CHILD SUPPORT LIABILITY AND PARTNERSHIP DISSOLUTION

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Abstract

This paper studies the determinants of partnership dissolution and focuses on the role of child support. We exploit the variation in child support liabilities driven by an important UK policy reform to separately identify the effects of children from the effect of child support liability. We find strong evidence that an increase in the child support liability significantly *reduces* dissolution risk. Our results suggest that child support criteria that are based on the non-custodial parent's income, compared to criteria based on aggregate incomes of both parents, would imply much smaller separation rates.

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1. Introduction

It has been more than a quarter of a century since Becker, Landes and Michael (1977) published their pioneering study on the economics of marriage. The main implication of their theory is that maximization of marital incomes by men and women would induce strong segregation in the marriage market. Since then, a growing economics literature of theoretical and applied research has been successful in promoting a better understanding of family behaviour (for recent surveys, see Weiss (1997) and Ermisch (2003)). However, despite this well developed theoretical framework, and the motivation provided by the growth in divorce that has occurred over time across many countries, there are surprisingly few empirical studies of the determinants of partnership dissolution.

In recent years, partly because of the dramatic growth of divorce amongst parents, child support (CS) has become a major policy issue. High rates of lone parenthood and low levels of child support have resulted in growing numbers of lone parents, almost all mothers, who rely on welfare. A dramatic reform was introduced in the UK in 1993 which created a Child Support Agency which, for the first time, mandated child support payments. However, the levels of child support liabilities were often extremely high and accumulated arrears frequently amounted to thousands of pounds. Moreover, the reform was implemented in a way that made no allowance for earlier agreed settlements, the incentives for many lone parents to seek child support was limited because of the interaction between CS and the welfare system, and the rules that determined the obligations were complex. Thus, the levels of compliance remained low and the costs of enforcement were high. A subsequent reform, that was not implemented until 2003, made the CS formula much simpler, reduced the interaction with the welfare system, and reduced typical liability levels.

Separation has typically been associated with a large drop in income for the custodial parent and it is the purpose of obligatory child support to offset this. In Walker and Zhu (2003) we show how separation affects the distribution of equivalised incomes between parents and show how the level of child support requirements, and compliance with them, affects this redistribution. However, child support not only changes the nature of the payoffs to spouses should separation occur.

By raising the financial obligation of the absent parent, almost always the father, child support raises the costs of separation to the absent parent[1]. However, child support also lowers the cost of separation to the custodial parent, almost invariably the mother. Thus, in addition to providing for a redistribution of resources should separation occur, child support obligations, to the extent that they exceed what would otherwise have occurred, also changes the incentive to separate.

Since child support will generally generate greater separation disincentives for fathers and greater incentives for mothers the net effect is unclear *a priori*. However, child support often interacts with welfare receipts for poor households and, in some cases, child support payments may be tax deductible and hence will interact with the tax system. Thus, it will often be the case that net payments of child support 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 implications of CS for the probability of parents separating and this paper is specifically concerned with the empirical modelling of how child support affects separation.

Section 2 reviews the existing literature. Section 3 explains the theoretical framework and Section 4 outlines the empirical specification. Section 5 presents the UK data and Section 6 focuses on the role of child support in partnership dissolution and explains how contemporaneous child support (and the present value of future) liabilities are constructed. Section 7 presents the results and interpretation while Section 8 analyses the implied separation rates under a child support criterion that is based on the income of the custodial parent compared to a criterion based on the income of both separated parents. Section 9 concludes and evaluates.

2. 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. Since then a number of papers have been stimulated by changes in US

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

welfare rules that followed the 1996 Personal Responsibility and Work Opportunity Reconciliation Act (PRWORA)[2]. An important study that postdates Moffitt's survey is Eissa and Hoynes (1999) which exploits changes in the entitlements for the Earned Income Tax Credit (EITC), the major in-work transfer programme in the US. They show how the expansion in EITC has affected the incentives to have a partner and shows that the phase-in range of EITC encourages partnership and the phase-out discourages it. However, very few papers consider the role of child support explicitly. Hoffman and Duncan (1995) include predicted child support as a regressor in their model of divorce using US Panel Study of Income Dynamics (PSID) data but find that it is statistically insignificant[3]. It is worth noting, however, that the predicted child support was based on the small subsample of 171 divorced women not receiving Aid for Families with Dependent Children (AFDC) in the first two years post-divorce.

In the UK there is very little quantitative research on the economic determinants of separation. Recently Böheim and Ermisch (2001) studied partnership dissolution in the UK using the first eight waves of the British Household Panel Survey (BHPS). Using a discrete-time transition rate model, they estimate the probability of the union dissolving at time *t* as a function of the duration of the partnership as of *t-1* and a vector of economic and partnership characteristics also measured at *t-1*. One major focus of the paper is on the differences between a couple's expectations at *t-2* of their financial situation in the following year and an evaluation of the realised outcome at *t-1*, as predictors of partnership dissolution. It is shown that couples experiencing unexpected improvements in finances have lower dissolution risks while couples experiencing negative shocks are at higher risks: a result which is consistent with the theoretical prediction that income "surprises" affect partnership dissolution.

[2] See Bitler *et al* (2003) which examines the effects of the switch from AFDC to Temporary Assistance to Needy Families (TANF) and of state waivers on flows into and out of marriage. Work on welfare system effects has recently been complemented by Gruber (2003) who exploits the move to unilateral divorce regulations to show a significant increase in the odds of an adult being divorced and of a child living with a divorced parent.

[3] Several papers investigate the role of child support on *remarriage*. Yun (1992) finds a positive effect of the *availability* of child support but a negative effect of actual *payments*. Beller and Graham (1993) and Hu (1994) find no significant effect of child support.

In this paper, we argue that 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* from *t-1*, they would compare their potential net incomes after partnership dissolution with the status quo rather than look at changes in net incomes within the partnership, i.e. we assume that couples are forward looking rather than backward looking. Although the “surprises” might well be one of the factors that determine of 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 (see Hoffman and Duncan (1995) and Weiss and Willis (1997) who use prediction errors from econometric estimates of one period ahead individual incomes).

Child support is the key variable that links the net incomes before and after relationship dissolution for both partners. Indeed, when we abstract from any labour supply or repartnership effects on incomes, child support 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 child support liabilities and receipts. Only two papers directly address this issue: Nixon (1997) uses Current Population Survey data and finds a statistically significant and positive relationship between marital status and child support enforcement, while Helm (2004) uses state-level data and exploits variation in child support enforcement over time and finds no significant effect. Neither paper explore the complex relationship between child support, taxes and transfers which serve to make liabilities and receipts differ.

In this paper we use the prevailing child support rules set by the government to calculate, under plausible assumptions, the estimated child support liability and the implied levels of receipt, for each time period 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. Indeed, because separation is likely to be regarded as permanent, we also compute the present values of child support liabilities and receipts.

Despite its popularity in the media, and to a lesser extent in the psychology and sociology literatures, the “empty nest syndrome” which refers to “feelings of depression, sadness, and/or grief experienced by parents and caretakers after children coming of age leave their childhood homes" (see http://www.psychologytoday.com/), has drawn little attention from economists. The only exception appears to be Heidemann, Suhomlinova and O'Rand (1998), who have found that the onset of the "empty nest" stage increases the risk of marital disruption. However, their sample is households of middle-aged women from the National Longitudinal Survey of Mature Women and their model does not take into account child support variables. This is important because child support liability usually only arises for dependent children (in the UK this is defined as under 17, or 19 if in further education) so that children leaving the nest empty (or, at least, less full) are associated with reductions in CS liability and receipt that changes the incentives.

3. The Theory of Partnership Dissolution and the Econometric Model

The seminal work in this area is Becker (1981) and Becker, Landes and Michael (1977) and there is an excellent review in Ermisch (2003). Their framework has served as the basis for much subsequent research – for example, Peters (1993), Moffitt (1990), Nixon (1997), and Weiss and Willis (1985). In the Becker framework divorce occurs if the combined utility of the partners is higher outside the partnership than inside. So if U is utility (assumed to be transferable), D indicates divorced and M married (we ignore the possibility of cohabitation for the moment) and h and w indicate husband and wife, then divorce requires that $U_{Dw}+U_{Dh} > U_{Mw}+U_{Mh}$. Thus the change in utility should divorce occur is $\Delta U_D = \Delta U_{Dh} + \Delta U_{Dw}$ which can be approximated by $\Delta U_D = -\lambda_h .C_L + \lambda_w .C_R$ where λ is the marginal utility of income, C is child support and L and R indicate liability and receipt, which can be different because of tax and welfare rules. In the absence of mandatory child support we might still expect altruistic parents to make transfers although the data typically suggests that this does not occur. In general, we cannot sign the total utility change but if $C_L = C_R = C$ then $\Delta U_D = (\lambda_w - \lambda_h) .C$ and if wives have lower incomes than husbands following divorce, so that $\lambda_w > \lambda_h$, then we would expect child support to increase the probability of divorce since the transfer would be worth more to the wife than to the husband[4].

[4] We are assuming that prior to separation marginal utilities are equalised across spouses because individual incomes are pooled.

If the wife expects to be on out-of-work welfare in the event of divorce then $C_R = 0$, since the welfare system taxes child support at 100%. We would then expect divorce to be unlikely since $\Delta U_D = -\lambda_h . C_L < 0$, and this would be all the more unlikely the richer is the husband.

Overall, we would expect divorce to be more likely between partners where the husband would have higher post-divorce income than the wife, since then we would expect $(\lambda_w - \lambda_h) > 0$. Moreover, if C_L attracts tax relief, and the tax system is progressive, this would make divorce even more likely.

However, even in these special cases there is a presumption that prior to divorce $\lambda_w = \lambda_h$ because income is assumed to be pooled within intact households and it seems unlikely that divorce would occur in these circumstances, except because of unanticipated shocks. Thus, this framework is quite unlikely to be entirely comprehensive and, even if it were, its empirical implications are only unambiguous in special cases. Nevertheless, the framework is helpful for providing a structure for thinking through these issues.

4. Empirical Specification

The empirical analogue of the theoretical framework assumes that ΔU_D is a latent variable and divorce then occurs, i.e. $D = 1$, if this latent variable is positive and not, i.e. $D = 0$, otherwise. We estimate both a discrete-time transition rate model and hazard models. 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 child support liabilities driven by an important policy reform, to separately identify the effects of children from the effect of child support. The reform replaced *ad hoc* CS arrangements which almost invariably results 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[5].

Moreover, we allow for the potential impact of the departure of all children (the empty nest effect) in our wider sample which also includes childless couples. We

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

use this simple model to home in on a parsimonious specification which we then pursue using a duration modelling framework, which is less restrictive in its distributive assumptions than the simple transition probit.

The discrete-time transition rate model used by Böheim and Ermisch (2001), has the desirable property that probability of survival at time period *t* only depends on survival probability upto period *t-1* and a vector of explanatory variables also measured at *t-1*. 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 β in equation (1), due to the convenient cancelling result.

$$\text{(1)} \qquad \Pr[\, D_{it} = 1 \mid X_{it\text{-}1},\ duration_{it\text{-}1}] = \Phi(\alpha \ln(duration_{it\text{-}1}) + \beta X_{it\text{-}1})$$

where Φ denotes the cumulative standard normal distribution, ln(duration$_{it-1}$) is the log duration of partnership of couple *i* as of *t-1*, and X_{it-1} is a vector of explanatory variables also measured at *t-1*, and α and β are parameters of interest. Note that the assumption of residual homoscedasticity is standard practice in the literature (see Hoffman and Duncan (1995) and Weiss and Willis (1997)).

The hazard function offers a convenient way of defining duration dependence. Positive duration dependence means that the probability that a spell will end shortly increases as the spell increases in length. It specifies the instantaneous rate of failure at $T = t$ conditional upon survival to time *t* as

$$\text{(2)} \qquad \lambda(t) = \lim_{\Delta t \to 0^+} \frac{\Pr(t \le T < t + \Delta t | T \ge t)}{\Delta t} = \frac{f(t)}{S(t)}$$

where *f(t)* and *S(t)* are density and survival functions.

Here we are interested in estimating three of the most popular parametric survival distributions, namely the Exponential, the Weibull and the Lognormal parameterisations which allows for no duration dependence, monotonic and non-monotonic duration dependence respectively. The Generalized Gamma Model is extremely flexible, nesting all three as special cases.

(3)
$$S(t)=\begin{cases}1-I(\gamma,u), & \text{when } \kappa>0\\ 1-\Phi(z), & \text{when } \kappa=0\\ I(\gamma,u), & \text{when } \kappa<0\end{cases}$$
$$f(t)=\begin{cases}\dfrac{\gamma^{\gamma}}{\sigma t\sqrt{\gamma}\Gamma(\gamma)}\exp(z\sqrt{\gamma}-u), & \text{when } \kappa\neq 0\\ \dfrac{1}{\sigma t\sqrt{2\pi}}\exp(-z^{2}/2), & \text{when } \kappa\neq 0\end{cases}$$

where $\gamma=|\kappa|^{-2}$, $z=sign(\kappa)\{\ln)(t)-\mu\}/\sigma$, $\mu=\gamma\exp(|\kappa|z)$, and $\Phi(z)$ is the standard normal cumulative distribution function and *I(a,x)* is the incomplete Gamma function (for details see, for example, Kalbfleisch and Prentice (2002) and Stata Corp (2003)). In other words, the Generalized Gamma distribution reduces to the Weibull distribution when κ=1, to the exponential when κ=1 and σ=1, and to the lognormal case when κ=0.

The one parameter exponential distribution is widely used as a model for duration data. It is simple to work with and to interpret, and is often an adequate model for durations that do not exhibit much variation. The exponential distribution is obtained by taking the hazard function to be a constant, $\lambda(t) = \gamma > 0$, over the range of t. The instantaneous failure rate is independent of t so that the conditional chance of failure in a time interval of specified length, is exactly the same as the unconditional chance of failure. However, in empirical work, the exponential distribution is sometimes found to be less flexible in fitting data than one would like. The two parameter Weibull distribution is an important generalisation of the exponential distribution, which allows for a duration dependence of the hazard on time. The hazard function of the Weibull function is given by

(4)
$$\lambda(t)=\gamma p t^{p-1}$$

where $\gamma > 0$ and $p > 0$. This hazard is monotonically decreasing for $p < 1$, increasing for $p > 1$, and reduces to the constant exponential hazard if $p = 1$. The shape of the hazard function depends critically on the value of p, which is sometimes called the shape parameter. As duration dependence is independent of the parameter γ, γ is sometimes known as the scale parameter. The probability function, the density function and the survival function are, respectively

$$
(5) \qquad \begin{aligned} F(t) &= 1 - \exp(-\gamma t^{p}) \\ f(t) &= \gamma p t^{p-1} \exp(-\gamma t^{p}) \\ S(t) &= \exp(-\gamma t^{p}) \end{aligned}
$$

It is clear that the three parameterisations are general in different ways. The Cox proportional hazard model does not require any specific probability distribution for the survival times. However, when a Weibull or a lognormal distribution is appropriate for the observed survival data, the distinction between which depends on the monotonicity of the duration dependence, it will provide more efficient estimates.

5. Data

This paper uses a sample of couples, drawn from the BHPS, who are at risk of partnership dissolution in the forthcoming year having survived to that time, until they are either censored or the risk has materialised. 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 OSMs after reaching 16, are interviewed each successive year, together with all adult members of their families, even if the OSMs split off from their 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, cohabitation, and fertility and employment transitions, 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 paper includes all women who were either married or cohabiting, and were aged 60 or less, at the time of the second wave. 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 and where the outcome can be either directly observed or imputed with certainty. This leaves us with 15,800 couple-years, of which 319 (just 2.0%) end up in dissolution. For presentation purposes, we choose the woman as the representative for a couple.

Table 1 gives the means and standard deviations of continuous partnership characteristics by partnership outcomes. It 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 to insufficient search. The indication that the probability of a partnership dissolving declines with elapsed partnership duration might reflect either heterogeneity, say in risk aversion, or the hypothesis that couples invest in partnership-specific capital over time. Table 1 also shows that conditional on employment, there is hardly any difference between the net weekly earnings of women who experience a separation and women who remain in partnership. In contrast, women who continue their partnership have partners with higher earnings than those who separate, again conditional on male partners working.

Table 2 reports summary statistics of the indicator variables used in the empirical model. Cohabiting couples are almost five times as likely to separate as legally married couples. This huge difference might reflect the difference in the level of commitment, or it might be due to difference in characteristics between these two groups. Note that for child support purposes, married and cohabiting couples are treated equally. 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 simple correlation and might capture the effect that households with younger children tend to have shorter relationship durations.

Table 1: Means (SD) of Continuous Variables by Partnership Outcome

	Continue	Dissolve
Partnership Characteristics		
Age at start of partnership	23.50 (5.80)	23.06 (5.96)
Log duration of partnership$_{t-1}$	2.76 (0.76)	2.10 (0.97)
Age difference		
Woman's age – partner's age	-2.52 (4.65)	-2.55 (5.57)
Labour Market		
Net Labour income$_{t-1}$	175 (129)	170 (107)
Partner's net labour income$_{t-1}$	346 (263)	314 (189)
N (couple-years)	15481	319

Note: Earnings are in £/week and in January 2004 prices (zero values excluded)

Table 2: Summary Statistics of Indicator Variables

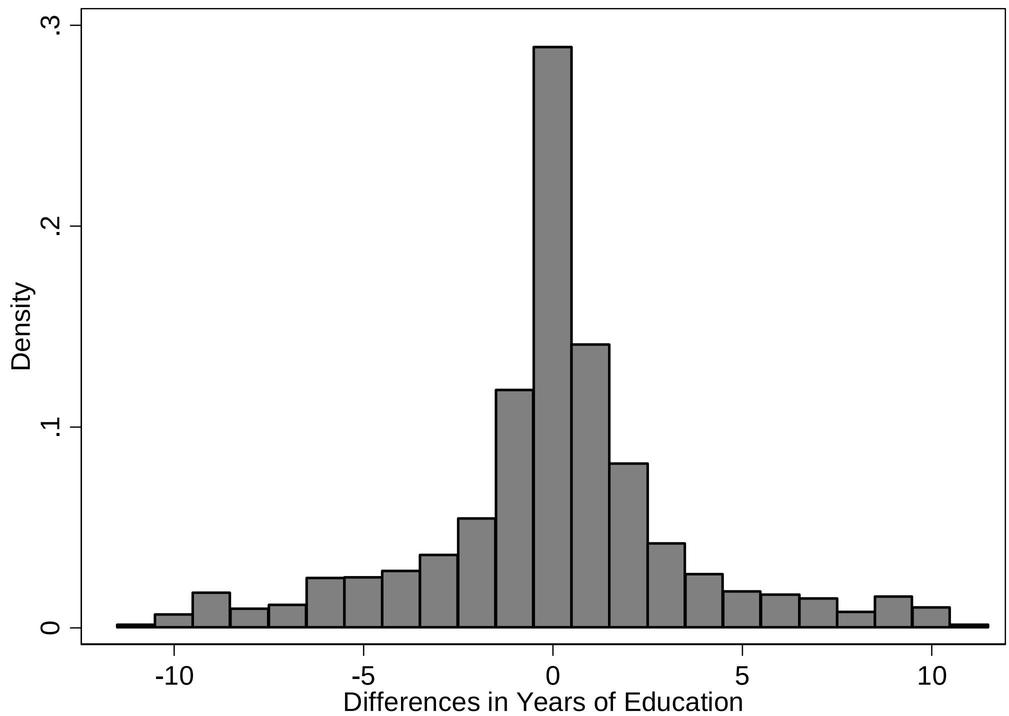
	% dissolving	N
Partnership characteristics		
Married$_{t-1}$ (cohabiting$_{t-1}$)	1.6 (7.6)	14631 (1169)
Partners are different (same) ethnic group	7.2 (1.9)	333 (15467)
Partners have different (same) religion	2.5 (1.6)	7426 (8374)
Youngest child <5(≥5) years at t-1	3.0 (1.8)	3084 (12716)
Education		
Degree	2.0	1426
Other higher	2.0	3570
A-levels	2.4	1348
O-levels	2.4	3841
Basic formal education	1.9	1799
No formal education	1.3	3650
Partner has different (same) education level	2.0 (2.0)	11854 (3946)
Partner's Education		
Degree	1.5	1893
Other higher	1.8	4348
A-levels	2.3	1727
O-levels	2.6	2641
Basic formal education	1.5	1300
No formal education	1.6	3103
Labour market		
Employed$_{t-1}$ (not employed$_{t-1}$)	2.0 (2.1)	10501 (5299)
Partner Employed$_{t-1}$ (not employed$_{t-1}$)	1.9 (2.4)	12287 (3513)
Receipt Income Support between t-2 and t-1	4.2 (1.8)	1156 (14644)
Financial change indicators		
Better financial situation$_{t-1}$	1.9	4110
Same financial situation$_{t-1}$	1.6	7635
Worse financial situation$_{t-1}$	2.6	3901
Partners view financial future differently (similarly)	2.4 (1.6)	6158 (9256)
Surprise indicators (N=15555)		
Large positive surprise	0.0	257
Positive surprise	1.4	2529
No surprise	1.6	7877
Negative surprise	2.3	3377
Large negative surprise	4.2	709
Surprise missing	3.7	806

Note: Sample size for surprise indicators is reduced as 3 consecutives waves are required for this analysis.

The literature of the economics of marriage suggests that education is the main determinant of expected earnings and so it should be a key sorting device in partnership formation. Here, we find some evidence, consistent with the idea of assortative mating, that the difference in number of years of education between partners is important for dissolution. Figure 1 shows that the distribution in the difference id age left full-time education is symmetric with almost 30% of couples having exactly the same number of years of education and over half of all couples' Differences being no more than one year. Our data shows that, of couples with similar number of years of education (i.e. with a difference no more than one year), only 1.7% separate each year compared to 2.3% of couples with larger (either positive or negative) differences.

Employment of either partner is associated with lower risk of partnership dissolution while receipt of Income Support (the UK welfare programme for those with low income – in this case, mostly lone parents with little or no labour income) increases the risk. Interestingly, it is the couples who face the same financial situation as last year, rather than those experiencing improved financial outcomes that are having the lowest risks. As expected, couples experiencing worse outcomes and couples with different views on financial developments face higher risks.

Figure 1: Histogram of the differences in years of education (woman's - man's)



Following Böheim and Ermisch (2001), we also construct "surprise" variables, by comparing people's expectations formed at *t-2* of their financial situation at *t-1* with their evaluation of the actual outcomes at *t-1*, in order to test the hypothesis that new information affects partnership dissolution. Table 3 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. Roughly half of all women correctly predict their financial situations in the following year. Of the remaining half, more women seem to be over-optimistic (i.e. experiencing negative surprises) than to be over-pessimistic (i.e. experiencing positive surprises). More importantly, there is a monotonic increase in the probability of partnership dissolution as we move from large positive to large negative surprises.

Table 3: Expectations and realisations regarding financial situation

Expectation_{t-2}	Evaluation_{t-1}		
	Better off	About the same	Worse off
Better off	= (11.4%)	- (9.4%)	-- (4.8%)
About the same	+ (13.3%)	= (35.9%)	- (13.5%)
Worse off	++ (1.8%)	+ (3.9%)	= (6.0%)

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

6. 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 child support 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, Symons and Walker (1995) and Bingley, Lanot, Symons and Walker (1995) investigate the potential effects of the proposed Child Support reform on net incomes and labour supply of lone mother headed households. More recently, Paull *et al* (2000) investigate the potential effects of the proposed child support reform on net incomes and labour supplies of lone mother headed households.

Despite its apparent importance, there appears to be no research that analyses the impact of potential CS liabilities on partnership dissolution in the literature so far. CS is the key variable that links the net incomes before and after relationship dissolution for both partners. Indeed, when we abstract from any labour supply or repartnership effects, child support 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 changes in CS.

The system of CS that was applied by the Child Support Agency (CSA) in the UK during our sample period of 1992-2001 was based on the principle of income-shares, i.e. child support liability depends primarily on the net incomes of *both* natural parents, subject to deductions which include housing costs, travel-to-work costs and allowances for new children in the second family. Child Poverty Action Group (2000) and Paull, Walker and Zhu (2000) explain how the CS liability is calculated.

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 took place. So we make some naïve, 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 the Parent with Care (PWC)) and stays in the original house;

(3) 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;

(4) Both PWC and NRP’s welfare benefit entitlements are reassessed on separation under the assumptions given by (1)-(3);

(5) Finally, CS liability is calculated under the system of child support described above, based on observed earnings and hours, observed/imputed housing costs for the NRP/PWC and predicted welfare benefit receipts from step (4).

(6) We only include contemporaneous child support liability in the partnership dissolution model, so we are implicitly assuming NRPs are myopic. In principle, we should use instead the present value of the total child support liabilities for each NRP, which also depends on his discount rate and age structure of the

qualifying children (recall that child support payment ceases when a qualifying child reaches 16, or up to 18 if he/she stays on school) and even on the CS liability associated with planned, but yet unborn, children. We reserve this extension for further work.

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

Using the official equivalence scales we then calculate the equalised net incomes pre and post partnership dissolution for the sub-sample of couples with qualifying children (N=8856), which accounts for almost half of the whole sample. Table 4 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 pre and post separation, using before housing cost (BHC) and after housing cost (AHC) scales. We can see couples with dependent children in our sample have a mean weekly total net income of £463 in January 2004 prices, with 22.6% and 66.8% coming from women and men’s labour income respectively, 8.9% from benefits and 1.7% 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 housing costs with a mean of £75 and using the alternative equivalence scale, we get a mean equivalised income of £275 after housing cost.

The PWC and the children will suffer a loss of equivalised income in the magnitude of 25% or 34% on average, depending on whether we use the BHC or AHC measure, despite a 170% increase in total social security transfers and full compliance of child support of the NRPs. Note that PWCs only benefit from less than half of the child support paid by the NRPs, due to the fact that the income support system imposes a 100% tax on all child support receipts. On the other hand, NRPs seem to be better off on both BHC and AHC measures of equivalised income post separation, with a net gain in the magnitude of 30%-45%.

Table 4: Mean equivalised household incomes for PWC (and children) and NRP pre and post separation, by sources of income, BHC and AHC

	Mother with children		Non-resident father	
	Amount	%	Amount	%
Pre-separation:				
Own Net earnings	104.71	22.6	309.18	66.8
Partner's net earning	309.18	66.8	104.71	22.6
Total net benefit	41.09	8.9	41.09	8.9
Other income	7.80	1.7	7.80	1.7
Total net income	462.78	100.0	462.78	100.0
Equivalence scale (BHC)	1.41		1.41	
Equivalised income (BHC)	328.21		328.21	
Equivalence scale (AHC)	1.41		1.41	
Housing cost	75.34		75.34	
Equivalised income (AHC)	274.78		274.78	
Post-separation:				
Own Net earnings	104.71	41.8	309.18	120.2
Partner's net earning	-		-	
Total net benefit	110.41	44.1	12.74	5.0
Other income	2.17	0.9	5.62	2.2
Child support	32.97	13.2	-70.26	-27.3
Total net income	250.26	100.0	257.28	100.0
Equivalence scale (BHC)	1.02		0.61	
Equivalised income (BHC)	245.35		421.77	
Housing cost	75.34		37.73	
Equivalence scale (AHC)	0.96		0.55	
Equivalised income (AHC)	182.21		399.18	

Note: AHC = after housing costs, BHC = before housing costs

7. Estimation Results

We start by re-estimating the Böheim and Ermisch (2001) specification, including their "surprise" variables, as the baseline model[6]. To facilitate comparison, we replicate, in column 1 of Table 5, the results of their main model[7]. Column 2 represents an attempt to replicate their results using all eleven waves of BHPS data now available, instead of just the original eight. Unsurprisingly, the two sets of results including goodness of fit measures, are remarkably similar, with perhaps the exception of the coefficients of labour incomes which are insignificant in any case[8]. In

[6] The alternative specification with the financial change variable in t-1 as a measure of new information yields statistically insignificant coefficients on the financial change dummies.

[7] See Böheim and Ermisch (2001), page 204.

[8] The differences in magnitude presumably reflect the differences in the units of measurement.

the last column we apply the same model specification to the wider sample of couples which also include couples without dependent children. It is worth noting that the fit measures improve significantly as a result of the sample size more than doubling.

This baseline specification include partnership characteristics, age differences between partners, employment and unemployment dummies and the net weekly earnings of each partner, as well as financial "surprises". The estimation results suggest that cohabiting couples are more likely to separate than legally married couples, but the difference is nowhere near that suggested by simple correlation in Table 1. The number of previous marriages also increases the risk. In line with the theoretical predictions, women who started relationship later are less likely to separate while the probability of partnership dissolution also declines with the duration of the relationship. Consistent with the hypothesis of sorting, partners with the same race, religion are less likely to dissolve. Having a non-religious husband does not seem to have an effect. The presence of a pre-school child decreases the risk of partnership dissolving, contrasting with the positive simple correlation. An increase in number of qualifying children in the family increases the risk. Age difference dummies are generally insignificant, except when the woman is at least five years older than the man. Women's earnings significantly reduce the risk of partnership dissolution while her partner's earnings do not make a difference, a result which contradicts Böheim and Ermisch (2001)'s findings, although their estimates are not precisely determined.

Earlier work also suggested that women's economic independence might increase the risk of partnership dissolution. "Surprise" variables do turn out to be significant as a whole, with couples experiencing positive shocks less likely to separate and couples with negative chocks much more likely to dissolve. This result gives strong support to the importance of new information in marital dissolution decisions.

Table 6 presents five model specifications, from the most general which nests the Böheim and Ermisch (2001) model, to a parsimonious model from systematically testing-down. To facilitate model evaluation and selection, we report the change in the probability for an infinitesimal change in each independent continuous variable and the discrete change in the probability for dummy variables, rather than reporting coefficients of the probit model. We also report P-values instead of standard errors.

Table 5: Comparing with the Böheim and Ermisch (2001) results

	Böheim and Ermisch sample Waves 1-8	Böheim and Ermisch sample but 11 waves	Böheim and Ermisch Full sample
Partnership characteristics (at t-1):			
Cohabiting	0.625 (0.171)	0.400 (0.128)	0.356 (0.079)
Number of ex-marriages	0.188 (0.110)	0.073 (0.131)	0.254 (0.083)
Age at start of partnership	-0.043 (0.010)	-0.045 (0.012)	-0.036 (0.007)
Log duration of partnership	-0.307 (0.078)	-0.449 (0.085)	-0.394 (0.037)
Partners same ethnic group	0.293 (0.373)	-0.720 (0.149)	-0.553 (0.116)
Partners have same religion	0.186 (0.090)	-0.073 (0.072)	-0.090 (0.053)
Partners not religious	-0.040 (0.091)	0.029 (0.074)	0.078 (0.054)
Youngest child <5 years	-0.346 (0.110)	-0.173 (0.090)	-0.161 (0.069)
Number of children	0.098 (0.049)	0.106 (0.042)	0.104 (0.026)
Partners different education	0.055 (0.093)	0.051 (0.083)	0.027 (0.060)
Age difference			
Woman 5+ years older	0.385 (0.254)	0.495 (0.203)	0.326 (0.146)
Woman 3-5 years older	0.543 (0.217)	0.092 (0.207)	-0.018 (0.151)
Woman 0-3 years older	0.134 (0.145)	-0.040 (0.127)	-0.010 (0.096)
Partner 2 to 4 years older	0.022 (0.118)	0.005 (0.117)	0.065 (0.087)
Partner 4+ years older	0.180 (0.115)	0.021 (0.125)	-0.027 (0.094)
Labour Market (as of t-1):			
Labour income	0.022 (0.052)	-0.00011 (0.00009)	-0.00013 (0.0001)
Partner's labour income	-0.137 (0.080)	-0.00003 (0.000)	-0.00004 (0.000)
Employed	0.367 (0.316)	0.090 (0.094)	0.084 (0.071)
Unemployed	0.047 (0.101)	0.272 (0.263)	0.109 (0.173)
Partner employed	-0.005 (0.159)	-0.303 (0.130)	-0.097 (0.088)
Partner unemployed	-0.019 (0.144)	-0.179 (0.159)	-0.041 (0.117)
Surprise indicators			
Large positive surprise	a	a	a
Positive surprise	-0.292 (0.148)	-0.149 (0.108)	-0.067 (0.080)
Negative surprise	0.083 (0.098)	0.107 (0.082)	0.100 (0.063)
Large negative surprise	0.218 (0.145)	0.287 (0.123)	0.274 (0.099)
Missing surprise indicator	-	-	0.385 (0.094)
Constant	-0.925 (0.540)	0.843 (0.448)	0.169 (0.268)
N (couple-years)	4451	6837	15262
Chi-square (df)	103.3 (24)	144.8 (24)	342.5 (25)
Pseudo R^2	0.092	0.089	0.111
Log Pseudo-likelihood	-458.4	-694.7	-1311.6
Akaike Information Criterion	**0.2172**	**0.2105**	**0.1753**

Note: Standard errors in parentheses are adjusted to allow for multiple observations per couple. Labour incomes are in £/Month in January 1998 prices.

Böheim and Ermisch (2001) sample: Couples where both partners co-reside before a dissolution and both are interviewed in 3 consecutive waves, the women are aged 60 or less, and at least one dependent child (aged 16 or less) is living in the household.

Full sample: Böheim and Ermisch sample plus childless couples. 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 and where the outcome can be either directly observed or imputed with certainty.

Model 1 represents the full specification nesting the Böheim and Ermisch specification. For this wider sample which includes childless couples, we add 12 more variables to the baseline specification, including both partner's unearned net incomes and working hours, an indicator for having dependent children, as well as two separate measures of the "empty nest effect" to the baseline specification. Most important of all, we include the calculated child support liability, two dummies for the wife's predicted benefit status (in-work and out-of-work benefits respectively) and their interactions with CS.

The wife's unearned income has a positive effect on the risk of partnership dissolution and is marginally significant, while the husband's unearned income has the opposite sign but insignificant. Having any dependent children at all appears to increase the risk. The child support liability has a large negative effect on the hazard of divorce, although the benefit dummies and their interactions with child support liability appear to be insignificant. The post empty nest dummy, which indicates the departure of all children from parental homes in the sample period, is strongly positive. In contrast, the years since empty nest variable is negative and significant, with a magnitude which suggests that that the overall empty nest effect will only remain positive for about six years. Working hours are highly significant, with wife's own hours increasing the risk and husband's working hours reducing the risk. The baseline specification variables still display a similar pattern after the inclusion of new variables, although there is some change in the magnitude of the coefficients. Finally, the goodness-of-fit measures suggest that this full model represents an improvement over the baseline model, despite the apparent over-parameterisation[9].

Model 2 drops all variables in the baseline specification which are individually and jointly insignificant from Model 1. The new omitted surprise category effectively includes people with positive and negative surprises, as well no surprises. Model 3 drops the statistically significant working hours from Model 3, in an attempt to reduce the potential multicollinearity problem of the child support variables. But still, the benefit type dummies and their interactions with child support are not jointly

[9] We assume full compliance throughout our analysis although only about one third pay any CS and only half of those that do pay the full amount. It may be more reasonable to assume that separation depends on the expected liability 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).

significant. Model 4 drops the insignificant benefit type dummies and their interactions with the current CS liability from Model 3 and adds back working hours. It turns out that the retained income support dummies also become insignificant. Model 5 represents the preferred parsimonious specification, after dropping all variables which are not significant at the 5% level, with the exception of current earnings and unearned incomes. This preferred specification only has just over half as many regressors as the previous one, with all but the current income variables significant at the 5% level. This parsimonious specification also represents the best fit among all five specifications according to the AIC.

We apply this specification to a duration model framework, which is less restrictive in its distributional assumptions. The Generalized Gamma Model estimates are presented in the first two columns of Table 7. The Generalized Gamma Model is extremely flexible, nesting as special cases the Weibull, the exponential, and the lognormal. Our Wald tests of the null that κ=1 (the Weibull distribution) has $\chi^2(1)$= 0.44, Prob>χ^2 = 0.5059, while the test of: κ=1 and σ=1 (the Exponential distribution) has $\chi^2(2)$=48.49, Prob>χ^2=0.0000, and the null that κ=0 (the Lognormal distribution) has $\chi^2(1)$=7.95, Prob > χ^2 = 0.0048. Hence the Wald tests overwhelmingly reject the Exponential and Lognormal distributions and are strongly in favour of a Weibull distribution, the results of which are presented in the last two columns of Table 8.

To facilitate easy comparison, both the Gamma and the Weibull model in Table 7 are fitted in the accelerated failure-time metric, in which a positive coefficient implies an increase in the expected time of survival. The two sets of results are remarkably similar, implying that the Weibull is a very good approximation of the more general Gamma model. Indeed, the Weibull model gives a better fit according to the AIC. In contrast to the discrete time model, partner's unearned income now becomes significantly positive. All other coefficients retain their sign and level of significance in the discrete time specification. The shape parameter *p* is very precisely determined with an estimate of 0.5 indicating negative overall duration dependence and a rather sharp decline in the hazard of separation immediately after the formation of a partnership.

Table 6: Probit Model of Partnership Dissolution: changes in probability, P-values in parentheses

	Model (1)	Model (2)	Model (3)	Model (4)	Model (5)
Incomes					
Wife's net earning (£1000/week)	-0.035 (0.002)	-0.037 (0.001)	-0.021 (0.011)	-0.035 (0.001)	-0.035 (0.001)
Partner's net earning (£1000/week)	0.003 (0.275)	0.003 (0.280)	0.001 (0.710)	0.003 (0.323)	0.002 (0.485)
Wife's Unearned Income – (£1000/wk)	0.024 (0.179)	0.026 (0.148)	0.025 (0.148)	0.022 (0.211)	0.020 (0.281)
Partner's Unearned Income –(£1000/wk)	-0.022 (0.183)	-0.024 (0.139)	-0.017 (0.257)	-0.024 (0.137)	-0.025 (0.129)
Child support related variables					
Indicator for qualifying children	0.006 (0.068)	0.007 (0.066)	0.007 (0.072)	0.004 (0.221)	
Current CS liability (£1000/wk)	**-0.110 (0.017)**	**-0.113 (0.016)**	**-0.135 (0.006)**	**-0.067 (0.019)**	**-0.058 (0.025)**
Indicator for wife on IS if divorced	-0.003 (0.370)	-0.003 (0.326)	-0.005 (0.082)	-0.001 (0.670)	
CS*Indicator for wife on IS if divorced	0.040 (0.418)	0.045 (0.371)	0.052 (0.323)		
Indicator for wife on FC if divorced	-0.004 (0.268)	-0.004 (0.270)	-0.004 (0.406)		
CS*Indicator for wife on FC if divorced	0.082 (0.204)	0.081 (0.217)	0.086 (0.229)		
Characteristics					
Empty Nest dummy	**0.036 (0.002)**	**0.039 (0.002)**	**0.040 (0.001)**	**0.039 (0.002)**	**0.036 (0.003)**
Years since empty nest	**-0.006 (0.017)**	**-0.006 (0.015)**	**-0.006 (0.011)**	**-0.006 (0.015)**	**-0.006 (0.016)**
Own working hours/week	0.0002 (0.003)	0.0002 (0.003)		0.0002 (0.003)	0.0003 (0.000)
Partner's working hours/week	-0.0001 (0.004)	-0.0001 (0.002)		-0.0001 (0.002)	-0.0001 (0.002)
Cohabiting	0.015 (0.000)	0.015 (0.000)	0.016 (0.000)	0.015 (0.000)	0.015 (0.000)
Number of ex-marriages	0.007 (0.004)	0.007 (0.005)	0.008 (0.004)	0.007 (0.005)	0.008 (0.004)
Age at start of partnership	-0.001 (0.000)	-0.001 (0.000)	-0.001 (0.000)	-0.001 (0.000)	-0.001 (0.000)
Log duration of partnership	-0.012 (0.000)	-0.012 (0.000)	-0.013 (0.000)	-0.012 (0.000)	-0.011 (0.000)
Partners from same ethnic group	-0.030 (0.000)	-0.031 (0.000)	-0.033 (0.000)	-0.032 (0.000)	-0.033 (0.000)
Partners have same religion	-0.003 (0.100)				
Partners not religious	0.002 (0.179)				
Youngest child <5 years	-0.004 (0.034)	-0.004 (0.030)	-0.005 (0.014)	-0.004 (0.032)	
Number of qualifying children	0.003 (0.005)	0.003 (0.006)	0.004 (0.004)	0.004 (0.004)	0.004 (0.000)
Partners have different education	0.001 (0.733)				
Age difference					
Woman more than 5 years older	0.012 (0.039)	0.012 (0.016)	0.013 (0.013)	0.013 (0.015)	0.013 (0.015)
Woman 3-5 years older	-0.000 (0.914)				
Woman 0-3 years older	-0.000 (0.890)				

Partner 2 to 4 years older	0.002 (0.437)				
Partner more than 4 years older	-0.001 (0.862)				
Labour Market (as of t-1):					
Employed	-0.001 (0.815)				
Unemployed	0.005 (0.449)				
Partner employed	-0.000 (0.948)				
Partner unemployed	-0.002 (0.498)				
Surprise indicators					
Large positive surprise	[a]	[a]	[a]	[a]	[a]
Positive surprise	-0.002 (0.471)				
Negative surprise	0.003 (0.124)				
Large negative surprise	0.012 (0.004)	0.011 (0.005)	0.011 (0.007)	0.011 (0.006)	0.011 (0.006)
Missing surprise indicator	0.015 (0.000)	0.015 (0.000)	0.016 (0.000)	0.015 (0.000)	0.016 (0.000)
N (couple-years)	15262	15262	15262	15262	15262
Chi-square (df)	392.28 (37)	365.15 (24)	351.52 (22)	363.73 (21)	351.17 (18)
Pseudo R^2	0.1235	0.1192	0.1135	0.1185	0.1168
Log Pseudo-likelihood	-1293.5	-1299.9	-1308.3	-1300.9	-1303.5
Akaike Information Criterion	**0.1745**	**0.1736**	**0.1744**	**0.1734**	**0.1733**

Note: 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.

a) All 255 women with large positive surprises continue their partnership in the forthcoming year, hence this variable is dropped from the estimation

b) AIC = 2(-lnL+k)/n where lnL is the log-likelihood, k is the number of parameters and n is the sample size. A lower AIC implies a better fit (Maddala (2004) p488).

Model 1: Full specification nesting the Böheim and Ermisch specification (indicator for planned but not yet born children, education for both partners, and housing costs pre and post separation are statistically insignificant and have already been left out).

Model 2: Dropping religion dummies, different education dummy, insignificant age difference dummies, and employment status dummies and insignificant surprise indicators from Model 1. The new omitted surprise category effectively includes people with positive and negative surprises, as well no surprises.

Model 3: Dropping (statistically significant) working hours from Model 3, in an attempt to reduce collinearity of the CS variables.

Model 4: Dropping statistically significant benefit type dummies and their interactions with the current CS liability from Model 3 and adding back working hours

Model 5: Dropping partner's observed income, indicator for having qualifying children and indicator for children under 5 from Model 4.

Table 7: The Generalized Gamma and the Weibull Models:

	Generalized Gamma Model		**Weibull Model**	
	Coeff.	**P-value**	**Coeff.**	**P-value**
Income				
Wife's net earning (£1000/week)	4.614	0.004	4.338	0.004
Partner's net earning (£1000/week)	-0.167	0.706	-0.166	0.696
Wife's Unearned Income (£1000/week)	-2.667	0.303	-2.576	0.308
Partner's Unearned Income (£1000/week)	5.306	0.042	5.310	0.043
Current CS liability (£1000/week)	8.646	0.022	8.747	0.020
Partnership characteristics				
Empty nest dummy	-2.400	0.018	-2.298	0.016
Years since empty nest	0.887	0.026	0.862	0.023
Own working hours/week	-0.037	0.001	-0.035	0.000
Partner's working hours/week	0.019	0.009	0.018	0.011
Cohabiting	-1.799	0.000	-1.754	0.000
Number of ex-marriages	-0.973	0.008	-0.952	0.009
Age at start of partnership	0.124	0.000	0.124	0.000
Partners from same ethnic group	2.291	0.000	2.205	0.000
Number of qualifying children	-0.602	0.000	0.591	0.000
Woman more than 5 years older	-1.266	0.025	-1.341	0.000
Surprise indicators				
Large positive surprise	25.599	0.000	27.054	0.000
Large negative surprise	-1.074	0.008	-1.051	0.008
Missing surprise indicator	-1.321	0.002	-1.295	0.002
Constant	-0.506	0.620	-0.345	0.734
$\ln\sigma$	0.692	0.000	-	-
κ	0.809	0.005	-	-
P	-	-	0.504	0.000
N (couple-years)	15501		15501	
Chi-square (df)	139.88 (18)		164.29 (18)	
Log Pseudo-likelihood	-672.5		-672.8	
Akaike Information Criterion	0.0895		0.0894	

Figure 2 shows the Weibull survival functions by years of duration of partnership for different levels of CS liability. From the bottom upwards, the five curves indicate survival rates evaluated at zero CS, one standard deviation below the mean, the mean CS level, and one and two standard deviations above the mean CS level respectively, while holding all other regressors at their mean levels. It is clear that the survival rates decline more rapidly in the early years of the partnership and for lower levels of CS.

Figure 3 is an attempt to evaluate the likely impact of CS reform on partnership dissolution. The solid line indicates survival rates evaluated where CS=0, the dotted line shoes the predicted hazard function under the the new CS system which has only been enforced from 2003 (and hence out of our sample period), while the dashed line shows the predicted hazard 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 divorce rate, potentially reducing the divorce probability by around 10% for a 20 year old marriage if all child supports liabilities are fully enforced. On the other hand, the latest reform seems likely to reverse the trend, at least partially, through reducing typical child support liabilities. These results are broadly consistent with our simple simulation results which suggest that the introduction of CS (compared to no CS at all - which was quite typical prior to 1993) has increased the instantaneous hazard by around 14.5% over what we predict it would have been while the new CS reform will decrease the hazard, from the peak level, by about 2%.

Figure 2: Impact of CS liability on Survival Rates

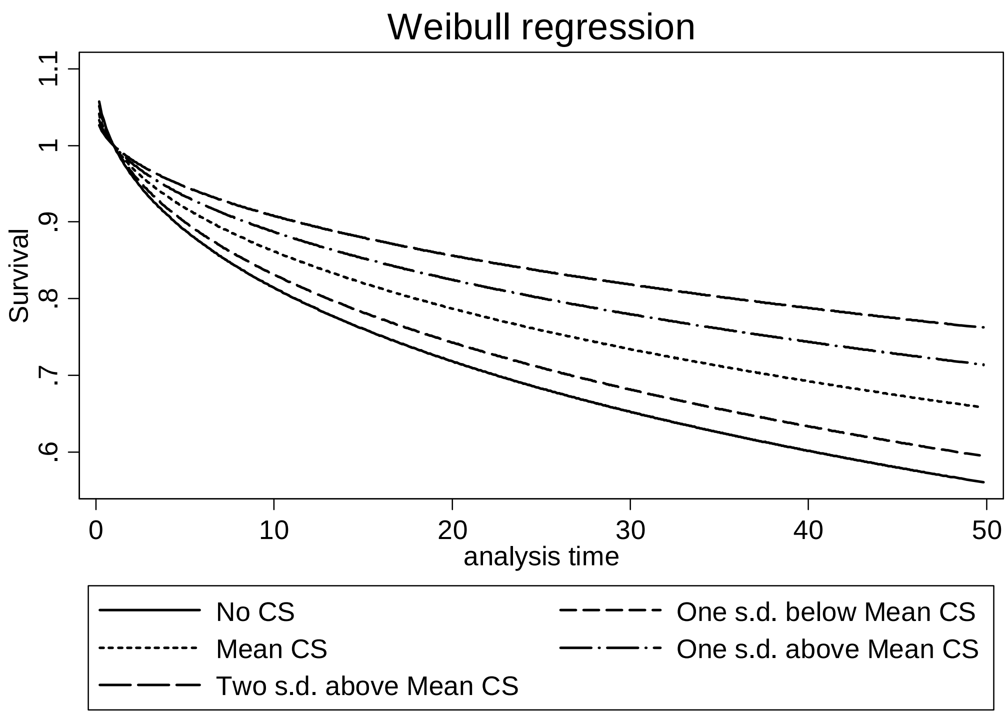
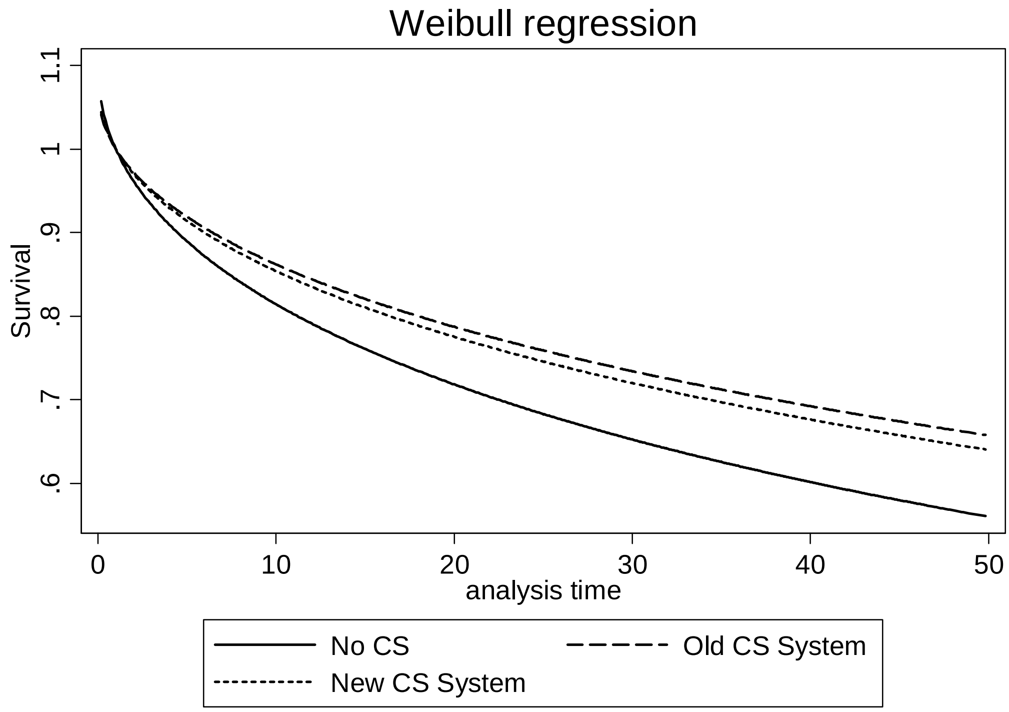


Figure 3: Predicted Impact of CS reform on Survival Rates



8. Simulating CS Design Effects on Separation

We have shown that our econometric results suggest important CS effects on separation. In the UK the CS system has been reformed from a system based on the income of both the households of both separated parents to one where liability is based entirely on the income of the NRP (usually the father). In the USA many states have a system based on NRP income while the others use a weighted sum of both incomes. Here we take a stylised system where the CS liability of the NRP is determined by the weighted average of both parent's net incomes[10], and we then vary the weights while fixing the expected value of total amount of CS liability at some level. That is, we assume the system is given by $CS_i = b\left(a.y_f + (1-a).y_m\right)$ where b is a scale parameter indicating the generosity of the CS system, while a is the parameter that weights the separated parents together[11]. Thus a=0 implies that CS

[10] Of course, in practice CS formulae may be more complicated – as the UK one was. The new UK formula has a=0 but $b(y_f)$ is piecewise linear.

[11] Applying OLS to the sample of BHPS separated couples reveals that, for the UK in the mid to late 1990's, b=0.213 and a=0.812.

liability is independent of the PWC's income while *a*=0.5 implies that the NRP's liability falls by 50% of an increase in PWC's income.

Figure 4 shows the amount of CS contribution from the NRP to the PWC as the weight attached to PWC's income rises. Figure 5 shows the CS contribution from the NRP to the PWC as a percentage of NRP and PWC's actual net earnings. Both figures are drawn for varying weights of the NRP's CS liability (i.e. value of parameter *a*). The two figures suggest that a system which is based entirely on NRP's net earnings would result in a weekly liability of £71.50 per week for the father, which amounts to 22.8% of his actual net earnings. However, if the system was based on the unweighted sum of both parents' earnings, holding the level of total CS liability constant, the NRP's liability would be reduced to £53.1 per week, or 16.9% of their respective net earnings, with the PWC (notionally) contributing an equal share of (typically) her net earnings to make up the balance.

Figure 6 shows the predicted effects of parameter *a* on the survival rate evaluated at the mean liability. For example, the probability of surviving to 10 year is approximately 6% higher if *a*=0 compared to *a*=0.5. This corresponds to an instantaneous separation rates of 1.60% if *a*=0 compared to 1.73% for a system that was based on the unweighted sum of both parents' incomes, holding the level of CS liability constant.

Figure 4 NRP's CS Liability by Weight on PWC Income

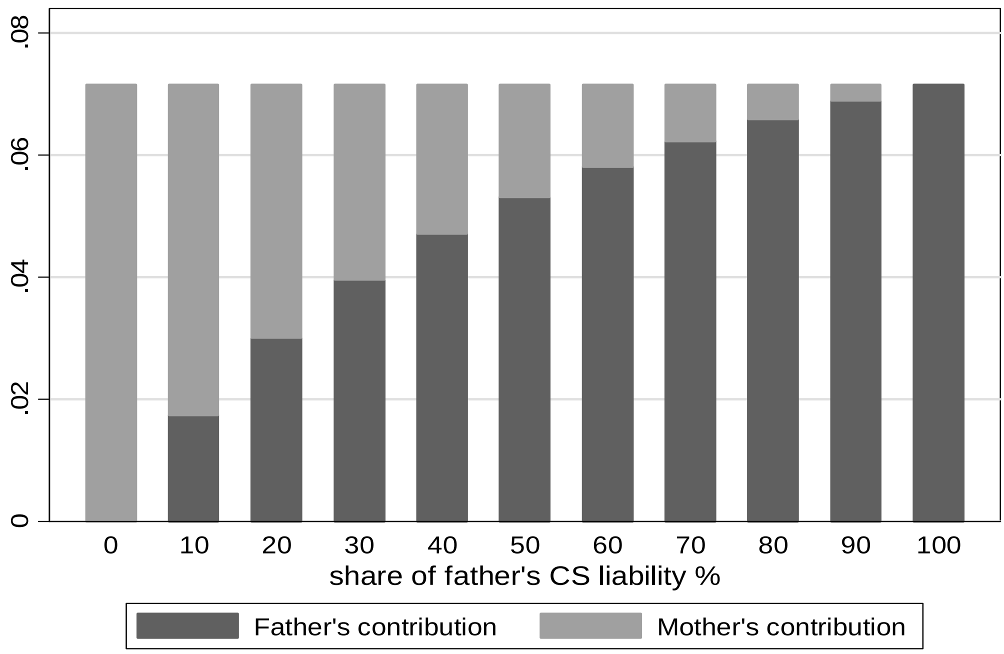


Figure 5 NRP's CS as Share of Net Income

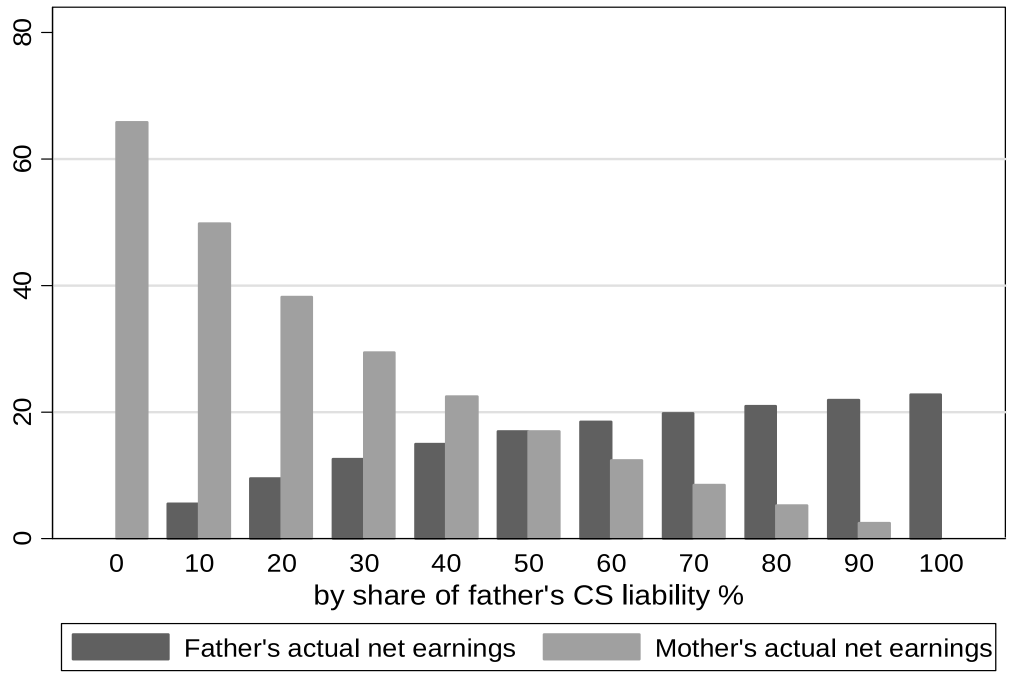
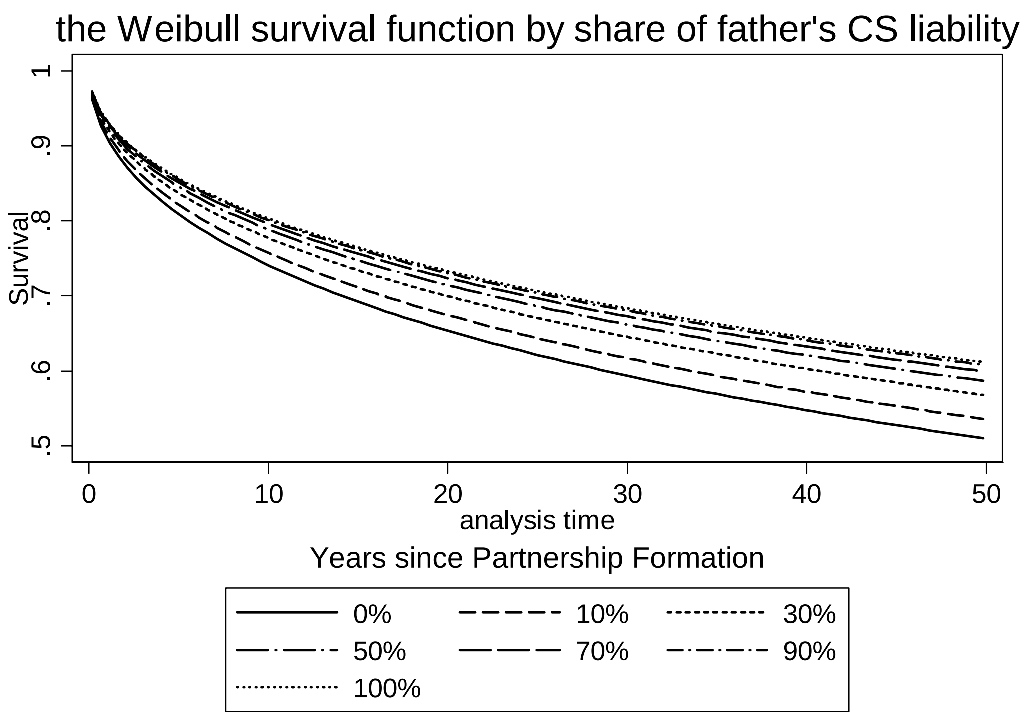


Figure 6 Effects of Income Sharing Rule on Predicted Survival Rates



9. Conclusions

This paper studies the determinants of partnership dissolution in the UK using the British Household Panel Study (BHPS). 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 have a substantial impact on the probability of partnership dissolution. Moreover, we exploit the variation in child support liabilities, driven by an important policy reform, to separately identify the effects of children in the household from the effect of child support liability. We find there is very strong evidence that an increase in the implied child support liabilities significantly reduces the dissolution risk while an increase in the wife's current earnings has the opposite effect. Moreover, we find the departure of all children (an empty nest) has a large positive effect.

We use the estimates to simulate the effect of CS on separation rates. We calculate that divorce rate would have been 14.5% higher (i.e. 2.16% instead of 1.85%) were it not for the introduction of a CS formula in the UK, and that the very recent reform which has reduced typical liabilities may well modestly increase the separation rate from 1.85% to 1.89%.

We also use the estimates to simulate the effect of alternative CS designs – we find that a system which is based entirely on the non-resident parent's income would result in a separation rate of 1.60% compared to 1.73% for a system that was based on the unweighted sum of both parents' incomes, holding the level of CS liability constant.

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[12]. The assumptions of no labour supply or repartnership effects are maintained hypotheses could also be tested[13]. But despite our reservations about these assumptions we believe these existing findings do have significant policy

[12] 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.

[13] The Appendix shows how the working and repartnership behaviour of the partners varies up to and beyond separation. 20% repartner shortly after divorce while there seems to be little change in labour supply behaviour.

implications. For instance, our results suggest that the current child support reform (Department of Social Security (1999)), and the CS pass-through that has been a feature of CS design in some US states, might have effects on divorce rates through changing child support liabilities and receipts that are largely unintended.

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 together a partnership that would otherwise dissolve, welfare of all parties has improved. There is little research on the impact of separation on well-being and further research needs to be done to separate out the effects of separation from its financial consequences, especially on outcomes for children, including their well-being.

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Appendix

Table A: *Equivalised Income, Poverty Rates, Labour Market Participation and Repartnership*

	Equivalised Income		Poverty Rates (%)		Labour Market Participation (%)		Repartnership (%)	
Years Separated	Fathers	Mothers	Fathers	Mothers	Fathers	Mothers	Fathers	Mothers
<=-4	139.6	139.6	43.7	43.7	83.6	56.8		
-3	136.7	136.7	43.0	43.0	88.6	62.0		
-2	130.1	130.1	40.2	40.2	87.0	65.2		
-1	141.9	141.9	34.2	34.2	86.7	58.3		
0	142.8	142.8	31.3	31.3	84.7	58.0		
1	171.0	98.3	34.6	59.1	85.0	56.7	22.8	20.5
2	187.1	116.6	28.8	49.0	81.7	57.7	29.8	21.2
3	191.3	131.9	32.6	47.2	86.5	53.9	29.2	21.3
4	209.1	136.7	21.1	44.7	82.9	48.7	38.2	32.9
>=5	229.1	166.7	19.9	24.0	91.1	62.3	55.5	35.6
Total	167.1	135.1	33.2	41.0	85.8	58.2		