Child Support 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 and entitlements driven, in part, by the introduction of a new set of complex rules that determined child support liability, and by their interaction with welfare rules. Our sample has the virtue that the post 1992 variation in child support liability for all couples in partnerships formed prior to 1992 is unanticipated. We find strong evidence that the resulting large child support liabilities significantly *reduced* dissolution risks. Simulations based on our results suggest that child support criteria that are driven only by the non-custodial parent's income, compared to criteria driven by the 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 the maximization of marital incomes by men and women would induce strong segregation in the marriage market in the form of (positive) assortative mating and so separation largely results from uncertainty or unfavourable outcomes.

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 the motivation provided by the growth in divorce that has occurred over time and across many countries, and the widespread acceptance that divorce has strong adverse affects on children[1], there are surprisingly few empirical studies of the determinants of partnership dissolution. Moreover, despite the policy reforms that have occurred to increase child support entitlements and compliance, little of the research that has been done has considered the impact of child support (CS). CS will generally generate 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 paper is specifically concerned with the empirical modelling of how CS affects separation. The main contribution of this paper is to quantify the net effect of CS on separation incentives and evaluate the implications that CS system design might have for separation rates.

In recent years, partly because of the dramatic growth of separation amongst parents, CS has become a major policy issue. High rates of lone parenthood and low

[1] See, for example, Cherlin *et al* (1995).

levels of child support have resulted in growing numbers of lone parents, almost all mothers, many of whom 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[2] for cases who were entitled to welfare payments[3]. However, the levels of child support liabilities were often extremely high and accumulated arrears sometimes 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 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[4].

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[5]. 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.

[2] 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.

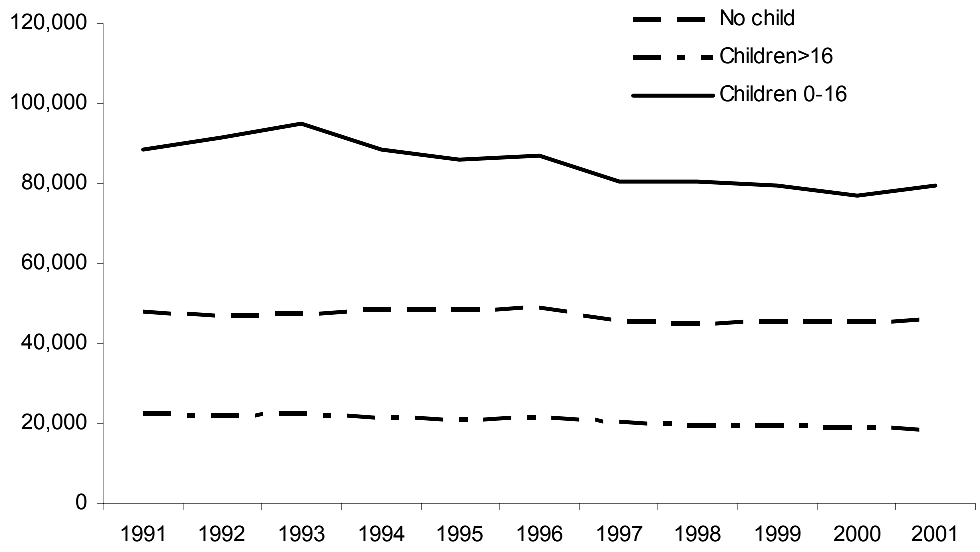
[3] Cases not on welfare could be considered for a CSA assessment at the request of the parent with care. Thus, the CSA rules act as a focal point for other negotiated transfer and/or settlement payments for non-welfare cases.

[4] 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.

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

Aggregate data in Figure 1 shows that the number of divorces in families with children under 16 fell by 15% from 1993 to 2001 relative to those with no children, and by almost this degree relative to those whose youngest child was 16+[6]. This would be consistent with greater CS liabilities reducing separation incentives.

Figure 1 Divorces, 1991-2001, England and Wales



Source: Office of National Statistics

The methodology we use is to estimate the determinants of the hazard of separation using a panel of couples who we can follow post separation. 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 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.

[6] CS liability ceases in the UK when the child is 16 or 18 if in full-time education.

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 welfare rules that followed the 1996 Personal Responsibility and Work Opportunity Reconciliation Act which reformed the US welfare system[7]. An important study that postdates Moffitt's survey is Eissa and Hoynes (2001) 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 separation using US Panel Study of Income Dynamics (PSID) data but find that it is statistically insignificant[8]. It is worth noting, however, that the predicted child support was based on the small subsample of 171 separated women not receiving Aid for Families with Dependent Children (AFDC) in the first two years post-separation.

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

[7] 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.

[8] 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. Carlson *et al.* (2003) investigates the effects of CS on partnership formation in a sample of unmarried mothers and finds no effects on marriage.

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.

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* from *t-1*, 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[9].

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, 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. However, neither of these papers explore the complex relationship between child support, taxes and transfers which serve to make liabilities and receipts differ.

In this paper we use the UK CS rules prevailing from 1993 to 2003 to calculate, under plausible assumptions, the estimated child support liability and the implied levels of receipt, for each time period that a couple is at risk of dissolution.

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

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 experiment with including the present values of child support liabilities and receipts[10].

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 separation occurs if the combined utility of the partners is higher outside the partnership than inside. So if U is utility (assumed to be transferable between partners), D indicates separated and M married (we ignore the possibility of cohabitation for the moment) and h and w indicate husband and wife, then separation requires that $U_{Dw}+U_{Dh} > U_{Mw}+U_{Mh}$. Thus the change in utility should separation 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 the interaction between CS and the 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 is not quantitatively significant. In general, we cannot sign the total utility change but if there are no tax or welfare reasons for receipts to differ from liabilities then $C_L = C_R = C$ and $\Delta U_D = (\lambda_w - \lambda_h) . C$. So if wives have lower incomes than husbands following separation, so that $\lambda_w > \lambda_h$, then we would expect child support to increase the

[10] 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 upto 16, or 18 if in full-time 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. Thus, while we are not especially concerned about the hypothesis we do control for empty nests in our analysis to ensure that any effect does not contaminate our estimates of CS effects.

probability of separation since the transfer would be worth more to the wife than to the husband.

If the wife expects to be on out-of-work welfare in the event of separation then $C_R = 0$, if the welfare system taxes child support at 100%. We would then expect separation 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 separation to be more likely between partners where the husband would have higher post-separation 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 separation even more likely.

However, even in these special cases there is a presumption that prior to separation $\lambda_w = \lambda_h$ because income is assumed to be pooled within intact households. Of course, if this were really true then it seems unlikely that separation would occur except because of unanticipated shocks. Thus, this simple theoretical framework is quite unlikely to be entirely applicable 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 separation 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 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[11]. Receipts differed from liabilities because of complex interactions with the tax and welfare systems.

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 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 the 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 the determinants of the explanatory variables on the hazard.

In addition to modelling a simple transition model we also estimate three of the most popular parametric survival distributions, namely the Exponential, the Weibull and the Lognormal parameterisations which, respectively, allow for no duration dependence, monotonic and non-monotonic duration dependence. 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[12]. In addition we estimate the

[11] See Paull *et al* (2000) and Appendix A for details.

[12] The hazard function of the Weibull function is given by $\lambda(t) = \gamma p t^{p-1}$ 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.

Generalized Gamma Model which is extremely flexible, nesting all three as special cases. In particular, we estimate

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

where $\gamma=|\kappa|^{-2}$, $z=sign(\kappa)\{\ln)(t)-\mu\}/\sigma$, $\mu=\gamma\exp(|\kappa|z)$, $\Phi(z)$ 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 if κ=1, to the exponential if κ=1 and σ=1, and to the lognormal if κ=0.

5. Data

This paper uses a sample of couples, drawn from the BHPS wave 1 and later, who are at risk of partnership dissolution in the forthcoming year having survived to that time, until they are either censored (at the latest wave) or the risk has materialised (i.e. they have experienced their first separation). 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 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 couples (married or cohabiting), who have had their children, at the time of the first wave and adds all future couples with children up to wave 12, the latest available for analysis. In this paper we deliberately exclude couples who have not started families to minimize the problem of having to deal with the endogenous fertility decision[13]. However, our sample does include couples whose children have either become too old to qualify for child support or who have had children that have now left the parental homes. We do this to facilitate comparison with other studies which have largely focussed on the non-economic determinants of separation. For example, this literature has found a well-determined "empty nest" effect where it is the exit of children from the household that contributes to the separation hazard[14]. However, our focus is on economic variables and, in particular, child support.

For people experiencing multiple relationship dissolutions over the sample period, we only focus on the first relationship[15]. 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 19,442 couple-years, of which 389 (just 2.00%) end up in dissolution. To test the robustness of our results we also conduct analyses over just the couples who have children since these are the main focus of our interest. For presentation purposes, we choose the woman as the representative for a couple.

Table 1 gives the means and standard deviations of continuous variables relating to partnership characteristics broken down by partnership outcomes. The table 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

[13] See Aizer and McLanahan (2004) and Case (1998) for research on the impact of CS on fertility. They find no statistically significant effects.

[14] None of this empty nest research, to our knowledge, allows for the potential endogeneity of leaving home.

[15] We do not model second partnership formation (or even first) and, in any event, the number of multiple separations is too small to allow analysis of multiple partnership spells.

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 over six 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 despite the fact that they have all had children. 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 simply reflect the effect that households with younger children tend to have shorter relationship durations.

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

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

	Continue	Dissolve
Partnership Characteristics		
Age at start of partnership	23.60 (6.14)	23.10 (5.67)
Log duration of partnership$_{t-1}$	2.94 (0.77)	2.21 (1.00)
Age difference		
Woman's age – partner's age	-2.48 (4.40)	-2.52 (5.18)
Labour Market		
Net Labour income$_{t-1}$	157 (132)	155 (112)
Partner's net labour income$_{t-1}$	349 (264)	321 (187)
N (couple-years)	19053	389

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

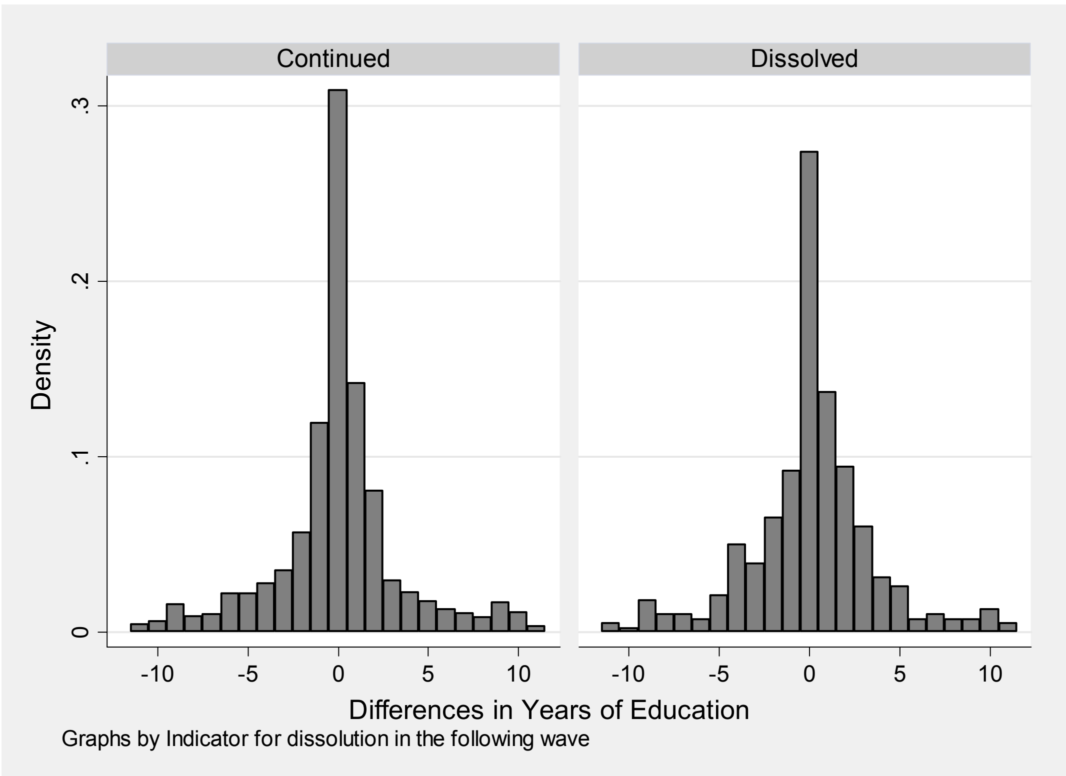
Table 2: Summary Statistics of Indicator Variables

	% dissolving	N
Partnership characteristics		
Married$_{t-1}$ (cohabiting$_{t-1}$)	1.6 (10.0)	18458 (984)
Partners are different (same) ethnic group	3.8 (1.9)	1359 (18083)
Partners have different (same) religion	2.6 (1.5)	9176(10266)
Youngest child <5(≥5) years at t-1	4.0 (1.5)	4151(15291)
Education		
Degree	1.3	1264
Other higher	2.2	3739
A-levels	3.0	1347
O-levels	2.5	4291
Basic formal education	2.4	2340
No formal education	1.2	5976
Partner has different (same) education level	2.2 (1.5)	14540 (4902)
Partner's Education		
Degree	1.7	1743
Other higher	1.9	4847
A-levels	2.0	1793
O-levels	2.0	2874
Basic formal education	1.6	1618
No formal education	1.5	4458
Labour market		
Employed$_{t-1}$ (unemployed$_{t-1}$)	2.0 (5.0)	9803(220)
Partner Employed$_{t-1}$ (unemployed$_{t-1}$)	2.1 (3.8)	11490 (678)
Receipt Income Support between t-2 and t-1	4.6 (1.8)	1409 (18033)
Financial change indicators		
Better financial situation$_{t-1}$	2.5	4242
Same financial situation$_{t-1}$	1.5	9792
Worse financial situation$_{t-1}$	2.5	4947
Partners view financial future differently (similarly)	2.3 (1.6)	7761 (11304)
Surprise indicators (N=16547)		
Large positive surprise	0.9	228
Positive surprise	1.6	2556
No surprise	1.4	8736
Negative surprise	2.2	3390
Large negative surprise	4.5	646
Surprise missing	2.3	991

partners is important for dissolution. Figure 2 shows that the distribution in the difference in 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 of no more than one year), only 1.7% separate each year compared to 2.4% of couples with larger (either positive or negative) differences. Similarly, Table 2 shows that couples with the same level of highest education qualification have a separation risk of 1.5% per year compared to 2.2% for couples with different levels of qualifications.

Unemployment is associated with higher risk of partnership dissolution for both men and women (the omitted category being inactivity). 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) more than doubles the risk. Interestingly, it is the couples who face the same financial situation as last year, rather than those experiencing improved financial outcomes that have the lowest risks. As expected, couples experiencing worse outcomes and couples with different views on financial developments face higher risks.

Figure 2: Histogram of the differences in years of education (woman's - man's)



Following Böheim and Ermisch (2001), we 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: Dissolution Rates by Expectations and realisations regarding financial situation (N=15525)

Expectation$_{t-2}$	Evaluation$_{t-1}$		
	Better off	About the same	Worse off
Better off	2.7% = (9.2%)	2.6% - (8.4%)	4.5% -- (4.2%)
About the same	1.8% + (12.0%)	1.2% = (39.8%)	1.9% - (13.6%)
Worse off	0.9% ++ (1.5%)	1.1% + (4.7%)	1.2% = (6.7%)

Note: ++ large positive surprise, + positive surprise, = as expected, - negative surprise, -- large negative surprise. Numbers in parentheses are relative frequencies. Sample size for surprise indicators is reduced as 3 consecutives waves are required for this analysis.

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 this Child Support system on net incomes and labour supply of lone mother headed households.[16]

[16] More recently, Paull *et al* (2000) investigate the potential effects of a reform to the child support system, which was not implemented until 2003, on net incomes and labour supplies of lone mother headed households.

The contrast between the UK and the US is interesting. US states have always been able to design their own specific child support mechanisms and states have divided into two broad camps. In the *income-shares* camp, child support is a proportion of the combined incomes of both natural parents. The UK system analysed here broadly falls into this category, with the liability of the non-resident parent prorated between the parents according to each share of their combined incomes. In contrast, the newly reformed UK system falls into the *percent-of-income* camp where child support is a percentage of the non-resident parent's income with the percentage varying with the number of children. That is, in the 1993-2003 CS system liability depends primarily on the net income of both natural parents. Exemptions from this income include allowances for new children, which may be partially offset if the new partner has sufficiently high income. For the non-resident parent, the presence of stepchildren and the income of a new partner also affect the maximum and minimum levels of liability[17].

The way in which Child Support interacts with the tax and welfare system is also important. A major part of the recent reform dealt with the benefit disregards for receipt of Child Support and proposed the introduction of a £10 disregard for Income Support and also proposed increasing the current Family Credit (FC) disregard of £15 such that Working Families Tax Credit (WFTC) which replaced FC in 1999 would disregard all child support payments no matter how large.

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 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 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[18];

[17] See the Appendix A for details.

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

(3) Both PWC and NRP's welfare benefit entitlements are reassessed on separation under the assumptions given by (1)-(2);

(4) 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 (3).

(5) In principle, we should use the present value of the expected 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 planned, but yet unborn, children. In practice, we found that only current liabilities were empirically important and reserve this for more detailed 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=9900), which accounts for almost half of the whole sample. Table 4a 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. Couples with dependent children in our sample have a mean weekly total net income of £440 in January 2004 prices, with 22.8% and 66.4% coming from women and men's labour income respectively, 9.1% from benefits and 1.8% from all other income. With a mean equivalence scale of 1.40, this results in an equivalised income of £322 for the family before housing cost. After deducting the housing costs with a mean of £76 and using the alternative equivalence scale, we get a mean equivalised income of £266 after housing cost. The PWC and the children will suffer a loss of equivalised income in the magnitude of 22% or 30% on average, depending on whether we use the BHC or AHC measure, despite a 180% increase in total social security transfers and full compliance of child support of the NRPs. Note that, on average, PWCs only benefit from less than half of the child support paid by the NRPs, due to the fact that the Income Support (IS) system, that provides income for those out-of-work, imposes a 100% tax on child support receipts in excess of the level of IS entitlement. 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 25%-41%.

Table 4a *Mean equivalised household incomes for PWC (and children) and NRP pre and post separation, by sources of income, (N=9900)*

	Mother with children		Non-resident father	
	Amount	%	Amount	%
Pre-separation:				
Own Net earnings	100.18	22.8	291.78	66.4
Partner's net earning	291.78	66.4	100.18	22.8
Total net benefit	39.85	9.1	39.85	9.1
Other income	7.93	1.8	7.93	1.8
Total net income	439.75	100.0	439.75	100.0
Equivalence scale (BHC)	1.40		1.40	
Equivalised income (BHC)	321.67		321.67	
Equivalence scale (AHC)	1.41		1.41	
Housing cost	76.35		76.35	
Equivalised income (AHC)	265.58		265.58	
Post-separation:				
Own Net earnings	100.18	40.8	291.78	119.2
Partner's net earning	-		-	
Total net benefit	111.30	45.3	13.92	5.7
Other income	2.29	0.9	5.64	2.3
Child support	31.99	13.0	-66.58	-27.2
Total net income	245.75	100.0	244.77	100.0
Equivalence scale (BHC)	1.01		0.61	
Equivalised income (BHC)	252.46		401.26	
Housing cost	76.35		39.12	
Equivalence scale (AHC)	0.96		0.55	
Equivalised income (AHC)	185.07		373.91	

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

Table 4b shows the expected CS liabilities faced by the NRP and the expected CS receipts net of social security benefits for the PWC, broken down by the predicted benefit status of the mother upon separation. The proportion of mothers expected to be on out-of-work benefits, are quite high, because we assume no change in employment status following separation in our model.

Table 4c is an attempt to assess the goodness of fit of our CS entitlement prediction, by exploiting the subsample of couples who are observed both before and after separation. To minimize the problems of changes in employment and repartnerships (and having children in the second families which will affect CS liabilies) and to a lesser extent attrition, we restrict our sample to the first two waves immediately after separation. It is clear that our CS routine has done a reasonably good job of predicting CS entitlement. The estimated compliance rate is just under 60%, in line with official statistics. For the subgroup of NRPs who do pay any CS, the difference between our predicted entitlement and the actual payment is less than £1 per week on average. It is interesting to note that of the subsample of NRPs 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 PWCs, and it is also possible that some PWCs only perceive CS receipts net of social security benefits as real CS.

Table 4b *Child Support Entitlements and Receipts by Predicted Benefit Status of the Caring Mother upon Separation (N=9900)*

	Not on Benefit	On Out-of-Work Benefit	On In-Work Benefit	Total
CS Entitlement (£/wk)	70.83	64.83	64.92	66.58
CS Receipt net of benefits(£/wk)	70.83	6.51	32.86	31.98
N	2847	4432	2621	9900
%	28.8	44.8	26.5	100.0

Table 4c *Comparing Predicted Child Support Entitlement with Actual Payments/Receipts using the matched couple sample (N=184)*

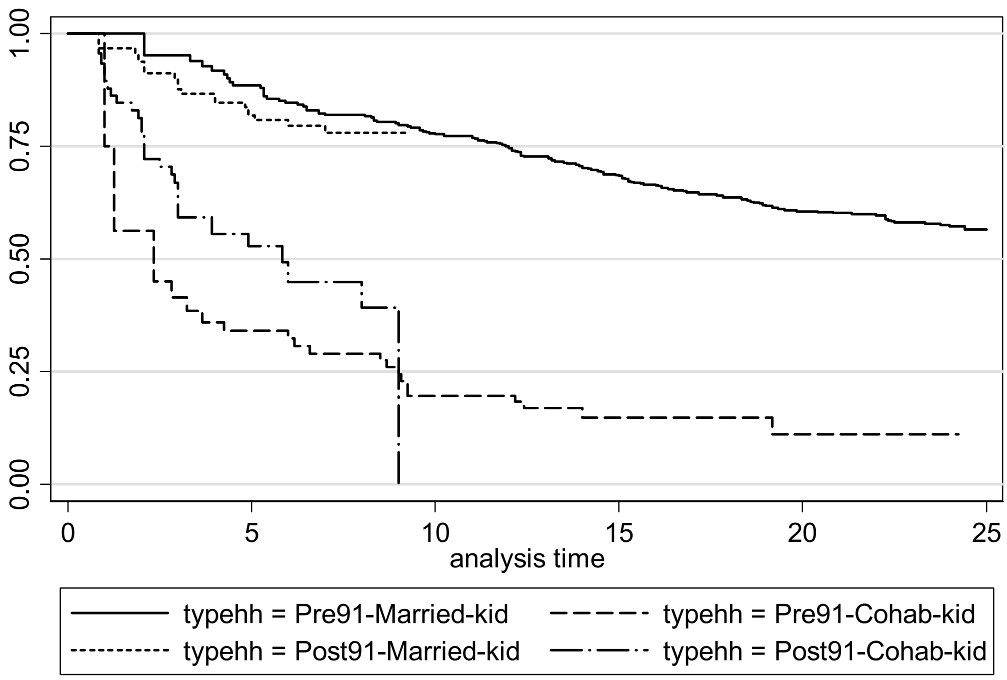
	NRP Paying	NRP Not Paying
Predicted CS entitlement (£/week Jan 2004 prices)	97.26	71.51
Actual Payment (£/week Jan 2004 prices)	96.46	-
% of PWC report receiving CS	56.6	16.7
N	106	78
%	57.6	42.4

Note: sample based on all matched couples entitled to positive child support in the first two waves immediately after separation.

7. Estimation Results

Figure 3 shows Kaplan-Meier plots of the separation hazards for different types of households with dependent children up to 25 years since the formation of the current relationship[19]. For post 91 partnerships the CS is no longer a surprise and is presumably taken into account in the original partnership decision. So that post 91 partnerships should have higher survival rates - partnerships of marginal robustness will no longer be formed. The finding that post 91 cohabitants have a higher survival rate than that of pre 91 cohabitants is consistent with this theoretical prediction, although there seems to be little difference between married couples pre and post reform.

Figure 3: Kaplan-Meier Survival Functions by Household Types



We start by re-estimating the Böheim and Ermisch (2001) specification, including their "surprise" variables, as the baseline model[20]. To facilitate comparison, we

[19] We don't include households without qualifying children here because there are relatively few observations within this range. For instance, pre 1991 married couples without dependent children have a mean duration of 33 years.

[20] 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.

replicate, in column 1 of Table 5, the results of their main model[21]. Column 2 represents an attempt to replicate their results using all twelve 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. In 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 sample size is more than doubled. 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. Consistent with Böheim and Ermisch (2001)’s findings, women’s earnings do not make a difference while an increase in their partners’ earnings significantly reduces the risk of partnership dissolution. “Surprise” variables do turn out to be significant as a whole, with couples experiencing positive shocks being less likely to separate and couples with negative shocks much more likely. This result gives strong support to the importance of new information in marital dissolution decisions.

Table 6 presents three model specifications from the most general, which nests the Böheim and Ermisch (2001) model, to a parsimonious model by systematically testing-down. To facilitate model evaluation and selection, we report the change in the probability for an infinitesimal change in each independent continuous

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

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

	Böheim and Ermisch sample Waves 1-8	Böheim and Ermisch sample but 12 waves	Böheim and Ermisch including childless couples
Partnership characteristics (at t-1):			
Cohabiting	0.625 (0.171)	0.546 (0.109)	0.545 (0.096)
Number of ex-marriages	0.188 (0.110)	0.083 (0.065)	0.184 (0.046)
Age at start of partnership	-0.043 (0.010)	-0.056 (0.009)	-0.043 (0.007)
Log duration of partnership	-0.307 (0.078)	-0.344 (0.061)	-0.381 (0.048)
Partners same ethnic group	0.293 (0.373)	-0.257 (0.127)	-0.150 (0.096)
Partners have same religion	0.186 (0.090)	-0.087 (0.067)	-0.136 (0.056)
Partners not religious	-0.040 (0.091)	-0.018 (0.069)	0.039 (0.058)
Youngest child <5 years	-0.346 (0.110)	-0.134 (0.083)	-0.191 (0.077)
Number of children	0.098 (0.049)	0.087 (0.041)	0.101 (0.027)
Partners different education	0.055 (0.093)	0.045 (0.077)	0.081 (0.063)
Age difference			
Woman 5+ years older	0.385 (0.254)	0.536 (0.199)	0.479 (0.151)
Woman 3-5 years older	0.543 (0.217)	0.127 (0.183)	0.052 (0.142)
Woman 0-3 years older	0.134 (0.145)	-0.106 (0.111)	-0.032 (0.092)
Partner 2 to 4 years older	0.022 (0.118)	-0.115 (0.102)	-0.054 (0.084)
Partner 4+ years older	0.180 (0.115)	-0.048 (0.107)	-0.063 (0.089)
Labour Market (as of t-1):			
Labour income	0.022 (0.052)	-0.005 (0.069)	-0.027 (0.067)
Partner's labour income	-0.137 (0.080)	-0.087 (0.057)	-0.135 (0.05)
Employed	0.367 (0.316)	0.039 (0.082)	0.035 (0.068)
Unemployed	0.047 (0.101)	0.086 (0.261)	0.366 (0.174)
Partner employed	-0.005 (0.159)	-0.041 (0.125)	0.084 (0.090)
Partner unemployed	-0.019 (0.144)	-0.125 (0.164)	-0.090 (0.129)
Surprise indicators			
Large positive surprise	-	-0.561 (0.369)	-0.268 (0.294)
Positive surprise	-0.292 (0.148)	-0.086 (0.096)	-0.041 (0.079)
Negative surprise	0.083 (0.098)	0.141 (0.077)	0.103 (0.065)
Large negative surprise	0.218 (0.145)	0.280 (0.115)	0.274 (0.104)
Missing surprise indicator	-	-	0.169 (0.102)
Constant	-0.925 (0.540)	0.333 (0.360)	-0.081 (0.292)
N (couple-years)	4451	7195	16033
Chi-square (df)	103.3 (24)	169.9(25)	374.6 (26)
Pseudo R^2	0.092	0.093	0.125
Log Pseudo-likelihood	-458.4	-841.5	-1266.7
Akaike Information Criterion	**0.2172**	**0.2411**	**0.1614**

Note: Standard errors in parentheses are adjusted to allow for multiple observations per couple. Labour incomes are in £1000/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: All couples who have had children, i.e. Böheim and Ermisch sample plus couples whose children have grown up or even have left parental homes. 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.

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.

Model 1 represents the full specification nesting the Böheim and Ermisch specification, with the exception of earnings which are dropped to minimize the multicollinearity problem. For this wider sample which includes couples without qualifying children, we add 14 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, together with dummies for wife's predicted benefit status (in-work and out-of-work benefits respectively) and post-91 sample and their interactions with CS[22].

The wife's unearned income has a positive effect on the risk of partnership dissolution and is significant, while the husband's unearned income has the opposite sign but insignificant. Having any dependent children at all appears to have no effect. The child support liability has a large negative effect on the hazard of separation, 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 but insignificant., with a magnitude which suggests that that the overall empty nest effect will only remain positive for about ten 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[23].

[22] 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.

[23] We assume full compliance throughout our analysis although according to Table 4c 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

Model 2 drops partner not religious dummy, different education dummy, insignificant age difference dummies and insignificant surprise indicators, all of which are both individually and jointly insignificant in Model 1. Model 3 represents the preferred parsimonious specification, after dropping all variables which are not significant at the 5% level, with the exception of the child support related variables and unearned incomes. This parsimonious specification also represents the best fit among all three 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 $\kappa = 1$ (the Weibull distribution) has $\chi^2(1) = 1.43$, Prob>$\chi^2 = 0.23$, while the test of: $\kappa = 1$ and $\sigma = 1$ (the Exponential distribution) has $\chi^2(2) = 19.37$, Prob> $\chi^2 = 0.00$, and the null that $\kappa = 0$ (the Lognormal distribution) has $\chi^2(1) = 2.81$, Prob > $\chi^2 = 0.09$. Hence the Wald tests strongly reject the Exponential and to a lesser extent the Lognormal distributions and are strongly in favour of a Weibull distribution, the results of which are presented in the last two columns of Table 7.

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 (i.e. a lower risk of partnership dissolution). The two sets of results are remarkably similar, implying that the Weibull is a very good approximation of the more general Gamma model. Most of the

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).

Our attempts to model compliance, for separated matched couples, showed that there are essentially no observable characteristics in BHPS that explain compliance although the size of liability is significant. Details available on request. 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 6: Probit Model of Partnership Dissolution: changes in probability, P-values in parentheses

	Model (1)	Model (2)	Model (3)
Incomes			
Wife's Unearned Income – (£1000/wk)	0.033 (0.011)	0.033 (0.011)	0.030 (0.027)
Partner's Unearned Income –(£1000/wk)	-0.015 (0.216)	-0.017 (0.191)	-0.026 (0.052)
Child support related variables			
Indicator for qualifying children	0.003 (0.349)	0.004 (0.288)	
Current CS liability (£1000/wk)	**-0.084 (0.018)**	**-0.087 (0.017)**	**-0.069 (0.034)**
Indicator for wife on IS if separated	0.003 (0.235)	0.003 (0.247)	0.003 (0.198)
CS*Indicator for wife on IS if separated	0.014 (0.723)	0.015 (0.695)	0.008 (0.823)
Indicator for wife on FC if separated	-0.003 (0.362)	-0.003 (0.356)	-0.002 (0.479)
CS*Indicator for wife on FC if separated	0.078 (0.139)	0.079 (0.137)	0.067 (0.199)
Indicator for post 91 partnership	-0.004 (0.174)	-0.004 (0.181)	-0.004 (0.210)
CS*post 91 partnership	0.029 (0.569)	0.026 (0.604)	0.018 (0.721)
Characteristics			
Empty Nest dummy	**0.019 (0.009)**	**0.019 (0.010)**	**0.008 (0.022)**
Years since empty nest	-0.002 (0.170)	-0.002 (0.188)	
Own working hours/week	0.0001 (0.089)	0.0001 (0.101)	0.0001 (0.051)
Partner's working hours/week	-0.0001 (0.003)	-0.0001 (0.003)	-0.0001 (0.001)
Cohabiting	0.026 (0.000)	0.026 (0.000)	0.025 (0.000)
Number of ex-marriages	0.004 (0.000)	0.004 (0.001)	0.005 (0.001)
Age at start of partnership	-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)
Partners from same ethnic group	-0.005 (0.127)	-0.005 (0.133)	
Partners have same religion	-0.004 (0.014)	-0.004 (0.009)	-0.004 (0.005)
Partners not religious	0.001 (0.573)		
Youngest child <5 years	-0.005 (0.015)	-0.005 (0.014)	-0.005 (0.012)
Number of qualifying children	0.002 (0.015)	0.002 (0.020)	0.003 (0.000)
Partners have different education	0.002 (0.252)		

Age difference			
Woman more than 5 years older	0.020 (0.003)	0.023 (0.000)	0.023 (0.000)
Woman 3-5 years older	0.002 (0.695)		
Woman 0-3 years older	-0.001 (0.670)		
Partner 2 to 4 years older	-0.001 (0.558)		
Partner more than 4 years older	-0.002 (0.432)		
Labour Market (as of t-1):			
Employed	0.0003 (0.900)	0.0003 (0.903)	
Unemployed	0.015 (0.039)	0.014 (0.043)	0.015 (0.036)
Partner employed	0.004 (0.137)	0.004 (0.141)	
Partner unemployed	-0.004 (0.201)	-0.004 (0.173)	-0.005 (0.047)
Surprise indicators			
Large positive surprise	-0.006 (0.285)		
Positive surprise	-0.001 (0.544)		
Negative surprise	0.003 (0.116)	0.004 (0.050)	0.004 (0.047)
Large negative surprise	0.010 (0.012)	0.011 (0.006)	0.011 (0.006)
Missing surprise indicator	0.006 (0.089)	0.006 (0.055)	0.007 (0.048)
N (couple-years)	16033	16033	16033
Chi-square (df)	414.52 (38)	408.79 (30)	405.98 (25)
Pseudo R^2	0.1341	0.1325	0.1304
Log Pseudo-likelihood	-1253.02	-1255.33	-1258.37
Akaike Information Criterion	**0.1612**	**0.1605**	**0.1602**

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) 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 (2001) specification, with the exception of earnings which are dropped to minimize the multicollinearity problem.

Model 2: Dropping partner not religious dummy, different education dummy, insignificant age difference dummies and insignificant surprise indicators from Model 1.

Model 3: Dropping indicator for having any qualifying children, years since empty-nest and indicator for partner from same ethnic group and dummies for being employed from Model 2

Table 7: The Generalized Gamma and the Weibull Models:

	Generalized Gamma Model		**Weibull Model**	
	Coeff.	**P-value**	**Coeff.**	**P-value**
Income				
Wife's Unearned Income (£1000/week)	-2.832	0.144	-2.541	0.137
Partner's Unearned Income (£1000/week)	5.301	0.008	4.939	0.015
Child support related variables				
Current CS liability (£1000/wk)	**7.578**	**0.073**	**7.278**	**0.082**
Indicator for wife on IS if separated	-0.638	0.052	-0.585	0.051
CS*Indicator for wife on IS if separated	-0.731	0.878	-1.589	0.729
Indicator for wife on FC if separated	0.103	0.811	0.043	0.921
CS*Indicator for wife on FC if separated	-6.674	0.304	-6.330	0.354
Indicator for post 91partnership	-0.207	0.609	-0.311	0.418
CS*Indicator for post 91 partnership	-1.517	0.768	0.112	0.982
Partnership characteristics				
Empty nest dummy	**-0.645**	**0.127**	**-0.573**	**0.124**
Own working hours/week	-0.020	0.016	-0.019	0.010
Partner's working hours/week	0.011	0.036	0.011	0.041
Cohabiting	-2.023	0.000	-1.971	0.000
Number of ex-marriages	-0.644	0.000	-0.594	0.000
Age at start of partnership	0.110	0.000	0.108	0.000
Partners have same religion	0.487	0.009	0.480	0.007
Youngest child < 5 years	-0.011	0.968	-0.005	0.985
Number of qualifying children	-0.478	0.000	-0.425	0.000
Woman more than 5 years older	-1.643	0.001	-1.690	0.000
Wife unemployed	-1.238	0.030	-1.106	0.019
Partner unemployed	0.442	0.254	0.376	0.294
Surprise indicators				
Negative surprise	-0.464	0.049	-0.398	0.050
Large negative surprise	-0.952	0.006	-0.855	0.006
Missing surprise indicator	-0.699	0.069	-0.612	0.063
Constant	2.411	0.000	2.582	0.000
$\ln\sigma$	0.462	0.000	-	-
κ	0.583	0.094	-	-
P	-	-	0.687	0.000
N (couple-years)	16033		16033	
Chi-square (df)	265.58 (24)		293.12 (24)	
Log Pseudo-likelihood	-622.94		-623.99	
Akaike Information Criterion	**0.0808**		**0.0810**	

coefficients retain their sign although they tend to become slightly less significant than in the discrete time specification. The shape parameter p is very precisely determined with an estimate of 0.687 indicating negative overall duration dependence and a rather sharp decline in the hazard of separation immediately after formation

Figure 4 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 5 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 shows the predicted hazard 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 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 separation rate, potentially reducing the separation probability by around 10% for a 20 year old marriage if all child support 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 decreased the instantaneous hazard by around 29.6% over what we predict it would have been while the new CS reform will increase the hazard, by about 6.9%, for couples with dependent children.

Figure 4: Impact of CS liability on Survival Rates

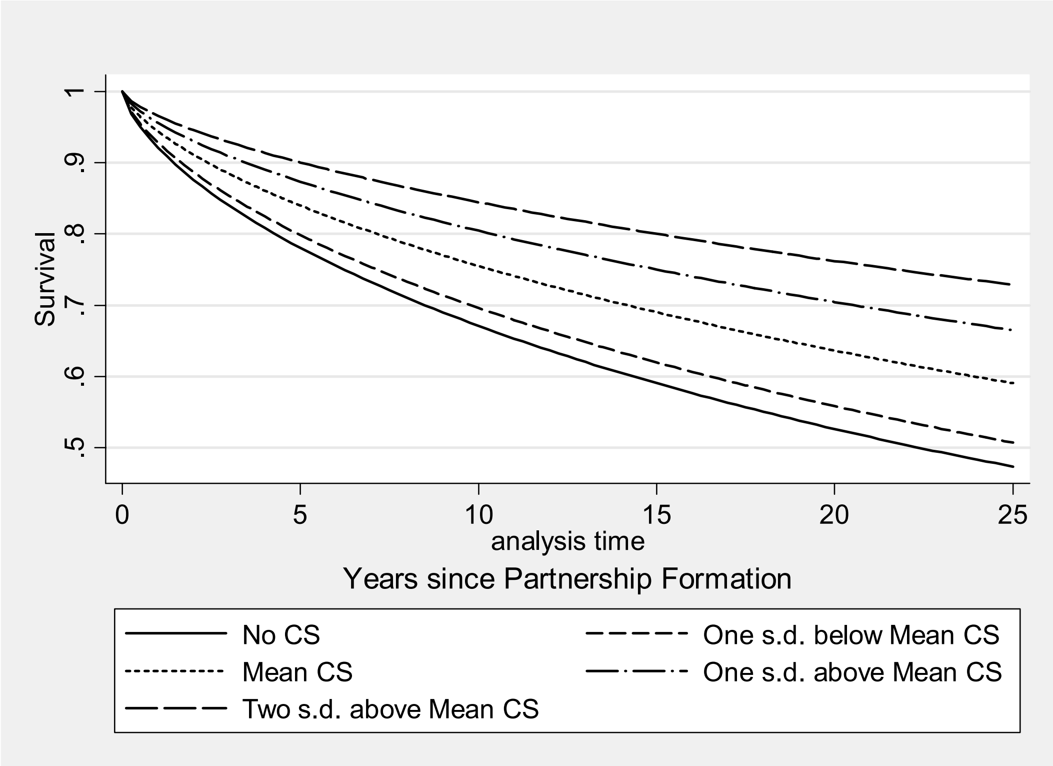
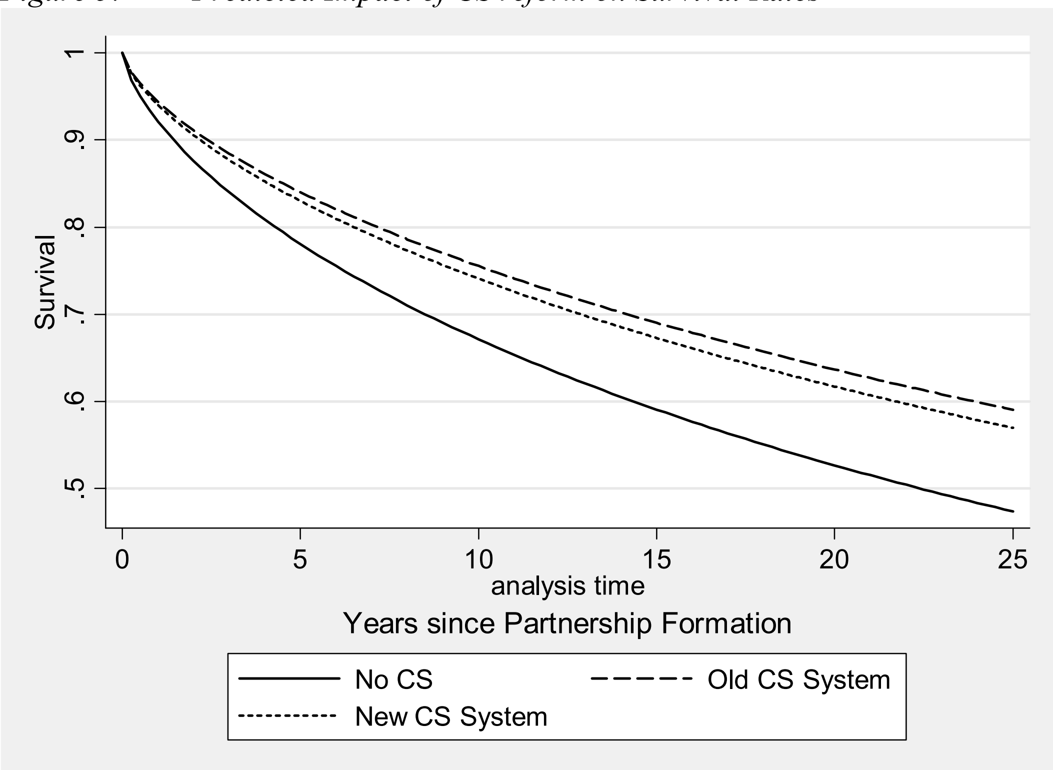


Figure 5: Predicted Impact of CS reform on Survival Rates



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[24], 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 $CS_i = b\left(a.y_{ih} + (1-a).y_{iw}\right)$ where b is a scale parameter indicating the generosity of the CS system, while a is the parameter that weights the separated parents together[25]. Thus $a=1$ implies that CS liability is independent of the PWC's (assumed here to be w) income while a=0.5 implies that the NRP's (assumed here to be h) liability falls by 50% of an increase in PWC's income.

Figure 6 shows the amount of CS contribution from the NRP to the PWC as the weight attached to NRP's income rises. Figure 7 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 £69.0 per week for the father, which amounts to 23.0% 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 £51.2 per week, or 17.0% of their respective net earnings, with the PWC (notionally) contributing an equal share of (typically) her net earnings to make up the balance.

Figure 8 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

[24] 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.

[25] Applying OLS to the sample of BHPS couples with dependent children reveals that, for the UK in our sample period, b=0.207 and a=0.853.

approximately 6% higher if *a=1* compared to *a*=0.5. This corresponds to an instantaneous separation rates of 2.45% if *a=1* compared to 2.68% for a system that was based on the unweighted sum of both parents' incomes, holding the level of CS liability constant.

Figure 6 NRP's CS Liability by Weight on NRP Income

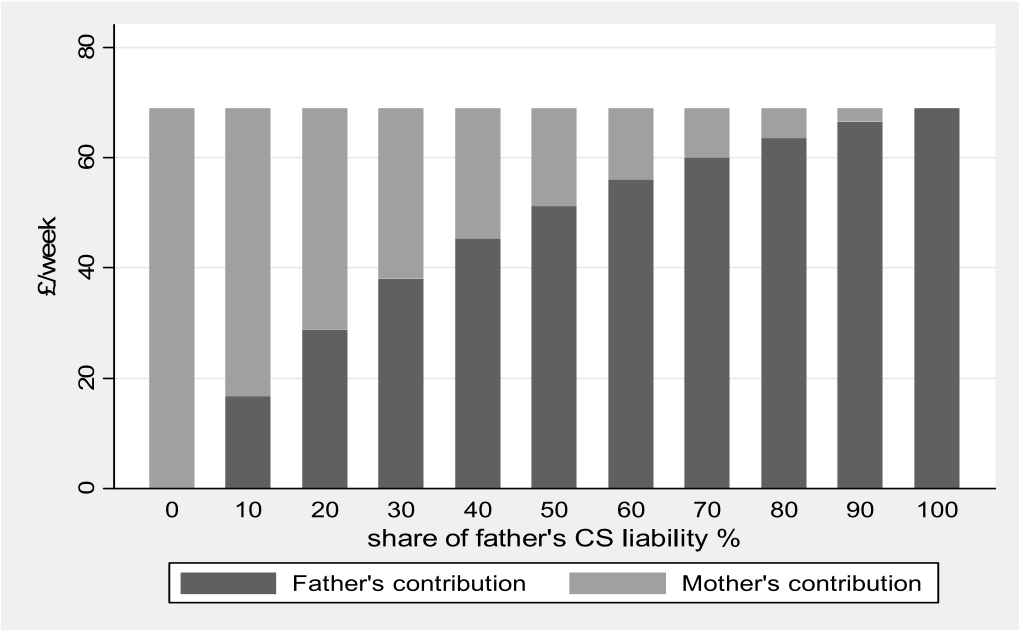


Figure 7 NRP's CS Liability as Share of Net Income

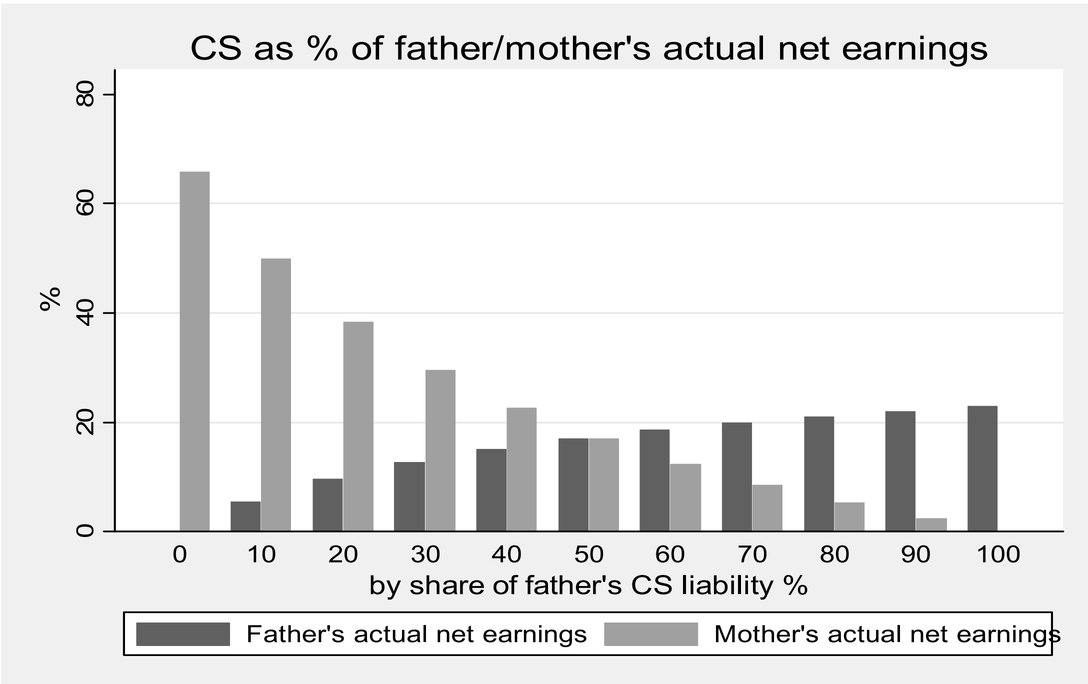
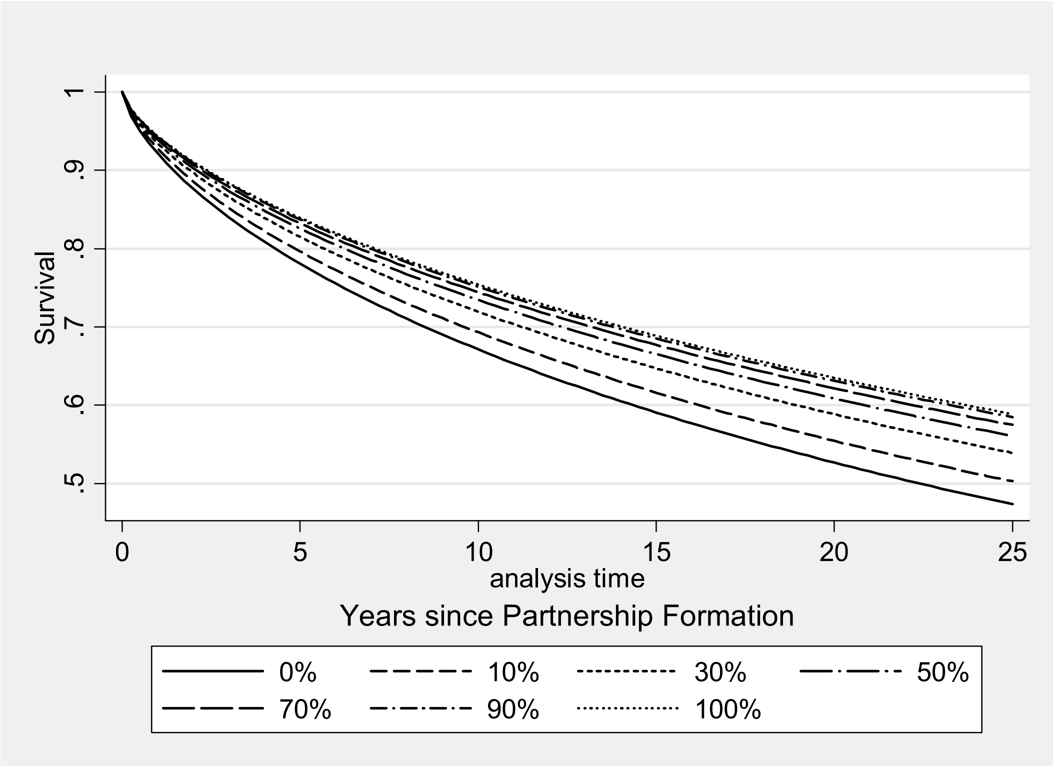


Figure 8 Effects of Income Sharing Rule on Predicted Survival Rates



9. Conclusions

This paper studies the determinants of partnership dissolution in the UK using the British Household Panel Survey (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. In particular, we find that the variation in child support liabilities arising from the introduction of complex rules for CS, that surprised partners in 1992 had important implications for the subsequent separation experience. We find there is very strong evidence that an increase in the implied child support liabilities significantly reduced the dissolution risk. This result still holds when we restrict our sample to couples with qualifying children or dropping the benefit status and post 91 sample dummies and their interactions with CS (which are statistically insignificant), see Table A6 in the Appendix. Indeed, we find that this effect is strong enough to make the effect of the departure of all children (an empty nest) to become insignificant.

We use the estimates to simulate the effect of CS on separation rates on the sample of couples with dependent children. We calculate that the separation rate would have been over 40% higher (i.e. 3.46% instead of 2.44%) were it not for the introduction of the complex CS formula in the UK in 1992 that resulted in large liabilities.

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 2.45% compared to 2.68% 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[26]. The assumptions of no labour supply or repartnership effects are maintained hypotheses could also be tested[27]. But despite our reservations about these assumptions we believe these existing findings do have significant policy implications. For instance, our results suggest that the current child support reform (Department of Social Security (1993)), 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 child support 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. The logic of the present paper 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 given children are 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, welfare of all parties has improved. There is little research on the impact of

[26] 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.

[27] Table A in 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.

separation on well-being of the partners and further research needs to be done to separate out the effects of separation *per se* from its financial consequences, especially on outcomes for children, including their well-being[28].

[28] 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, McLanahan and Robins (1994).

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Appendix A

The system of child support analysed here is that reigning from 1993 to 2003 and is described in some detail in CPAG (1999) and here their notation is used to facilitate comparison between our summary exposition and the fine details. Their steps in the formula can be compressed into the following single relationship, which is broadly based around the "proposed amount" (P) for the parent with care and non-resident parent:

$$P = 0.5*F \qquad \text{if } F+G < 2A$$

$$P = c*F + (1 - 2*c)*A*(F/(F+G)) \qquad \text{if } F+G \geq 2A$$

where: $F = D - B$ (= 0 if non-resident parent or new partner on Income Support (IS) or Job Seekers Allowance (JSA)) where D = net income for the non-resident patents, and B = exempt income for the non-resident parent; $G = E - C$ (= 0 if parent-with-care or new partner on IS, JSA, Disabled Workers Allowance (DWA) or Working Families Tax Credit (WFTC) which is the UK equivalent to Earned Income Tax Credit, EITC) where E = net income for the parent-with-care and C = exempt income for the parent-with-care; A = maintenance requirement; and c = 0.15, 0.20 and 0.25 for 1, 2 and 3 plus qualifying children respectively. Since net income is set to zero for the listed benefit recipients and also excludes several other types of benefits, it mainly captures net earnings and investment income[29]. Exempt income includes an allowance for supporting qualifying and new children[30] in the household, but this is reduced if a new partner has sufficient income to help support any new children. Exempt income also includes housing costs and travel-to-work costs. The maintenance requirement depends on the number and ages of the qualifying children. Note that non-resident parents on IS or JSA have a zero proposed amount.

In addition, the final liability (L) is subject to three separate maximums, partly to ensure that non-resident parents are left with adequate resources to support themselves and their families:

$$L = \max(P,\ J,\ 0.3*D,\ 0.85*(R - V))$$

[29] It also includes the income of own children (qualifying or new).

[30] Qualifying children are the natural children of the separated parents. New children are defined as children of one of the parent and a new partner. Stepchildren are defined as natural children only of the new partner of one of the parents.

where: J = maximum dependent on modified values of A, F and G; R = family income for the non-resident parent; and V = protected income for the non-resident parent family. The family income for the non-resident parent includes all income except certain benefits for the non-resident parent, any new partner and any dependent children. The protected income includes an allowance for family size and ages of children, housing costs, net council tax and travel-to-work costs. There is also a minimum liability of roughly 10% of the current Income Support rate for a single person, which currently stands at £5.20 a week. Those exempt from this minimum have a zero liability if L is below this minimum and exemptions include all those non-resident parents with any dependent children in their new household.[31]

The relationship between the liability and non-resident parent income has three steps. At low levels of income, the liability is fixed at the minimum or at zero depending upon whether the non-resident parent is exempt. Past the point where income is sufficiently high for L to exceed £5.20, the liability rises at a rate of 50% with any additional income. If income is higher than the point where the children's needs are deemed to have been met ($F+G \geq 2A$), the liability rises at a lower rate with income to allow the children to share in the good fortune of a high income non-resident parent. The income of the parent-with-care affects the liability only in the third of these steps and in determining the point where the third step begins. The higher the income of the parent-with-care, the lower the amount of non-resident parent income where the third step begins and the slower the increase in the liability with non-resident parent income in the step. Hence, increases in parent-with-care income reduce the liability, but in a non-linear fashion.[32] The number of qualifying children influence the liability both directly in the third step for non-resident parent income and indirectly by increasing the exempt income for the parent-with-care. Finally, a rise in the non-resident parent's housing or travel-to-work costs reduces the liability through its impact on exempt income. Similarly, a rise in the parent-with-care's housing to travel to-work costs increases the liability. Hence, there are incentives to increase spending on either of these items.

[31] Exemptions include those non-resident parents with any dependent children in their new family, those receiving certain disability benefits, those under the age of 16, those under the age of 19 and in full-time education and those with net income below the minimum.

[32] In addition, the higher the parent-with-care income, the lower the maximum liability level set in J.

Appendix B

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

Years Separated	Equivalised Income		Poverty Rates (%)		Labour Market Participation (%)		Repartnership (%)	
	Fathers	Mothers	Fathers	Mothers	Fathers	Mothers	Fathers	Mothers
<=-4	140.1	140.1	38.3	38.8	80.4	54.7		
-3	146.8	146.8	38.1	38.1	88.1	64.3		
-2	142.1	142.1	38.0	38.0	85.0	65.0		
-1	159.6	159.6	30.8	30.8	88.0	60.9		
0	152.3	152.3	30.7	30.7	84.0	60.7		
1	176.6	108.6	34.3	54.5	83.9	59.4	19.6	17.5
2	203.7	125.5	26.3	43.9	81.6	62.3	28.9	20.2
3	222.3	136.8	28.0	42.0	86.0	58.0	31.0	22.0
4	218.7	142.0	22.1	39.5	83.7	52.3	36.0	31.4
>=5	241.5	171.7	19.9	23.7	87.2	67.3	57.3	42.2
Total	180.1	144.2	30.0	37.0	84.6	60.6	37.3	28.4

Table A6: Specification Checks (Probit Model of Partnership Dissolution: changes in probability, P-values in parentheses)

	Couples who have had children (main sample N=16033)		Subsample of Couples With qualifying children (N=7834)	
	Baseline Model (1)	Drop Insignificant CS related variables (2)	Baseline Model (3)	Drop Insignificant CS related variables (4)
Incomes				
Wife's Unearned Income – (£1000/wk)	0.030 (0.027)	0.033 (0.016)	0.079 (0.002)	0.078 (0.003)
Partner's Unearned Income – £1000/wk)	-0.026 (0.052)	-0.026 (0.052)	-0.030 (0.392)	-0.030 (0.370)
Child support related variables				
Current CS liability (£1000/wk)	**-0.069 (0.034)**	**-0.047 (0.020)**	**-0.158 (0.024)**	**-0.081 (0.034)**
Indicator for wife on IS if separated	0.003 (0.198)		0.003 (0.701)	
CS*Indicator for wife on IS if separated	0.008 (0.823)		0.060 (0.462)	
Indicator for wife on FC if separated	-0.002 (0.479)		-0.007 (0.314)	
CS*Indicator for wife on FC if separated	0.067 (0.199)		0.156 (0.110)	
Indicator for post 91partnership	-0.004 (0.210)		-0.004 (0.588)	
CS*Indicator for post 91 partnership	0.018 (0.721)		0.040 (0.659)	
Characteristics				
Empty Nest dummy	**0.008 (0.022)**	**0.008 (0.024)**	**-**	**-**
Own working hours/week	0.0001 (0.051)	0.0001 (0.191)	0.0002 (0.223)	0.0001 (0.497)
Partner's working hours/week	-0.0001 (0.010)	-0.0001 (0.011)	-0.0002 (0.006)	-0.0002 (0.006)
Cohabiting	0.025 (0.000)	0.026 (0.000)	0.037 (0.000)	0.037 (0.000)
Number of ex-marriages	0.005 (0.001)	0.005 (0.000)	0.003 (0.357)	0.003 (0.283)
Age at start of partnership	-0.001 (0.000)	-0.001 (0.000)	-0.002 (0.000)	-0.002 (0.000)
Log duration of partnership	-0.013 (0.000)	-0.012 (0.000)	-0.017 (0.000)	-0.017 (0.000)
Partners have same religion	-0.004 (0.005)	-0.004 (0.006)	-0.005 (0.076)	-0.005 (0.090)
Youngest child <5 years	-0.005 (0.012)	-0.005 (0.012)	-0.007 (0.052)	-0.007 (0.053)
Number of qualifying children	0.003 (0.000)	0.003 (0.000)	0.003 (0.176)	0.003 (0.108)
Woman more than 5 years older	0.023 (0.000)	0.024 (0.000)	0.061 (0.000)	0.063 (0.000)
Wife unemployed	0.015 (0.036)	0.015 (0.031)	0.009 (0.503)	0.010 (0.452)
Partner unemployed	-0.005 (0.047)	-0.005 (0.072)	-0.010 (0.055)	-0.010 (0.065)
Negative surprise	0.004 (0.047)	0.004 (0.046)	0.008 (0.026)	0.009 (0.026)
Large negative surprise	0.011 (0.006)	0.011 (0.005)	0.019 (0.008)	0.020 (0.008)
Missing surprise indicator	0.007 (0.048)	0.007 (0.054)	0.008 (0.243)	0.008 (0.264)
N (couple-years)	16033	16033	7834	7834
Chi-square (df)	405.98 (25)	395.90 (19)	203.44 (24)	200.18 (18)
Pseudo R^2	0.1304	0.1280	0.1002	0.0977
Log Pseudo-likelihood	-1258.37	-1261.80	-899.13	-901.67
Akaike Information Criterion	**0.1602**	**0.1599**	**0.2359**	**0.2350**

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) 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 (see Maddala (1983) p488).