.348	–2.175*	–2.014*
Null hypothesis: Model 2 encompasses Model 1	–1.589	–0.532	–1.361	–0.393

Notes: t-statistics are in parentheses. ** denotes significance at 1 percent; * denotes significance at 5 percent. Critical values for regressions are: 1 percent, 5.99; 5 percent, 3.40. In Model 1 the structural break occurs in 1975; in Model 2 the structural break occurs in 1980. Cox test has standard normal distribution under the null hypothesis that Model 1 encompasses Model 2.

Figure 2 shows, the suicide data for the older male cohort in Ontario indicate unusually high recorded suicide rates for 1968 and 1971. (The recorded rates of 36 and 40 per 100,000 men were 2.1 and 2.8 standard deviations, respectively, from the mean of the 1960–1974 period.) Given the short sample period of data, not only do these outliers strongly influence the time-series analysis on whether and when there was a break in the trend, they also suggest significantly differing behavior between the two Ontarian male cohorts. If we were to treat these rates as genuine outliers, the results for the two Ontarian male cohorts would converge. Second, although the time-series analysis still rejects the null hypothesis that there is no break in the deterministic trend, it seems that this break does not occur until much later than the *Murdoch v. Murdoch* ruling. As discussed above, the short span of the sample data means that standard approaches to determining breaks in series when the timing of the break-point is unknown lack the power to reject the null. Therefore, as with our analysis of the data on female suicides, we examine the sequence of equation standard errors generated by estimating the model for the period 1970 to 1984. This process suggests that 1980 emerges as the most likely break-point for the male suicide rate in Ontario, in contrast to the 1974/75 break point observed for women. Notwithstanding the charge of data mining, we use this procedure to estimate the best-fitting versions of the model (equation [6]), presented in Table 2. These preferred specifications are then used to construct a second Cox (1961) nonnested encompassing test. It is performed under two specifications of the null hypothesis: (1) that the model specified with the structural break occurring in 1975 (Model 1) encompasses the same model specified with the structural break occurring in 1980 (Model 2), and (2) vice versa. Results are presented at the bottom of Table 2. They indicate that the "late-breaking" model variance-dominates the "early-breaking" one for male suicide rates in Ontario. Recall that the converse is clearly the case for the female rate in Ontario. It could be argued that the presence of this later structural break for males could be related to changes in divorce laws, with a rather long adjustment lag. If correct, this would seriously weaken the argument being advanced in this paper. It is, however, difficult to sustain the argument that the same phenomenon—changing the rules of divorce in favor of women—should induce an immediate positive response for female suicide rates but should take almost 6 years to induce a change in male suicide rates. It seems reasonable, therefore, to suggest that the break in male suicide rates at the end of the 1970s is due to factors other than changes in divorce law following *Murdoch v. Murdoch*.

FEMALE SUICIDES IN QUEBEC

The evidence presented thus far is broadly supportive of our hypothesis that changes in laws of settlement improved the well-being of married women in Ontario, as measured here by the fall in suicide rates of women aged 25–64. This legislative change does not seem to have had a strong effect on two groups whom we had hypothesized would be unaffected by this change: younger women (who were less likely to be married) and men. However, it could be argued that perhaps there was just some other factor that caused the suicide rates of women aged 25–64 to fall. We can test this hypothesis by exploiting a “natural experiment” in our data. Recall from our introduction that the legal change did not occur in Quebec. Therefore, unlike Ontario, we expect that there should be no change in the trend rate of suicide for women 25–44 and 45–64. Figure 3, which shows the suicide rates for Quebec women, disaggregated by age, is entirely consistent with this claim. There is no obvious change in the trend suicide rate for women in these age cohorts around 1975. This visual evidence is confirmed by econometric testing, reported in Table 3, which indicates that the null hypothesis (equation [2]) cannot be rejected for the older age cohorts in Quebec and is only weakly rejected for the youngest cohort.[10]

SUICIDE IN BRITISH COLUMBIA

We began our analysis with data from Ontario because, as suicide is a very rare event, it was necessary to draw on a large population and Ontario is Canada’s most populous province. We also had prima facie reasons for believing that these data were less prone to reporting errors. In this section, we extend our analysis to British Columbia, Canada’s third most populous province. We do so with some caution. British Columbia has a much higher proportion of Aboriginal peoples within its borders. The recording of suicides among these communities has proven to be particularly problematic (Health and Welfare Canada 1989), thus exacerbating problems of measurement error. Further,

[10] In order to save space, the full results for Quebec are not reported here, but are available on request.

Figure 3 Female suicide rates: Quebec

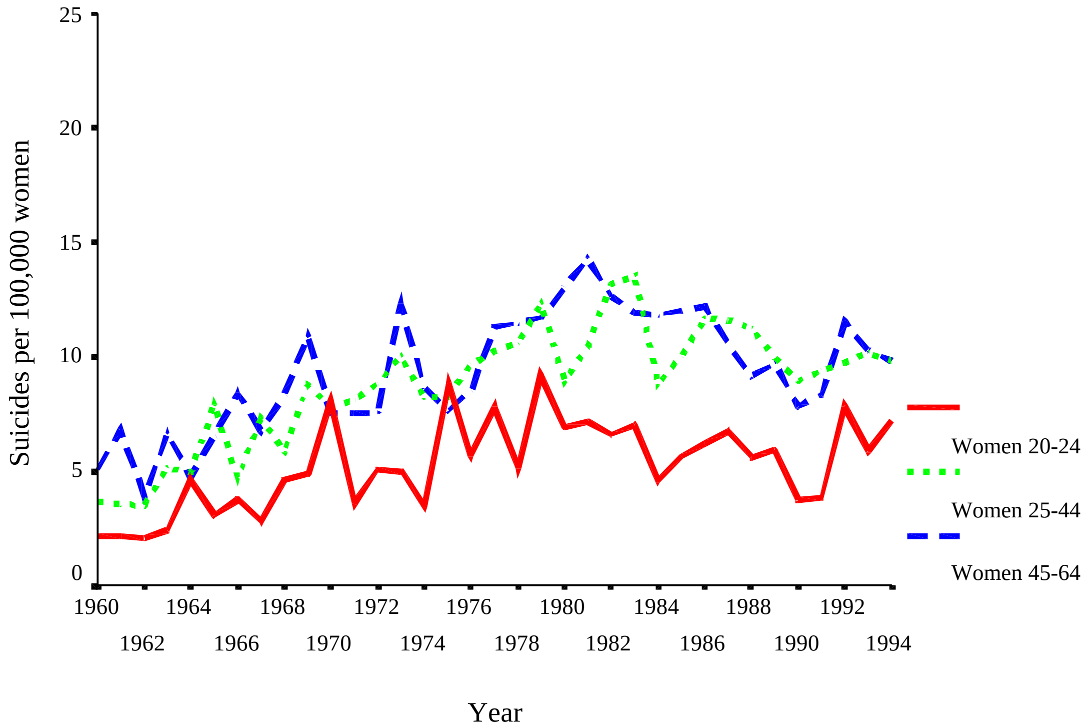


Table 3 F tests on structural break coefficients: Quebec women

	Null hypothesis		
Cohort	Intercept a1 = 0	Trend b1 = 0	Joint a1 = b1 = 0
Female 25–44	0.34	3.85	3.02
Female 45–64	1.98	1.45	0.68

Notes: ** denotes significance at 1 percent; * denotes significance at 5 percent.
Hypothesized structural break is 1975.

aboriginal males have a much higher propensity to commit suicide than white males (Statistics Canada 1997), reflecting a range of factors such as difficulties associated with integration into dominant Canadian society and the deterioration of traditional and cultural values (Health and Welfare Canada 1989). These factors, together with our short time series, suggest that it will be more difficult to test for the presence of a structural break. These fears are borne out in Figures 4 and 5, which provide age-specific suicide rates for women and men in British Columbia. The male series for all age cohorts appears to exhibit a rather high degree of variation when compared to both the Ontario males and the female series for all three provinces. Not surprisingly, with the exception of the female 25–44 and 45–64 age cohorts, the goodness-of-fit of equation (1) estimated for these data was much lower than that recorded for either adult males or females in Ontario.

Despite this weaker data, an interesting complication makes testing for a structural break worth pursuing. Rather than rushing in with stop-gap legislation, as was the case in Ontario, the British Columbia government established a Royal Commission in December 1973 to examine the whole issue of marital property. Its report, submitted in March 1975, “undoubtedly served to stimulate a desire to change the law governing matrimonial property” (McLeod and Mamo 1995, BC-3). The report led to the Family Relations Act, 1978. This legislation was broadly similar to that put into place in Ontario. Thus, although the change in legislation was passed in 1978 and came into force in 1979, as noted above, expectations of change were raised by the establishment of the Royal

Figure 4 Female suicide rates: British Columbia

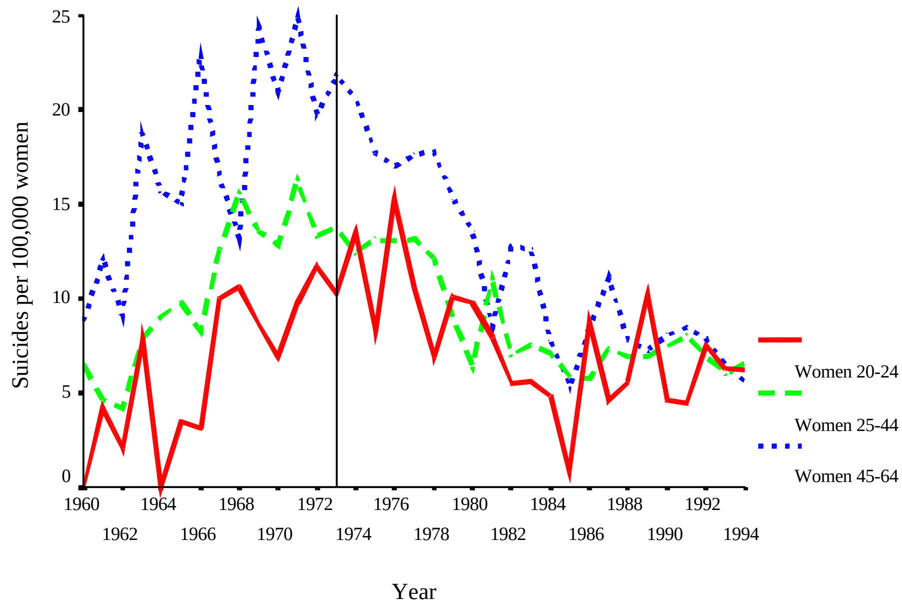
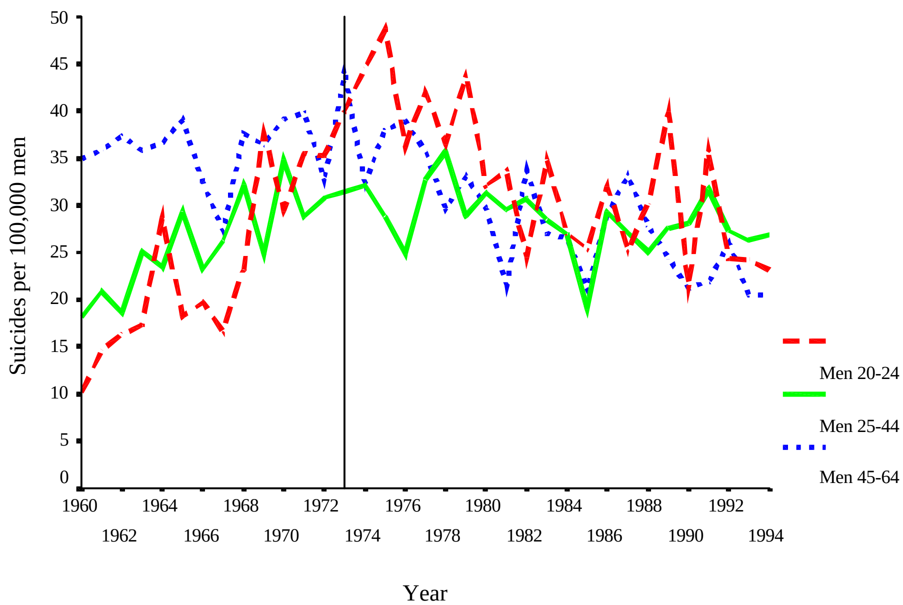


Figure 5 Male suicide rates: British Columbia



Commission in 1973. It is useful to recall the comment made by Dixit and Pindyck in their discussion of noneconomic applications to their model of investment under uncertainty. They write:

> Whatever its merits or demerits as descriptive theory, the Hamermesh-Soss model is *not rational enough* from the prescriptive viewpoint, because it forgets the option value of staying alive. Suicide is the ultimate irreversible act, and the future has a lot of ongoing uncertainty. Therefore the option value of waiting to see 'if something will improve' should be very large (Dixit and Pindyck 1994, 24).

The establishment of the Royal Commission can be thought of as something that generated the likelihood that "something will improve." This suggests that we would expect to see a break in the data slightly before actual legislation, as is the case in Ontario. We therefore propose that the break in the deterministic trend occurs from 1974 onwards. Mindful of our caveats regarding the data, we further hypothesize that no such break should occur for women aged 20–24 or men.

To save space, we report only the results of our summary test statistics in Tables 4 and 5. As in the case of Ontario, our F tests decisively reject the null hypothesis of no structural break in the trend rate of suicide for women aged 25–44 and 45–64. This result

Table 4 F tests on structural break coefficients: British Columbia women and men

Cohort	Null hypothesis: Intercept a1 = 0	Trend b1 = 0	Joint a1 = b1 = 0
Female 20–24	0.01	21.17**	11.55**
Female 25–44	9.55**	66.20**	30.84**
Female 45–64	5.35*	39.68**	19.91**
Male 20–24	2.16	46.89**	30.61**
Male 25–44	0.81	16.44**	8.29**
Male 45–64	0.48	8.47**	4.26*

Notes: ** denotes significance at 1 percent; * denotes significance at 5 percent. Hypothesized structural break is 1974.

Table 5 Cox encompassing tests for nonnested models: British Columbia women and men

Cohort	Null hypothesis: Model 1 encompasses Model 2	Null hypothesis: Model 2 encompasses Model 1
Women 20–24	–1.765	–4.455**
Women 25–44	1.877	–10.137**
Women 45–64	1.655	–8.886**
Men 20–24	–1.303	–3.433**
Men 25–44	–2.363*	–3.390**
Men 45–64	–2.603*	–1.836

Notes: ** indicates significance at the 1 percent level; * indicates significance at the 5 percent level. Cox test has standard normal distribution under the null hypothesis that Model 1 encompasses Model 2. Model 1: structural break occurs in 1975; Model 2: structural break occurs in 1980.

is consistent with our basic argument. But again like the results from Ontario, we also find evidence of structural breaks for individuals who we hypothesize are unaffected by these changes: younger women and men. As before, we infer the most likely break-point in the male data by examining the sequence of equation standard errors generated by estimating the model for the period 1970 to 1984. And a Cox test is performed under two specifications of the null hypothesis: (1) that the model specified with the structural break occurring in 1974 (Model 1) encompasses the same model specified with the structural break occurring in 1980 (Model 2), and (2) vice versa. The results provide mixed support for our hypothesis. On the one hand, the “early-breaking” model variance-dominates the “late-breaking” one for female suicide rates for age cohorts 25–44 and 45–64, lending further support to the Dixit and Pindyck inspired view that the *announcement* of an inquiry into divorce law reform in British Columbia best predicts the change in behavior. The results also lend support to the “late-breaking” model for the oldest male cohort. However, contrary to expectations, the early breaking model variance-dominates in the youngest cohorts for either sex and the results for men 25–44 are indeterminate.

5. CONCLUSION

The motivation for this paper lies in the difficulties faced by economists in devising discriminating empirical tests that distinguish unitary from collective models of the household. In this paper, we exploit a dramatic change in divorce law in Canada to test one variant of the collective approach, cooperative Nash-bargained household models.

We derive three specific tests: (1) the change in divorce law should only reduce suicides among married women; (2) it should have a larger effect on the suicide rates of married, older women; and (3) its impact is asymmetric. It should not affect the suicide rates of adult men. Using data from Ontario, the results of all three tests conform to our hypotheses. Further, we extend the tests to two other provinces. In Quebec, where no such legislative change took place, no structural break in the data is observed. The results for British Columbia also lend some support to the idea advanced by Dixit and Pindyck that the possibility that "something will improve" will reduce suicide. It is possible that the less conclusive results that we obtain for British Columbia may be a consequence of greater measurement error in the recording of suicides and the presence of a much larger Aboriginal population. Accordingly, we conclude that our results favor a variant of the collective approach over the unitary model.

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