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CHILD SUPPORT AND PARTNERSHIP DISSOLUTION*

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This article studies the determinants of partnership dissolution and focuses on the role of child support. We exploit the variation in child support liabilities and entitlements driven, in part, by the introduction of a new set of complex rules that determined child support liability in the UK, and by their interaction with welfare rules. Our panel has the virtue that the post 1993 variation in child support liability for all couples in partnerships formed prior to 1993 is unanticipated. We find strong evidence that the resulting large child support liabilities significantly *reduced* dissolution risks.

A growing economics literature of theoretical and applied research has been successful in promoting a better understanding of family behaviour in general – for recent surveys, see Weiss (1997) and Ermisch (2003) – and separation in particular; see Peters (1993) and Böheim and Ermisch (2001). However, despite the policy reforms that have aimed at increasing child support entitlements and compliance, little of the research that has been done has considered the impact of child support (CS) on separation. We would expect higher CS liabilities generally to imply smaller separation incentives for fathers and greater incentives for mothers so that the net effect is unclear *a priori*. However, CS often interacts with welfare receipts for poor households and, in some cases, CS payments may be tax deductible and hence will interact with the tax system. Thus, it will often be the case that net payments of CS may not equal net receipts and the difference will depend on individual circumstances in complicated ways. In general, because net payments and net receipts will not be equal, there will be some net implications of CS for the probability of parents separating and this article is specifically concerned with the empirical modelling of how CS affects separation. The main contribution of this article is to quantify the net effect of CS on separation incentives and so evaluate the implications that CS system design might have for separation rates.

CS has become a major policy issue in many countries partly because of the dramatic growth of separation amongst parents that has occurred since the 1970s. High rates of separation and lone parenthood, and low levels of CS, resulted in growing numbers of lone parents, almost all mothers, many of whom relied on welfare. A dramatic reform was introduced in the UK in 1993 which created a *Child Support Agency* (CSA) which, for the first time, mandated CS payments[1] for cases

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[1] A similar reform had been introduced in Australia in 1989 with an explicit formula and enforcement through the income tax authorities. Many changes have taken place across US states to increase entitlements of lone mothers and compliance has been improved through the creation of a federal compliance office.

who were entitled to welfare payments.[2] However, the levels of CS liabilities were often extremely high and accumulated arrears sometimes amounted to thousands of pounds. Moreover, the incentives for many lone parents to seek CS was limited because of the interaction between CS and the welfare system; and the rules that determined the obligations were complex and required many pieces of information from the non-resident parent which were difficult for the CSA to verify. Thus, the levels of compliance remained low and the costs of enforcement were high.[3]

Separation has typically been associated with a large drop in income for the parent with care (the custodial parent) and it is the purpose of obligatory CS to offset this fall. CS changes the nature of the payoffs to both spouses should separation occur. By raising the financial obligation on the non resident (absent) parent, almost always the father, CS raises the costs of separation to the absent parent.[4] However, CS also lowers the cost of separation to the custodial parent, almost invariably the mother.

The methodology we use is to estimate the determinants of the hazard of separation using a panel of couples who are at risk of separation in the following year. Assuming that couples form static expectations of their CS liabilities and receipts should separation occur, so that they suppose that those aspects of behaviour that affect CS do not change post-separation, we can compute expected CS liabilities and entitlements and so include these as determinants of separation.

Section 1 reviews the existing literature. Section 2 outlines the empirical specification. Section 3 presents the UK data and Section 4 focuses on the role of CS in partnership dissolution and explains how contemporaneous CS liabilities are constructed. Section 5 presents the results and their interpretation. Section 6 concludes and evaluates.

1. Existing Literature

There is an extensive literature that is concerned with the effect of welfare policies on separation. Moffitt (1992) surveys this literature and finds little support for the idea that separation is motivated by considerations of the potential welfare entitlements. However, very few papers have considered the role of CS. Hoffman and Duncan (1995) include predicted CS as a regressor in their model of separation using a very small subset of the US Panel Study of Income Dynamics data but find that it is statistically insignificant.[5] Recently Böheim and Ermisch (2001) study partnership dissolution in the UK using the first eight waves of the British Household Panel Survey (BHPS) and show that couples experiencing unexpected

[2] Cases not on welfare could be considered for a CSA assessment at the request of either parent. Thus, the CSA rules act as a focal point for other negotiated regular transfers and/or one-off settlement payments for non-welfare cases.

[3] A subsequent reform, that was not implemented until 2004, made the CS formula much simpler, reduced the interaction with the welfare system and reduced typical liability levels.

[4] Hereafter, we assume, for simplicity, that it is mothers who become the custodial parent, so it is fathers who are liable for CS.

[5] Several papers investigate the role of CS on *remarriage*. Yun (1992) finds a positive effect of the *availability* of CS but a negative effect of actual *payments*. Beller and Graham (1993) and Hu (1994) find no significant effect of CS. Carlson *et al.* (2003) investigates the effects of CS on partnership formation in a sample of unmarried mothers and finds no effects on marriage.

improvements in finances have significantly lower dissolution risks while couples experiencing negative shocks are at insignificantly higher risk.

However, the 'surprises' highlighted in Böheim and Ermisch (2001) only capture changes in a couple's economic circumstances *within* the partnership. When people decide whether to continue the partnership into $t + 1$ from t, they should compare their potential net incomes after partnership dissolution with the status quo rather than look at changes in net incomes within the partnership. In this work we assume that couples are forward looking rather than backward looking. Although the 'surprises' might well be one of the factors that determine the changes in net income arising from partnership dissolution, the former are nevertheless only a partial and indirect measure of the latter variable that directly enters the utility comparison framework.[6]

Indeed, CS is the key variable that links the net incomes before and after relationship dissolution for both partners. When we abstract from any labour supply or repartnership effects on incomes, CS is the main factor that determines the changes in net incomes caused by the marital dissolution. Other factors, such as child custody and housing arrangements, only affect changes in net incomes through their impact on CS liabilities and receipts. Only two papers directly address this issue: Nixon (1997) uses Current Population Survey data and finds a statistically significant and negative effect of CS enforcement on divorce; while Heim (2004) uses state-level data and exploits variation in CS enforcement over time and finds no significant effect. However, neither of these papers explores the complex relationship between CS, taxes and transfers which serve to make liabilities and receipts differ.

A simple theoretical rationale is provided by the cooperative model in Nixon (1997). If CS is just a simple transfer between separating partners then the effects will depend on the marginal utilities of consumption post separation relative to pre-separation. In the unlikely event that income prior to separation was pooled then we would expect the lower income of the mother to imply a higher marginal utility and so CS implies an aggregate net gain in welfare on separation. However CS is, in practice, not a simple transfer – the tax and welfare systems complicate matters. In one extreme, but common, case a separated mother with no other income will find that her CS receipt is means-tested at a rate of 100% so that the CS does not serve to increase her welfare at all and CS will therefore unambiguously reduce the likelihood of separation. While a non-cooperative model seems more natural for analysing divorce, the implications of such a model would be similar: if CS is a simple transfer (i.e. no tax or welfare effects) then both partners' threat points are affected and have offsetting effects on separation incentives, while the extreme case of receipt means tested at 100% implies an unambiguously lower probability of separation.[7]

[6] See Hoffman and Duncan (1995) and Weiss and Willis (1997) who use prediction errors from econometric estimates of one period ahead individual incomes.

[7] In the UK divorce does not require consent of both parties although consent implies that the proceedings are quicker. You have to show that your marriage has broken down irretrievably by proving that either your spouse has committed adultery, been cruel either mentally or physically, that you have been deserted for more than two years, or that you have lived apart for more than five years. You can agree with your spouse that your marriage is at an end and consent to a divorce provided you have lived apart for more than two years.

In this article we use the UK CS rules prevailing from 1993 to 2003 to calculate, under plausible assumptions, the estimated CS liability and the implied levels of receipt, for each time period that a couple is at risk of dissolution. Using the official adult equivalence scales (McClements, 1977) we then calculate the equalised net incomes for both partners pre and post dissolution of the relationship. The reform replaced *ad hoc* CS arrangements, which almost invariably resulted in little or no CS paid, with a set of rules where liability was a highly complex and non linear function of both partner's incomes and other variables. Liabilities were typically now large and varied considerably across individuals[8] and receipts differed from liabilities because of complex interactions with the tax and welfare systems.

2. Empirical Specification

We estimate both a discrete-time transition rate model and hazard models of separation. The discrete-time transition rate specification is used as a starting point, as it allows us to compare our results with those in the Böheim and Ermisch (2001) paper. We then extend the model by exploiting the variation in CS liabilities driven by an important policy reform, to separately identify the effects of children from the effect of CS.

The discrete-time transition rate model has the desirable property that the probability of survival at time period $t + 1$ only depends on survival probability up to period t and a vector of explanatory variables also measured at t. Jenkins (1995) has shown that once the total elapsed duration is included in the model, one can use a standard probit model to get consistent parameter estimates of the determinants of the explanatory variables on the hazard.

In addition to estimating a simple transition rate model we also estimate two of the most popular parametric survival distributions, namely the Exponential and the Weibull parameterisations, which allow for no duration dependence and monotonic duration dependence respectively. The two parameter Weibull distribution is an important generalisation of the Exponential distribution, to allow for a duration dependence of the hazard. The hazard function of the Weibull function is given by $\lambda(t) = \gamma p t^{p-1}$ where $\gamma > 0$ and $p > 0$. The shape of the hazard function depends critically on the value of p, which is sometimes known as the shape parameter; for details see, for example, Kalbfleisch and Prentice (2002) and Stata Corp (2003). The hazard is monotonically decreasing for $p < 1$, increasing for $p > 1$, and reduces to the constant exponential hazard if $p = 1$.

3. Data

The British Household Panel Survey (BHPS) is a nationally representative sample of some 5,500 households recruited in 1991, with around 10,000 original sample members (OSMs). These OSMs and their children, who also become sample members after reaching 16, are interviewed each successive year, together with all adult members of their families, even if the OSMs split off from their

[8] See Paull *et al.* (2000) for details.

original households to form new families and/or relocate to other areas of the UK. This sampling design ensures that the sample remains representative of the UK population over time. The core questionnaire of BHPS collects information on household organisation, housing, employment, education, health and incomes in all waves. In wave 2, BHPS also collected lifetime histories of marriage and cohabitation, which allow us to construct spells in progress of the current relationship for all couples in our sample, despite the fact that we are unable to observe the partnerships from the time of their formation.

The sample in this article includes all married or cohabiting couples with qualifying children in the BHPS,[9] at the time of the first wave, and adds all couples who have new children up to wave 12. For people experiencing multiple relationship dissolutions over the sample period, we only focus on the first relationship. We include all cases where the couples are at risk of partnership dissolution in the forthcoming year (i.e. period $t + 1$) and where the outcome can be either directly observed or imputed with certainty (that is, providing we can follow at least one of the partners post-separation and see a change in their partnership status, or a change in the identity of their partner). All explanatory variables are measured at period t. This leaves us with 7,834 couple-years, of which 219 couples (2.80%) end up in dissolution. Our aim throughout is to focus on CS and other economic variables.

Table 1 gives the means and standard deviations of all variables broken down by partnership outcomes used in the preferred specification of Böheim and Ermisch (2001), which will serve as a benchmark for our empirical analysis. For presentational purposes, we choose the woman as the representative for a couple.

Table 1 shows that there is not much difference between the net weekly earnings of women who are about to separate and those who will remain in partnerships. In contrast, women in stable partnerships have partners with higher earnings than those who are about to separate, partly reflecting the difference in participation rates of the partners. As expected, unemployment is associated with higher risk of partnership dissolution for both men and women (the omitted category being inactivity).

Cohabiting couples are much more likely to separate than legally married couples. This huge difference might reflect the difference in the level of commitment, or it might be due to difference in observable characteristics between these two groups despite the fact that they have all had children. Note that for CS purposes, married and cohabiting couples are treated equally. The Table also suggests that women who start a partnership later in life are slightly less likely to dissolve their partnership, while the elapsed partnership duration is negatively correlated with the risk of separation. The first finding seems to be consistent with the theoretical prediction that people who enter into a relationship early are more likely to regret the poor match arising from insufficient search or it may reflect a low probability of a rematch. The indication that the probability of a partnership dissolving declines with elapsed partnership duration might reflect either

[9] Our sample excludes couples whose children have either become too old to qualify for CS or who have had children that have now left the parental homes. Thus we abstract from any possible 'empty nest' effect where the exit of children from the household that contributes to the separation hazard. None of the existing empty nest research allows for the potential endogeneity of leaving home. Our results when we include this group, which results in a much larger sample, are essentially unchanged.

Table 1

Means SD of Explanatory Variables by Partnership Outcome

	Continuing	Dissolving	All
Incomes			
Earnings	1.052 (1.332)	0.870 (1.148)	1.047 (1.327)
Partner's earnings	3.032 (2.972)	2.089 (2.220)	3.006 (1.327)
Partnership characteristics:			
Cohabiting	0.054 (0.226)	0.228 (0.421)	0.059 (0.235)
Number of ex-marriages	0.204 (0.599)	0.169 (0.519)	0.203 (0.597)
Age at start of partnership	23.009 (4.625)	21.813 (3.712)	22.975 (4.606)
Log duration of partnership	2.510 (0.618)	2.122 (0.845)	2.499 (0.629)
Partners same ethnic group	0.951 (0.215)	0.900 (0.301)	0.950 (0.218)
Partners have same religion	0.492 (0.500)	0.429 (0.496)	0.490 (0.500)
Partners not religious	0.342 (0.475)	0.384 (0.487)	0.344 (0.475)
Youngest child <5 years	0.396 (0.489)	0.534 (0.500)	0.400 (0.490)
Number of children	1.841 (0.819)	1.886 (0.914)	1.843 (0.822)
Partners different education	0.770 (0.421)	0.804 (0.398)	0.771 (0.420)
Age difference			
Woman 5+ years older	0.025 (0.156)	0.059 (0.237)	0.026 (0.159)
Woman 3–5 years older	0.048 (0.214)	0.046 (0.209)	0.048 (0.214)
Woman 0–3 years older	0.214 (0.410)	0.183 (0.387)	0.213 (0.410)
Partner 2 to 4 years older	0.350 (0.477)	0.333 (0.472)	0.349 (0.477)
Partner 4+ years older	0.230 (0.421)	0.242 (0.429)	0.230 (0.421)
Labour Market:			
Employed	0.667 (0.471)	0.584 (0.494)	0.665 (0.472)
Unemployed	0.010 (0.101)	0.023 (0.150)	0.011 (0.103)
Partner employed	0.855 (0.352)	0.753 (0.432)	0.852 (0.355)
Partner unemployed	0.051 (0.220)	0.096 (0.295)	0.052 (0.222)
Surprise indicators			
Large positive surprise	0.015 (0.124)	0.005 (0.068)	0.015 (0.122)
Positive surprise	0.164 (0.370)	0.132 (0.340)	0.163 (0.369)
Negative surprise	0.229 (0.420)	0.269 (0.445)	0.230 (0.421)
Large negative surprise	0.055 (0.228)	0.110 (0.313)	0.056 (0.231)
Missing surprise indicator	0.056 (0.230)	0.068 (0.253)	0.056 (0.231)
Sample size (%)	7,615 (97.2%)	219 (2.8%)	7,834 100.0%)

Note. Earnings are in £100/week and in January 2004 prices.

heterogeneity, say in risk aversion, or the hypothesis that couples invest in partnership-specific capital over time.

Table 1 also shows that couples of the same ethnic group or religion are less likely to separate, a result consistent with the hypothesis of positive sorting by marriage. The presence of pre-school children is associated with higher risks. However, this is just a simple correlation and might simply reflect the fact that households with younger children tend to have shorter relationship durations.

Following Böheim and Ermisch (2001), we construct 'surprise' variables, by comparing people's expectations formed at $t-1$ of their (family's) financial situation at t with their evaluation of the actual outcomes at t, in order to test the hypothesis that new information affects partnership dissolution. Table 2 shows how the five 'surprise' categories, i.e. large positive surprises, positive surprise, as expected, negative surprise and large negative surprises are defined respectively, together with the corresponding relative frequencies. There is a monotonic increase in the probability of partnership dissolution as we move from large positive to large negative surprises. For instance, only 0.8% of couples experiencing large

Table 2

Dissolution Rates by Expected and Realised Financial Situation ($N = 7{,}369$)[10]

	Realisation$_t$		
Expectation$_{t-1}$	Better off	About the same	Worse off
Better off	3.3% = (13.2%)	3.3% − (11.2%)	5.4% − (6.0%)
About the same	2.4% + (13.9%)	2.1% = (32.9%)	3.3% − (13.3%)
Worse off	0.8% ++ (1.6%)	1.6% + (3.4%)	2.7% = (4.6%)

Note: ++ large positive surprise, + positive surprise, = as expected, − negative surprise, – large negative surprise. Numbers in parentheses are relative frequencies.

positive surprises will end up in separation in the following year, compared to 5.4% of couples with large negative surprises.

4. Child Support

Concern about growing child poverty has motivated recent research on the impact of partnership dissolution on the incomes of households with children and on child welfare. The overwhelming evidence from the US has indicated a positive role for CS in reducing child poverty among lone parent families; see e.g. Bartfeld (2000), Del Boca and Flinn (1995), Meyer and Hu (1999) and Meyer (1993). In the UK, Bingley *et al.* (1995*a,b*) investigate the potential effects of the CS system on net incomes and labour supply of lone mother headed households.[11]

For the vast majority of couples at risk of partnership dissolution in our sample we do not observe what would happen to them should separation take place. So we make some naive, but plausible, assumptions in our CS liability calculations:

(1) We abstract from any labour supply and repartnership effects and assume no implications for travel-to-work costs;
(2) Mother gets custody of all children (and so is referred to as the parent with care (*PWC*)) and stays in the original house, while the father becomes the non-resident parent (*NRP*), moves to a rented apartment, with rent set at the median of all rented housing of the region in that year;[12]
(3) Both *PWC*s and *NRP*s welfare benefit receipts upon separation are predicted under the previous two assumptions;
(4) Finally, CS liability is calculated under the system of CS described above, based on observed earnings and hours, observed/imputed housing costs for the *NRP*/*PWC* and predicted welfare benefit receipts from step (3).

[10] This sample is slightly smaller than those in all other Tables with $N = 7{,}834$, due to the exclusion of families with missing surprise variables. However, we include an indicator for missing surprise in all model specifications. Note also that we have imposed the 3 consecutive wave requirement at the outset in order to facilitate easy comparison and interpretation.

[11] Paull *et al.* (2000) investigate the potential effects of the subsequent reform to the CS system, which was not implemented until 2004 on net incomes and labour supplies of lone mother headed households.

[12] According to our own estimates, fathers account for less than 7% of lone parents in the BHPS.

While these assumptions are obviously abstractions, we would argue that the CS liability, and implied entitlement, derived in this way could be regarded by partners as a reasonable expectation resulting from a simple rule-of-thumb.[13]

Table 3 decomposes household incomes into earnings, benefit income and other incomes for both partners pre and post separation. It also shows equivalised incomes for *PWC* and *NRP* under both scenarios, using before housing cost (*BHC*) and after housing cost (*AHC*) scales. Our sample of couples with dependent children have a mean weekly total net income of £452 in January 2004 prices, with 23.1% and 66.4% coming from women and men's labour income respectively, 8.7% from benefits and 1.8% from all other income. With a mean equivalence scale of 1.41, this results in an equivalised income of £328 for the family before housing cost. After deducting the reported housing costs with a mean of £76 and using the alternative equivalence scale, we get a mean equivalised income of £273 after housing cost. The *PWC* and the children will suffer a loss of equivalised income in the magnitude of 22% or 31% on average following separation, depending on whether we use the *BHC* or *AHC* measure, despite a 183% increase in total social security transfers and full compliance of CS of the *NRP*s. Note that, on average, *PWC*s only benefit from less than half of the CS paid by the *NRP*s (i.e. £33 out of £69), mainly due to the fact that the Income Support (IS) system, that provides income for those out-of-work, imposes a 100% tax on CS receipts in excess of the level of IS entitlement. On the other hand, *NRP*s are better off on both measures of equivalised income post separation, with a net gain in the magnitude of 25%–40%.

Table 4 assesses the goodness of fit of our CS entitlement prediction, by exploiting the subsample of couples who are observed both before and after separation. We restrict our sample to all matched couples entitled to non-zero CS in the first two waves immediately after separation, as in Hoffman and Duncan (1995). This seems a reasonable restriction given that we are assuming that couples forecast their CS using static expectations.[14] It is clear that our CS routine has done a reasonably good job of predicting CS entitlement for those with compliant ex-partners, and the estimated compliance rate is just under 60%, in line with official statistics.[15] For the subgroup of compliant (fully or

[13] In principle, we should use the present value of the expected total CS liabilities for each *NRP*, which also depends on his discount rate and age structure of the qualifying children (note that CS payment ceases when a qualifying child reaches 16, or up to 18 if he/she stays on school) and even on planned, but yet unborn, children. In practice, we found that only current liabilities were empirically important and reserve this for more detailed further work.

[14] Evidence based on the subsample of matched couples in the BHPS (not shown here) suggests that there is very little change in employment status and repartnership remain at a relatively low level (20% for *PWC*s and 29% for *NRP*s) up to 2 years post separation.

[15] We found that there are essentially no observable characteristics in BHPS that explain compliance although the size of liability is significant. However, we cannot use this estimated compliance function to implement a propensity score because compliance is only observed for those who have separated. Moreover, the separation hazard contains no exclusion restrictions that would allow our compliance equation to be used as a control function for heterogeneity. Thus our estimated effect of CS on separation should be interpreted as a lower bound on the causal effect. We experimented with child gender, following recent work by Dahl and Moretti (2004), but found it to be entirely insignificant in both the separation and compliance equations.

Table 3

Mean Equivalised Household Incomes for PWC (and Children) and NRP, Pre and Post Separation, by Sources of Income (N = 7,834)

	Mother with children		Non-resident father	
	Amount	%	Amount	%
Pre-separation:				
Own net earnings	104.70	23.1	300.56	66.4
Partner's net earning	300.56	66.4	104.70	23.1
Total net benefit	39.15	8.7	39.15	8.7
Other income	8.09	1.8	8.09	1.8
Total net income	452.49	100.0	452.49	100.0
Equivalence scale (*BHC*)	1.41		1.41	
Equivalised income (*BHC*)	328.10		328.10	
Equivalence scale (*AHC*)	1.42		1.42	
Housing cost	75.66		75.66	
Equivalised income (*AHC*)	272.84		272.84	
Post-separation:				
Own net earnings	104.70	41.7	300.56	120.0
Partner's net earning	–		–	
Total net benefit	110.86	44.1	13.17	5.3
Other income	2.28	0.9	5.81	2.3
Child support	33.33	13.3	–68.98	−27.5
Total net income	251.16	100.0	250.56	100.0
Equivalence scale (*BHC*)	1.02		0.61	
Equivalised income (*BHC*)	254.55		410.75	
Housing cost	75.66		39.73	
Equivalence scale (*AHC*)	0.97		0.55	
Equivalised income (*AHC*)	188.95		383.32	

Note. *AHC* = after housing costs, *BHC* = before housing costs. All monetary figures are in £/week and in January 2004 prices.

Table 4

Comparing Predicted Child Support Entitlement with Actual Payments/Receipts using the Matched Couple Sample (N = 184)

	NRP Paying	*NRP* Not Paying
Predicted CS entitlement	97.26	71.51
Actual payment	96.46	–
N	106	78
%	57.6	42.4

Note. £/week Jan 2004 prices. Sample based on all matched couples entitled to positive CS in the first two waves immediately after separation.

partially) *NRP*s the mean weekly difference between predicted entitlement and actual payment is less than £1.[16]

It is clear that the way in which CS interacts with the tax and welfare system is also important. A major part of a subsequent CS reform, implemented in 2004,

[16] It is interesting to note that of the subsample of *NRP*s reporting paying CS, only 57% of their ex-partners actually report receiving CS, and even conditional on receiving, the mean received amount is one sixth less than the amount paid. This might reflect under-reporting of CS receipts by *PWC*s, and it is also possible that some *PWC*s only perceive CS receipts net of social security benefits as real CS.

Table 5
Means SD of Further Explanatory Variables by Partnership Outcome

	Continuing	Dissolving	All
Incomes			
Total Net Household Income	4.551	3.615	4.525
	(3.306)	(2.392)	(3.287)
Child support related variables			
Current CS liability	0.695	0.502	0.690
	(0.514)	(0.528)	(0.515)
Indicator for wife on IS if separated	0.426	0.527	0.429
	(0.495)	(0.501)	(0.495)
CS × Indicator for separated wife on IS	0.292	0.245	0.291
	(0.490)	(0.423)	(0.488)
Indicator for wife on FC if separated	0.287	0.265	0.286
	(0.452)	(0.442)	(0.452)
CS × Indicator for separated wife on FC	0.192	0.158	0.191
	(0.391)	(0.425)	(0.392)
Indicator for post 91 partnership	0.095	0.210	0.098
	(0.293)	(0.408)	(0.297)
CS × post 91 partnership	0.059	0.103	0.060
	(0.227)	(0.274)	(0.228)
Mother 40+ at wave 1	0.178	0.059	0.174
	(0.382)	(0.237)	(0.379)
CS × Mother 40+ at wave 1	0.120	0.035	0.118
	(0.348)	(0.219)	(0.346)

Note. Incomes and CS liabilities are in £100/week and in January 2004 prices.

dealt with the benefit disregards for receipt of CS and introduced a £10 disregard for Income Support and increased the Family Credit (FC) disregard of £15 such that Working Families Tax Credit (WFTC) which replaced FC in 1999 would disregard *all* CS payments no matter how large.

Table 5 gives the means and standard deviations of further explanatory variables (in addition to those in Table 1) we are going to use in our empirical modelling. We control for total net household income in all our specifications and so we identify the effect of CS through the introduction of the CS system with its complex nonlinearities and interactions. We cannot use CS receipt since this is only observed for the separated. We capture the effect of the welfare system on receipts by interacting liabilities with dummies for being eligible for in-work and out-of-work welfare programmes.[17] A further difficulty is that CS could affect fertility as well as separation. To examine the sensitivity of our estimates we distinguish between couples who have completed their fertility at the time when the CS reform occurs (mothers who are 40+ at wave 1)[18] and those who were younger. We allow

[17] We assume full compliance throughout our analysis although according to Table 4 only about 60% pay any CS and those that do pay seem to pay the full liability on average. Note that the real compliance rate could be higher than our estimate as arrears of maintenance due will be pursued by the CSA forcibly. It may be more reasonable to assume that separation depends on the expected payment and receipt. Omitted non-compliance is likely to be positively correlated with heterogeneity in the separation rate and this is likely to bias our estimates CS effect downwards (towards zero).

[18] Case (1998) employs an IV strategy and finds a sizable but insignificant effect of CS enforcement on fertility. Aizer and McLanahan (2004) suggest that stricter CS enforcement might affect the type of women who give birth out of wedlock.

for the effect of CS on separation to differ for couples which formed post-reform from pre-existing couples for whom the reform was a surprise.

Table 5 shows that the mean CS liability is £69 while the mean total net income is £452 per week. Families with lower total net incomes and lower CS liabilities are more likely to separate. Women who are predicted to receive IS are associated with a higher risk of separation while the converse is true for the predicted recipients of FC (or WFTC). Post-reform partnerships are associated with higher risks of dissolution, while women who have completed their fertility at the beginning of the sample are associated with lower risks. Note that due to the 3 consecutive wave requirement, pre-existing and new couples are not at risk of separation until 1992 and 1993 respectively. This, combined with the fact that partnership outcomes are based on information from year $t + 1$, ensures that all separations in our sample take place in 1993, the year the 1991 Child Support Act was implemented, at the earliest.

5. Estimation Results

Table 6 presents five specifications of the discrete-time transition rate model, starting with the Böheim and Ermisch (2001) specification which serves as a benchmark. To facilitate model evaluation and selection, we report marginal effects and p-values.

Model 0 is the Böheim and Ermisch (2001) specification and we find that only negative surprises are statistically significant for separation. Model 1 is a parsimonious specification which, nonetheless, nests the Böheim and Ermisch specification[19] and includes the calculated CS liability, together with dummies for wife's predicted benefit status (in-work and out-of-work benefits respectively) and post-91 sample, and the mother being 40+ in wave 1. Subsequent models introduce various CS interactions.

Our important finding is that while the effect of total net income is statistically insignificant and small in magnitude, the CS liability has a large negative effect on the hazard of separation and is significant at the 5% level in all specifications. Once we allow for interaction with the welfare system, the size of this effect is also remarkably stable across different specifications. A £100 per week increase in CS liability will reduce separation risk by 1.7 percentage point per year in our most comprehensive model specification, implying the introduction of the CS would result in a 42% drop in separation risk for a family with the mean CS liability of £69 per week at the sample mean (i.e. $0.690 \times 0.017/0.028$).

The benefit dummies and their interactions with child support liability appear to be insignificant and this, together with the stability of the CS effect, helps allay our concerns over the sensitivity of the effect of CS to endogenous fertility and to endogenous partnership formation.

[19] As neither partner's earning is statistically significant in Model 0, we use the more parsimonious total net household income in subsequent models to capture a separate income effect that is independent of any CS effect.

Table 6

Probit Model of Partnership Dissolution: Changes in Probability

	Model 0	Model 1	Model 2	Model 3	Model 4
Incomes					
Total Net Household Income (£100/wk)	–	0.0003 (0.426)	0.0003 (0.479)	0.0003 (0.469)	0.0003 (0.472)
Labour income (£100/wk)	−0.0001 (0.931)	–	–	–	–
Partner's labour income (£100/wk)	−0.0016 (0.113)	–	–	–	–
Child support related variables					
Current CS liability (£100/wk)	–	−0.009 (0.045)	−0.016 (0.026)	−0.016 (0.025)	−0.017 (0.019)
Indicator for wife on IS if separated	–	0.004 (0.304)	0.002 (0.802)	0.001 (0.811)	0.001 (0.841)
CS × Indicator for wife on IS if separated	–	–	0.005 (0.546)	0.005 (0.548)	0.005 (0.504)
Indicator for wife on FC if separated	–	0.002 (0.657)	−0.007 (0.329)	−0.007 (0.319)	−0.007 (0.326)
CS ×Indicator for wife on FC if separated	–	–	0.015 (0.129)	0.015 (0.125)	0.015 (0.122)
Indicator for post 91 partnership	–	−0.001 (0.883)	–	−0.003 (0.671)	−0.003 (0.690)
CS × post 91 partnership	–	–	–	0.029 (0.752)	0.040 (0.665)
Mother 40+ at wave 1	–	−0.006 (0.252)	–	–	−0.010 (0.173)
CS × Mother 40+ at wave 1	–	–	–	–	0.091 (0.412)
Surprise indicators					
Large positive surprise	−0.016 (0.130)	−0.016 (0.136)	−0.015 (0.141)	−0.015 (0.140)	−0.015 (0.134)
Positive surprise	−0.004 (0.373)	−0.004 (0.377)	−0.004 (0.346)	−0.004 (0.342)	−0.004 (0.363)
Negative surprise	0.007 (0.069)	0.007 (0.076)	0.007 (0.074)	0.007 (0.074)	0.007 (0.075)
Large negative surprise	0.018 (0.015)	0.017 (0.017)	0.017 (0.019)	0.017 (0.019)	0.017 (0.017)
Missing surprise indicator	0.005 (0.471)	0.005 (0.457)	0.005 (0.452)	0.005 (0.453)	0.005 (0.447)
Chi-square (df)	187.30 (26)	186.98 (30)	195.54 (30)	196.23 (32)	195.40 (34)
Pseudo R^2	0.0968	0.0982	0.0991	0.0992	0.1003
Log Pseudo-likelihood	−902.59	−901.14	−900.28	−900.20	−899.05

Note. $N = 7{,}834$. Rather than reporting coefficients, we report the change in the probability for an infinitesimal change in each independent, continuous variable and, by default, the discrete change in the probability for dummy variables. p-values in parentheses are adjusted to allow for multiple observations per couple. Other control variables included partnership characteristics, age difference dummies and labour market status as in Böheim and Ermisch (2001) (for details see Table 1).

We apply our most comprehensive specification to a duration model framework, which is less restrictive in its distributional assumptions. To facilitate easy comparison and interpretation, both the Exponential and the Weibull model in Table 7 are fitted in the log relative-hazard metric. The two sets of results are remarkably similar, implying that the Exponential is a very good approximation to the more general Weibull model. The results are also highly consistent with the discrete time specification. A £100 per week increase in CS liability will cut the hazard of separation by 62%. The dummy for mother being 40+ is now significant in both specifications. However, all CS interactions remain statistically insignificant, particularly the interaction with being a 40+ mother at Wave 1 which we feel captures the surprise effect of CS for those with completed fertility relative to potentially incomplete. This further strengthens our belief that effect of CS is robust to the endogeneity of partnership formation and fertility. The estimated shape parameter p is less than unity indicating negative overall duration dependence. However we can not reject the null hypothesis of no duration dependence at the conventional level of significance.[20]

Figure 1 shows the likely impact of CS reform on partnership dissolution using the Weibull estimates evaluated at sample means of all explanatory variables. The solid line indicates survival rates evaluated where CS = 0, the dotted line shows the predicted survival function under the new CS system which has only been enforced from 2004 (and hence out of our sample period), while the dashed line shows the predicted survival function under the CS system that prevailed from 1993 onwards. It suggests that the introduction of mandatory CS might have had an (unintended) impact on the separation rate, potentially reducing the separation probability by around 20% (for a 20 year old marriage) if all CS liabilities are fully enforced. On the other hand, the 2004 reform seems likely to reverse this tendency partially through reducing typical CS liabilities and introducing an IS disregard for CS.

The instantaneous hazards evaluated at the sample mean under the assumptions of zero CS, the old CS system and the new CS system are 5.10%, 2.80% and 3.12% respectively. Note that it is no coincidence that the instantaneous hazard under the old CS system is the same as the actual separation rate since we are applying the same rules to the whole sample and assuming full compliance in our estimation. These results imply that the introduction of CS (compared to no CS at all – which was quite typical prior to 1993) has decreased the instantaneous hazard by 45% over what we predict it would have been. However, one should bear in mind that this is an out of sample prediction. Moreover our estimates are for a linear effect, and hence rule out the possibility of diminishing returns. Simulation results using more realistic assumptions of compliance rates (say 50%) suggest that the effect of introducing the 1993 system from zero CS in 1992 is likely to have reduced separation risks for couples with dependent children by some 20%. In contrast, the

[20] We prefer the most comprehensive specification, as failure to take into account the interaction of CS with the welfare system will result in spurious negative duration dependence in the duration models.

Table 7

The Exponential and the Weibull Models

	Exponential Model		Weibull Model	
	Hazard Ratio	Robust p-value	Hazard Ratio	Robust p-value
Income				
Total Net Household Income (£100/wk)	1.015	0.407	1.015	0.407
Child support related variables				
Current CS liability (£100/wk)	0.381	0.009	0.381	0.009
Indicator for wife on IS if separated	1.073	0.788	1.067	0.805
CS × Indicator for wife on IS if separated	1.471	0.320	1.473	0.318
Indicator for wife on FC if separated	0.786	0.479	0.779	0.465
CS × Indicator for wife on FC if separated	2.113	0.148	2.122	0.145
Indicator for post 91 partnership	1.750	0.052	1.647	0.120
CS × post 91 partnership	1.101	0.814	1.115	0.788
Mother 40+ at wave 1	0.363	0.039	0.371	0.043
CS × Mother 40+ at wave 1	1.514	0.527	1.511	0.528
Surprise indicators				
Large positive surprise	0.313	0.227	0.313	0.229
Positive surprise	0.867	0.492	0.865	0.481
Negative surprise	1.316	0.096	1.315	0.096
Large negative surprise	1.755	0.014	1.743	0.015
Missing surprise indicator	1.235	0.426	1.235	0.426
P (H_0: $p = 1$)	–	–	0.950	0.583
Chi-squared (df)	219.12 (33)		170.43 (33)	
Log Pseudo-Likelihood	−449.01		−448.92	

Note: $N = 7{,}834$. Wald tests based on the Generalised Gamma Model; see Kalbfleisch and Prentice (2002); which nests both the Exponential and the Weibull model as special cases can reject neither. p-values in parentheses are adjusted to allow for multiple observations per couple. Other control variables included partnership characteristics, age difference dummies and labour market status as in Böheim and Ermisch (2001) (for details see Table 1).

introduction of the new CS reform is likely to increase the hazard by around 10% from its current level, assuming no change in compliance rates.

6. Conclusions

This article studies the determinants of partnership dissolution in the UK using the British Household Panel Survey. After allowing for heterogeneity in partnership characteristics, we still find couples to be highly responsive to changes in economic circumstances in deciding whether to continue their partnership. In line with previous studies we find that new information with regard to household finances has a substantial impact on the probability of partnership dissolution. In addition, we find that the variation in CS liabilities arising from the introduction of complex rules for CS, that surprised partners in 1993, had important implications for their subsequent separation experience. We find there is very strong evidence that an increase in the implied CS liabilities significantly reduced the dissolution risk. This result still holds when we allow for the interaction of CS with the welfare system, and potential endogeneity of fertility and partnership formation.

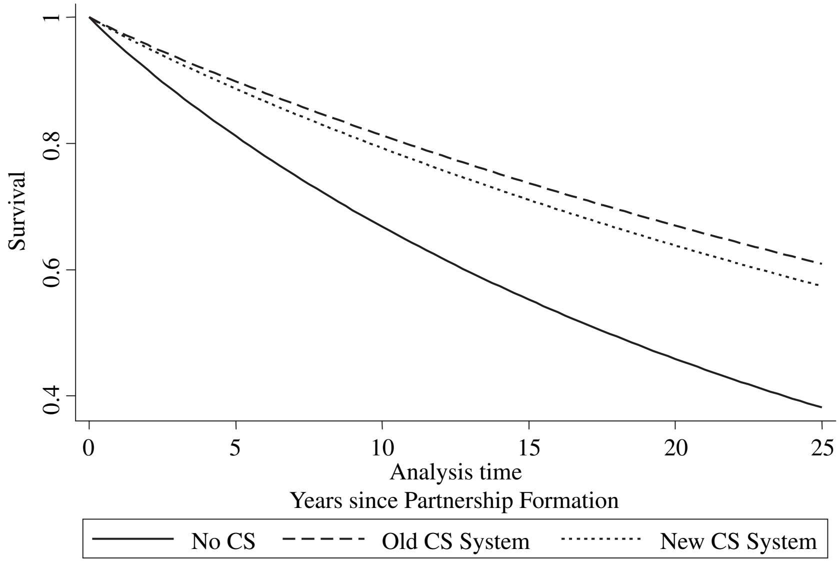

Fig. 1. *Predicted Impact of CS Reforms on the Survival Rate using Weibull Estimates*

We use the estimates to simulate the effect of CS on separation rates on the sample of couples with dependent children and find that this is dramatic. We calculate that the introduction of the 1993 system from zero CS in 1992 is likely to have reduced separation risks for couples with dependent children by over 20% under plausible assumptions of compliance rates. On the other hand, the 2004 reform will have the opposite effect, with an increase in hazard in the magnitude of 10%.

A natural extension in the future could take into account the labour supply and repartnership effects of dissolved couples, using the matched parent-with-care and non-resident-parent sample.[21] The assumptions of no labour supply or repartnership effects are maintained hypotheses but could also be tested. But despite our reservations about these assumptions we believe these existing findings do have significant policy implications. For instance, our results suggest that the 2004 child support reform and the CS pass-through that has been a feature of CS design in some US states, might have effects on separation rates through changing CS liabilities and receipts that are largely unintended. Aizer and McLanahan (2004) find that the risk of becoming a never-partnered mother is not significantly affected by CS but the logic of this present article suggests that CS would have effects on the joint risks of fertility and separation for the currently partnered, and we have only considered the risks of separation conditional on children being present.

Finally, while we have concentrated on the effect of CS on partnership dissolution we have not discussed the implications for the welfare of the parties concerned. It is unclear that, by holding a partnership together that would otherwise dissolve, the welfare of all parties has improved. There is little research on the

[21] Currently the sample in BHPS with matched separated mother-father information is probably too small to support such work, although we anticipate that would be possible after a few more waves.

impact of separation on well-being of the partners (Wilson and Oswald, 2005) and further research needs to be done to separate out the effects of separation *per se* from its financial consequences. Moreover, very little research has focussed on the outcomes for children, including their well-being.[22] Our current research (Walker and Zhu, (2005) is exploring the effects on children.

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[22] Some research suggests that CS has beneficial effects on outcomes for the children that exceed that of other forms of income. See, for example, Garfinkel *et al.* (1994). Recent work by Piketty (2003) and Gruber (2004) exploits natural experiments to estimate the impact of divorce on children. However, their findings are contradictory and further research is needed.

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