Child Support and Educational Outcomes: Evidence from the British Household Panel Survey

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Abstract

There is some evidence to support the view that Child Support (CS), despite low compliance rates and a strong interaction with the welfare system, has played a positive role in reducing child poverty among non-intact families. However, relatively little research has addressed the role of CS on outcomes for the children concerned. There are good reasons for thinking that CS could leverage better outcomes than other forms of income support and, using a sample of dependent children in non-intact families from the British Household Panel Survey (BHPS), we find that CS received has an effect which is at least 10 times as large as that associated with variations in other sources of total household net income for two key educational outcomes: namely school leaving at the age of 16, and attaining 5 or more good GCSEs. We show that this remarkable and strong result is robust and, in particular, can be given a causal interpretation.

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I. Introduction

The impact of parental separation (excluding death) on children has been a longstanding concern of social scientists. While there has been a general consensus that parental separation is associated with adverse outcomes for children (for surveys, see e.g. Amato and Keith (1991), Haveman and Wolfe (1995) and Amato (2001)), the extent to which this correlation is causal is far from clear. Causality becomes difficult to infer if there are important omitted variables that are likely to be correlated with separation. Most studies provide only the reduced form effect of separation - they do not provide evidence that it is separation *per se* that matters or the unobservables associated with separation. Thus, while such studies identify a problem, they fail to tell us what the problem is. For example, many such studies typically omit income despite the considerable evidence that income does affect child outcomes (see e.g. Dahl and Lochner (2005), Plug and Vijverberg (2003, 2005), Chevalier *et al.* (2005)and Francesconi *et al.* (2005)). That is, few studies have attempted to separate out the effects of one parent (mostly the father) leaving on the outcomes for the children, from the effects of that parent's income leaving.

Child support (CS) is the monetary transfer made by the non-custodial parent (usually the father) to the custodial parent (usually the mother), for the care and support of children of a relationship that has broken down. In most countries there is a system of CS that both reduces the financial effects of separation on children and raises the costs of separation to the non-custodial parent.[1]

It is widely believed that CS, both the possibility of any CS being paid and the amount paid (i.e. the extensive and the intensive margin), has played a positive role in reducing child poverty among non-intact families (see e.g. Meyer and Hu (1999)).[2] Despite the attention paid to child poverty because of its apparent adverse effects on outcomes for children, and to the role that CS plays in reducing child poverty, little research has been done on the role of CS on the outcomes for the children concerned. This gap in our knowledge is particularly troublesome in the UK because the system

[1] See Cancian *et al* (2003) for US evidence and González (2005) for evidence from across 16 countries.

[2] CS not only affects child outcomes through a direct income effect, it may also cause indirect incentive effects on maternal quality selection, fertility decisions, prenatal investments, and birth outcomes. Walker and Zhu (2006) show that an increase in CS liabilities arising from the introduction of complex CS rules in 1993 significantly reduced the risk of partnership dissolution for couples in the UK. Aizer and McLanahan (2006) find that stricter state CS enforcement leads to fewer out-of-wedlock births and an increase in both the average education of mothers and investment of prenatal care in the US.

of CS in the UK is in the process of being reformed: moving from a system where CS is determined by a very simple rule, to one that seems likely to entail a great deal more idiosyncratic variation in CS. In particular, it seems likely that the reform will reduce CS payments, unless compliance can be radically improved.[3]

Notable US exceptions are the studies by Knox (1996) and Argys et al. (1998), both of which investigate causation of CS on child educational outcomes by exploiting cross-state variation in CS enforcement using the National Longitudinal Survey of Youth (NLSY). The sample in Knox (1996) contains relatively young mothers aged between 23 and 31 and is disproportionately from lower socioeconomic status and minority families. Knox shows that an increase in CS of $1,000 per year increase children's reading and maths test scores for 5-8 year olds in single-parent families by around 7% of a standard deviation, which is one order of magnitude higher than that from other income sources. Moreover, the CS coefficient in the instrumental variables model using cross-state variation in CS levels, and other indicators of local economic environment, remains statistically significant at the 10% level. Knox also shows that CS predicted by economic conditions and state average CS levels do *not* help explain variations in achievement test scores *for children living with both parents*.

Argys et al. (1998) revisit the NLSY data a few years later, as the sample becomes more representative of children who are born to NLSY mothers. While showing that the CS coefficient is no longer statistically significant under the Knox (1996) specification with the updated sample, Argys et al. estimate a specification which controls for race and out-of-wedlock births and distinguishes between cooperative and non-cooperative awards. The additional positive effect of CS on children's cognitive outcomes is only found for blacks in the divorced/separated sample and for whites in the non-marital birth sample. Instrumenting CS and total family income using local economic and demographic indicators as well as policy variables relating to divorce, CS and welfare generosity of the state, the authors find that the effects of CS persist only for blacks in the divorced sample.

[3] The system of CS in the UK was reviewed by David Henshaw for the Department of Work and Pensions in July 2006. His main recommendations include encouraging most parents to make their own CS arrangement and significantly increasing the level of CS disregards in benefit calculation. For the details see http://www.dwp.gov.uk/childsupport/henshaw_report.asp.

Researchers have hypothesized the following mechanisms through which CS may have beneficial effects on children over and above income from other sources (Knox (1996), Aizer and McLanahan (2006)):

1) Kooreman (2000) presents empirical evidence that the marginal propensity to consume child clothing out of exogenous child benefits is much larger than that out of other income sources for households with one child in the Netherlands. He interprets this finding as evidence of the labelling effect of a child benefit system. Such a "labelling (hypothecation) effect" might also arise in the context of CS, so that CS may improve children's outcomes more than other sources of income, if mothers feel obliged to spend it directly on children. Del Boca and Flinn (1994) find that coefficients associated with CS and alimony income are significantly higher than those for other types of income in estimated Engel curves for expenditures on child-specific goods.
2) CS may affect relationships between the non-custodial father and the custodial mother and between the father and his child differently than income from other sources. On form of such improved family dynamics would be increased visitation. However, the effect is theoretically ambiguous *a priori* (e.g. Ermisch (2005), Del Boca and Ribero (2001)), and both Knox (1996) and Argys et al. (1998) found little evidence that contact between the child with the absent father matters.
3) It is possible that the adverse effects that may be associated with other forms of income are absent from CS. For example, welfare income may be stigmatised, and income from maternal employment may carry imply poorer childcare (see Gregg *et al.* (2005) and Rhum (2004)); and while step-parents might be associated with higher levels of household income step-parents may feel less altruistic towards a step-child (see, for example, Case et al. 1999).
4) Finally, there may be a degree of reverse causality. Aughinbaugh (2001) suggests that custodial mothers might invest *more* in the child as a signal to the absent father in order to secure future CS payments. Indeed, she does find that measures of child achievement have a significant positive effect on both the possibility of any CS being paid and the amount paid.

However, a plausible alternative to the above explanations is that the relationship between CS receipts and child's educational outcomes might reflect a selection effect (unmeasured heterogeneity among families) rather than a causal effect. For instance, payment of CS by an absent father might be correlated with his commitment, and perhaps that of the mother, to the well-being of the child. From a policy point of view, it is extremely important to be able to discriminate between these two explanations.

Our own empirical work is based on the British Household Panel Survey (BHPS). Unlike the NLSY which overrepresents younger mothers, our sample is representative of all non-intact families. Using a sample of dependent children under 16 in non-intact families whose real educational outcomes are observable, we find that CS received has an effect which is at least 10 times as large as that of other household net income for two key educational outcomes: namely school leaving at the age of 16, and attaining 5 or more GCSEs.[4] This result is remarkably robust with respect to adding controls for characteristics of the child and the custodial mother, which can be interpreted as evidence against a pure selection story. To investigate the extent to which the relationship is a causal one, we then instrument current CS receipts using retrospective information on mother's fertility, relationship and employment before the birth of the child. We find that the strong CS effect found in our simple probit models holds up in the IV specification.

II. Data

The British Household Panel Survey (BHPS) is an ongoing nationally representative sample of some 10,000 original sample members (OSMs) recruited in 1991. These OSMs, and all adult members of their families, are re-interviewed annually in subsequent years. The BHPS collects information on household composition, housing, employment, education, health and incomes in all waves. It is a particularly rich dataset with lifetime histories of marriage, cohabitation, fertility and employment transitions, allowing us to match all children to their natural parents and

[4] These are arguably the two most important measures of educational attainment in the UK: the former captures the dropout rate at the minimum school leaving age while the latter is a key quality indicator of the standard the pupil has achieved. The General Certificate of Secondary Education **(GCSE)** has been the principal means of assessing pupil attainment at the end of compulsory secondary education since 1988.

establish the timing of departure of any absent parents, even for partnerships dissolved before the sample began.

< Figure 1 here >

We focus on school leaving and the attainment of 5 or more GCSEs by the age of 16 for all children in non-intact families in this paper. Figure 1 reports the two outcomes and CS receipts by family types, where non-intact families are further divided into two groups depending on whether or not the parents were living together when the child was born (denoted as separated and single mothers in Table 1). A child is defined as living in a non-intact family, if he or she does not live with both natural parents at any point in time until he/she reaches the age of sixteen. Note that the way the family type is defined means that children living in the *same* family can be classified into *different* categories.[5] It is apparent that the outcomes and CS receipts differ substantially across intact and non-intact families, and among the latter category they also differ between children whose parents separated and children born out-of-wedlock. For instance, the dropout rates at 16 is 17% for children living with both natural parents, as opposed to 27% for children whose parents have separated during their childhood and 35% for children living in non-intact families since birth. On the other hand, the probability of attaining 5 or more good GCSEs for children born-out-wedlock is only slightly over half of that for children living in intact families, with children in separated families in between.

< Table 1 here >

Table 1 reports summary statistics by family types and highlights the importance of distinguishing between in-wedlock and out-of-wedlock births. There are 1791 distinct youths whose age 16 educational outcomes are observable in the first sixteen waves of the BHPS, of which 501 (or 28.0%) live in non-intact families. While over half of all separated mothers have received any CS, only three in ten single mothers do so. Conditional on receiving any CS, separated mothers received, on average, £51.30/week, twice as much as the £25.40/week received by single mothers.

[5] Although children of the same natural mother can be classified into different categories due to remarriages, our modest sample size precludes a full exploitation of the differences between siblings (see e.g. Bjorklund and Sundstrom (2006)). The sample size also does not allow us to further distinguish between (previously) married and cohabiting couples, who are treated symmetrically for social security and CS purposes in the UK.

The mean equivalised[6] total household net income for separated and single mothers is only 71% and 59% respectively of that for their intact counterparts, at £230 per week. CS only accounts for less than 9% of total household income for separated mothers, and less than 4% for single mothers.

Table 1 also reports child and mother characteristics which will be controlled for in the regression analysis later. The proportion of boys and twins is similar across family types. On the other hand, the differences in the average number of dependent children in the sample period and birth order might reflect complex interactions of fertility and repartnership. Recent studies suggest that family size and birth order are potentially important for children's educational attainment (see e.g. Booth and Kee (2005) and Lundberg (2005)). Single mothers have significantly lower education than either separated or intact mothers, while there is nothing to distinguish between the latter two groups. Table 1 also summarizes the potential instruments based on retrospective information on the mother's fertility, relationship and employment history up to the point the child concerned was born. It turns out that mothers in non-intact families, and single mothers in particular, are younger at their first birth and have shorter full-time employment experience. Perhaps more surprisingly, the dissolved partnership for separated mothers is almost as likely to have been a first marriage or cohabitation spell (lasting 3 months or more) as is the case for intact mothers. Our identification strategy relies on these lifetime events which predate the birth of the child.

III. Analysis and Results

1) Probit Models

We focus on the impact of CS on child educational outcomes in this paper. In the spirit of Knox (1996) and Argys et al. (1998), we decompose mean equivalised total household net incomes into three components according to sources: CS receipts, step-fathers' incomes and other total net incomes which are assumed to be attributable

[6] In our econometric analysis below, we allow for economies of scale in consumption using the popular "square root scale" which is commonly used in international comparisons (e.g. by the OECD). Note that this scale can be regarded as a special case of the Buhmann *et al.* (1988) class equivalence scale $M=h^{\theta}$ where h is the household size and $\theta=0.5$. However, our empirical results are remarkably robust with respect to the range of the value of the parameter θ commonly used in the literature, such as a value of 0.67, as in Jenkins and Cowell (1994); or indeed other popular equivalence scales, such as the "OECD-modified scale" which assigns a value of 1 to the household head, of 0.5 to each additional adult and of 0.3 to each child.

to the mother. Hence the differences in coefficients across the different income sources measure the differential impact of CS and step-fathers' incomes than those from other income sources. However, unlike Argys et al. (1998), we do not control for pre-separation family income, as we only observe separations taking place in the sample period for one third of the non-intact families.[7]

< Table 2 here >

Table 2 presents the marginal effects for the probit models of both outcomes for the sample of non-intact families. Model 1 is the simplest model specification which only controls for CS, step-fathers' incomes and other net incomes, all equivalised and averaged over non-intact years in the sample period. Model 2 adds such child characteristics as gender, number of dependent children in the family, an indicator for being twins and dummies for birth order of the child. Finally, Model 3 adds two important characteristics of the mother: the age she left full-time education and an indicator for being non-white.

It is clear that of the three key variables of interest; only CS is statistically significant in explaining the differences in child educational outcomes. A £10 per week increase in equivalised CS receipt (the sample mean is £12.3 per week) will decrease dropout rates by about 3 percentage points and increase GCSE pass rates by just over 4 percentage points respectively. Comparing to the other two income sources, this effect is an order of magnitude higher. On the other hand, step-father's income has no significant effect on either school dropouts or attaining GCSEs once we control for child characteristics. The residual category of other family net incomes turns out to have no statistically significant effect whatsoever on either outcome. Moreover, the size of the CS effects is remarkably robust with respect to the successive inclusion of child and mother characteristics. If the observed correlation between CS and educational outcomes are driven by a selection election, we would expect the size of the CS coefficients to decrease in absolute value as we add more controls. Hence we interpret the robustness of the CS effect in the simple probit models as evidence against a pure selection story.

2) Instrumental Variable Estimation

[7] Table 1 shows that this window of observation is 6.0 years for separated mothers and 7.4 years for single mothers on average.

CS estimates presented in Table 2 will suffer from endogeneity bias if the unobserved heterogeneity across households affect both CS receipts (and payments) and children's outcomes. Besides, measurement errors in CS arising from misreporting will also lead to biased results in probit models. The standard approach to deal with endogeneity and errors-in-variables is two-stage least squares[8].

The set of instruments we use is based on events that took place before the birth of the child, i.e. retrospective information on mother's fertility, relationship and employment. To maximize efficiency, we also carry out statistical tests which check for redundancy of (combinations of) instruments. The resulting preferred parsimonious specification is based on the following three excluded variables: the log of mother's age at first birth; an indicator for whether the dissolved partnership is the first marriage or cohabitation (lasting 3 months or more); and the number of years the mother has worked as a full-time employee before the birth of the child concerned.

< Table 3A+3B here >

Table 3A presents the first stage estimates for the CS equation. Neither step-fathers' incomes nor other net family incomes have any effect on CS receipts. As might be expected, both a later age at first birth and being a first marriage increase CS receipts, while more years of full-time employment by the mother before birth has the opposite effect. Note that the log of mother's age at first birth and years of full-time employment are all individually significant at the 5% level while first marriage is marginally significant at the 11% level[9]. The F-statistic for the joint significance of the excluded variables is well above 10 in this single endogenous regressor specification, thus readily satisfying the "rule of thumb" for weak identification not to be considered a problem (see Stock and Watson (2003)).

Table 3B presents the second stage 2SLS estimates for the main model specification which controls for child and mother characteristics as outlined in Model 3 in Table 2. In line with the probit estimates, neither step-fathers' incomes nor other total net income has a significant effect on the two educational outcomes. On the

[8] We fail to reject the exogeneity of step-fathers' incomes and other net family incomes at even the 30% level of significance for both the school leaving and the GCSE equations in the preferred specification in Table 3.

[9] We include first marriage in our preferred baseline 2SLS specification despite its marginal statistical significance because this instrument variable is key to the separate identification of the effect of the possibility of any CS being paid and the effect of the amount paid in the next subsection.

other hand, the CS effect found in the probit models remains statistically significant for both outcomes under this linear probability specification. While the CS effect has increased substantially in size, in comparison to the probit specification, so has the standard error.[10] Indeed, given the large standard errors, one cannot reject the null hypothesis that the coefficients are the same across the two different specifications, at least for the school leaving equation. In interpreting these findings, more emphasis should be placed on the direction and statistical significance rather than the exact size of the coefficients for CS. It is also worth noting that the IV models pass both the Anderson IV relevance test and the Sargan over-identification test of all instruments at conventional level of significance. Moreover, the Cragg-Donald Wald statistic is greater than the critical value for 5% maximal IV relative size for the Stock-Yogo weak instruments test, thus lending strong support to the validity of our instruments.

3) Two-stage Estimation with CS Predicted by a Selection Model

It is apparent from Table 1 that only half of separated mothers and one third of single mothers ever receive CS. From an econometric point of view, this implies that it is important to allow for left-censoring of CS (i.e. we don't observe negative CS in our data). From a policy perspective, it is also of interest to separate out the effect of receiving *any* CS from that induced by the variation in the *amount* of CS received.

In the following, we will estimate a two-stage model where CS is predicted by the Heckman Full Maximum Likelihood model in the first step (see Maddala (1983) and Heckman (1979)). This approach provides for consistent estimates in the presence of a censored endogenous variable and has been widely used in labour economics (see e.g. Jacobsen and Rayack (1996)).

< Table 4 here >

Table 4 shows the Heckman selection model estimates which are used to predict CS for the second stage. The separate identification of the possibility that the mother receives any CS over the sample period from that the amount received conditional on receiving is achieved using the same 3 instruments as in our 2SLS estimation in Tables 3A and 3B. In particular, first marriage is shown to have a large

[10] This is a common feature of the IV approach in general, given the trade-off between consistency and efficiency. Moreover, least square estimate has a persistent bias towards zero when the regressor concerned is measured with error. This is known as the attenuation effect (see e.g. Greene (2000)). However, it could also be consistent with a Local Average Treatment Effect (LATE) or a credit constraint story.

positive impact on receiving any CS. On the other hand, mother's age at first birth appear to have a positive effect on CS amount while mother's number of years as full-time employee before birth of the child has a negative effect. The signs of all 3 variables seem to make sense and are statistically significant. However neither the estimate of the cross-equation correlation, ρ, nor the inverse Mills ratio term, λ, is significantly different from zero.

< Table 5 here >

Using predicted CS from the MLE selection models and bootstrapping standard errors, we find the effects of CS in Table 5 are remarkably similar to our baseline 2SLS results. A £10 per week increase in equivalised CS received will reduce dropout rates by 5.6 percentage points while increase GCSE pass rates by 11.0 percentage points respectively. Although stepfather's income now appears to be statistically significant for attaining 5 or more good GCSEs, the size of the effect is still an order of magnitude smaller than that of the CS.

4) Sensitivity Analysis

One might be concerned with the use of a measure of CS, which is averaged over the non-intact years in the sample period. Table 6 presents the second-stage IV estimates[11] with binary measures for whether the mother has ever received CS during the sample period.

< Table 6 here >

It is striking that the CS effect remains significant for both outcomes, despite the loss of efficiency through using dichotomous measures of CS receipt as indicated by the drop in the Anderson IV relevance test statistic and the Cragg-Donald Wald F statistic.[12] These results support the hypothesis that improved family dynamics associated with the receipts of **any** CS (i.e. change at the extensive margin) are just as important as that induced by changes in the amount of CS received (i.e. change at the intensive margin) as far as children's educational outcomes are concerned.

< Table 7A+7B here >

[11] To save space, first stage IV estimates for the sensitivity analyses are not shown but will be available from the authors upon request.

[12] However, the test of joint significance of the excluded instruments still satisfied the "rule of thumb".

Table 7A and 7B present the second-stage IV estimates separately for non-intact families with below median family net incomes and above. For the poorest half of non-intact families which is overrepresented by single parents and benefit recipients, the CS effect is only statistically significant for attaining 5 or more GCSEs, although it still has the right sign for the school leaving equation. One possible explanation for this finding is that income has to be above a critical threshold for CS to have a significant effect. The complex interaction between CS and social security benefits that withdraw out-of-work benefits pound for pound might also be responsible. Besides, the measurement error problem is also likely to be more acute for benefit recipients.

In contrast, Table 7B clearly shows that the CS effects remain statistically significant for non-intact families with above median incomes, although only at 10% for school leaving decisions. The fact that the size of the CS effect on attaining 5 or more GCSEs are quite similar across the two subsamples while that on school leaving is almost twice as big for the poor half of the sample - implying that CS matters even more for the school leaving of children who are poor.

< Table 8 here >

Finally, Table 8 presents evidence for separated couples only, which accounts for around 80% of all non-intact families. Although the sizes of the CS effects are somewhat reduced, especially for the school leaving equation, they remain significant at the 5% significance level for both equations. This suggests that the causal effect we find earlier is not driven by the inclusion of children born out-of-wedlock.

Taken together, our sensitivity analysis suggests that the effect of CS on attaining 5 or more good GCSEs is remarkably robust across various types of non-intact families or the income distribution. However, the effect of CS on school leaving decisions appears to be more heterogeneous, with a stronger response for children born out of wedlock or living in poorer families.

IV. Conclusion

Our findings indicate that CS payments have a beneficial effect on educational outcomes for children living in non-intact families in the UK that is well in excess of the effect of income from other sources. This result is very robust with respect to the successive addition of controls for child and mother characteristics, and hence offers little support to the selection explanation of the CS effect.

Moreover, our instrumental variable estimated coefficients of the CS effect, based on retrospective information on fertility, relationship and employment of the mother, remain highly significant. We carry out sensitivity analyses to ensure that our results are not caused by weak instruments or the inclusion of children born out-of-wedlock. Our robustness checks also suggest that the CS change at the extensive margin is at least as important as the change at the intensive margin and the CS effects are not confined to any particular part of the income distribution.

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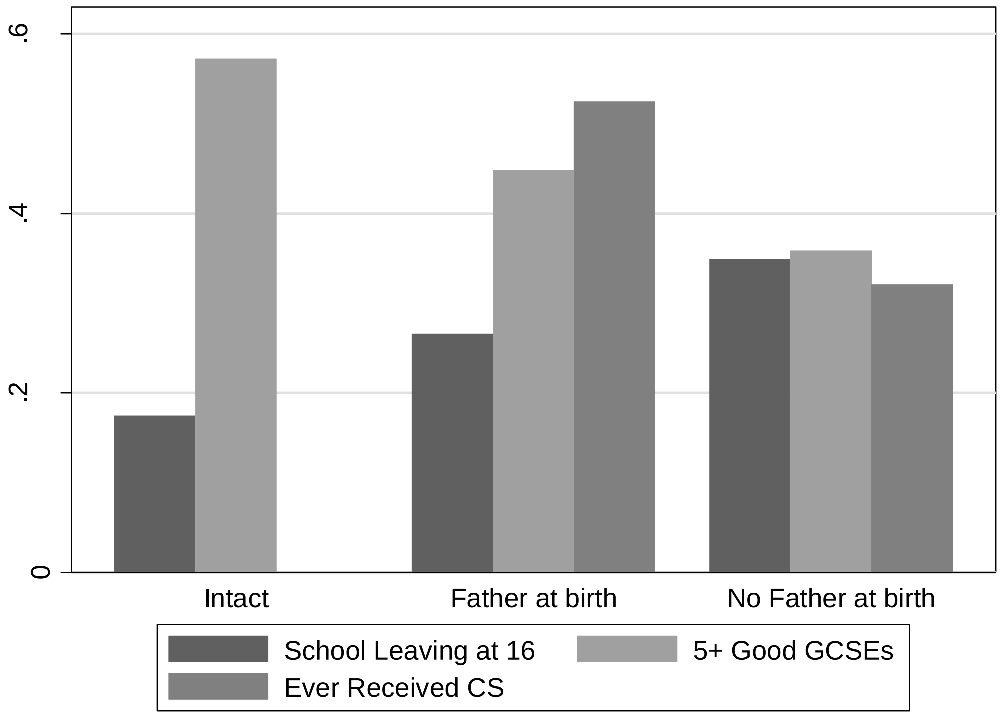
Figures

Figure 1: Educational Outcomes and CS receipts by Family Type



Tables

Table 1: Summary Statistics by Family Types

Family Type	Intact	Separated mothers	Single mothers
School leaving at 16 (%)	17.4	26.6	34.9
5 or more Good GCSEs (%)	57.2	44.8	35.8
Proportion of time mother repartnered (%)	-	37.3	33.9
Ever Repartnered (%)	-	53.9	50.9
Ever received CS	-	52.4	32.1
CS Received conditional on receiving (£/week)	-	51.3	25.4
Equivalence scale ($n^{0.5}$)	2.10	1.93	1.94
Equivalised CS receipt (£/week)	-	14.2	5.0
Equivalised Step-father's income (£/week)	-	42.0	33.6
Equivalised other income (£/week)	229.7	107.3	97.7
Equivalised total income (£/week)	229.7	163.6	136.3
Child's characteristics			
Child being boy (%)	50.1	51.1	55.7
Number of kids in the family	1.9	2.0	2.3
Child being twin (%)	2.3	1.5	2.8
1st natural child of mother	43.5	44.3	78.3
2nd natural child of mother	37.5	37.2	11.3
3rd natural child of mother	14.2	13.9	5.7
4th or higher order natural child of mother	4.8	4.6	4.7
Mother's characteristics			
Mother's age left full-time education	17.5	17.6	16.6
Mother non-white (%)	4.4	4.1	4.7
Retrospective information of the mother (IVs):			
Mother's age at 1st birth	24.9	23.3	20.6
(Dissolved) relationship first marriage (%)	86.0	84.6	-
Years as f/t employee before birth of the child	6.02	4.79	2.47
Proportion of time child not living with dad (%)	-	42.9	100.0
Number of waves observed in sample	7.1	6.0	7.4
Obs	1290	395	106

Note: CS and income are in Jan 2006 prices.

Table 2: Probit Estimates, Marginal Effects

	Left school at 16 = 1			Attained 5+ GCSEs = 1		
	Model 1	Model 2	Model 3	Model 1	Model 2	Model 3
Equiv. CS Receipt	**-0.376**	**-0.372**	**-0.323**	**0.478**	**0.439**	**0.433**
(£100/week)	(0.126)	(0.124)	(0.125)	(0.175)	(0.194)	(0.192)
Equiv. stepfather's	-0.047	-0.032	-0.019	**0.078**	0.058	0.054
income (£100/week)	(0.034)	(0.035)	(0.035)	(0.036)	(0.037)	(0.037)
Equiv. other (mother's)	-0.034	0.005	0.033	0.081	0.041	0.026
income (£100/week)	(0.047)	(0.039)	(0.033)	(0.051)	(0.044)	(0.042)
Child Characteristics	No	Yes	Yes	No	Yes	Yes
Mother characteristics	No	No	Yes	No	No	Yes
N	501	492	492	501	501	501
χ^2(d.f.)	11.07	36.04	45.86	14.19	36.38	43.35
	(3)	(8)	(10)	(3)	(9)	(11)
Log likelihood	-290.58	-272.99	-263.98	-323.92	-309.27	-306.66

Notes: Child characteristics include gender, number of dependent children, dummy for twins, dummies for order of births. Mother characteristics include age left full-time education and dummy for being non-white. Standard errors in parentheses. Bold and italic figures indicate statistical significance at the 5% and 10% level respectively.

Table 3A First Stage IV Results, N=501

Endogenous variable	**CS Receipts**
Equiv. stepfather's income (£100/week)	-0.0038
	(0.0154)
Equiv. other (mother's) income (£100/week)	-0.0027
	(0.0171)
Log mother's age at 1st birth	**0.5862**
	(0.0866)
First marriage	0.0374
	(0.0232)
Number of years mother in full-time employment before birth of child	**-0.0077**
	(0.0034)
Shea's Partial R^2	0.1052
F-statistic of joint significance of instruments	19.09
P-value (joint significance of IVs)	**0.0000**

Note: Standard errors in parentheses. Bold and italic figures indicate statistical significance at the 5% and 10% level respectively. All reported variables are excluded instruments except for the two income measures which are treated as exogenous.

Table 3B IV Estimates Second Stages, N=501

Outcomes	**Leaving School at 16**	**5+ GCSEs**
Equiv. CS Receipt (£100/week)	**-0.633**	**1.349**
	(0.252)	(0.297)
Equiv. stepfather's income (£100/week)	-0.018	0.050
	(0.029)	(0.034)
Equiv. other (mother's) income (£100/week)	0.040	0.013
	(0.032)	(0.038)
Anderson canon corr LR statistic Chi-sq (df)	52.726 (3)	
P-value	**0.000**	
Cragg-Donald Wald F Stat	**19.09**	
(Stock-Yogo 5% maximal IV relative bias)	(13.91)	
Sargan stat	1.414	1.936
Chi-sq (df)	(2)	(2)
P-value	0.493	0.380

Note: Other regressors include child and mother characteristics as outlined in Table 2. Standard errors in parentheses. Bold and italic figures indicate statistical significance at the 5% and 10% level respectively.

Table 4: CS Payments: Estimates from Selection Model with Full Maximum Likelihood, N=501

Outcomes	Paying any CS	Amount CS Paid (£100/week)
Equiv. stepfather's income (£100/week)	**0.201**	-0.029
	(0.094)	(0.024)
Equiv. other (mother's) income (£100/week)	0.084	0.001
	(0.101)	(0.026)
Log mother's age at 1st birth	-	**0.734**
		(0.147)
First marriage	**0.472**	-
	(0.128)	
Number of years mother in full-time employment before birth of child	-	**-0.012**
		(0.005)
ρ	-0.165	
	(0.194)	
λ	-0.045	
	(0.055)	
lnL	-351.03	

Note: Standard errors in parentheses. Bold and italic figures indicate statistical significance at the 5% and 10% level respectively. All reported variables are excluded instruments except for the two income measures which are treated as exogenous.

Table 5: Two-stage Estimation, Second Stage, Standard Errors Bootstapped with 1000 repetitions, N=501

Outcomes	**Leaving School at 16**	**5+ GCSEs**
Equiv. CS Receipt (£100/week)	**-0.561**	**1.103**
	(0.231)	(0.250)
Equiv. stepfather's income (£100/week)	-0.034	**0.079**
	(0.027)	(0.031)
Equiv. other (mother's) income (£100/week)	0.042	0.011
	(0.032)	(0.045)

Note: Other regressors include child and mother characteristics as outlined in Table 2. Standard errors in parentheses. Bold and italic figures indicate statistical significance at the 5% and 10% level respectively.

Table 6 IV Estimates Second Stages, Binary measure of CS receipt, N=501

Outcomes	**Leaving School at 16**	**5+ GCSEs**
Ever received CS	**-0.340**	**0.778**
	(0.146)	(0.182)
Equiv. stepfather's income (£100/week)	0.010	-0.014
	(0.032)	(0.039)
Equiv. other (mother's) income (£100/week)	0.038	0.016
	(0.033)	(0.041)
Anderson canon corr LR statistic Chi-sq (df)	39.681 (3)	
P-value	**0.000**	
Cragg-Donald Wald F Stat	**13.96**	
(Stock-Yogo 5% maximal IV relative bias)	(13.91)	
Sargan stat Chi-sq (df)	1.981 (2)	1.092 (2)
P-value	0.371	0.579

Note: Other regressors include child and mother characteristics as outlined in Table 2. Standard errors in parentheses. Bold and italic figures indicate statistical significance at the 5% and 10% level respectively.

Table 7A IV Estimates Second Stages, Below Median Income, N=251

Outcomes	Leaving School at 16	5+ GCSEs
Equiv. CS Receipt (£100/week)	-0.851	**1.422**
	(0.609)	(0.604)
Equiv. stepfather's income (£100/week)	-0.229	*0.333*
	(0.185)	(0.184)
Equiv. other (mother's) income (£100/week)	-0.167	**0.452**
	(0.193)	(0.191)
Anderson canon corr LR statistic Chi-sq (df)	46.493 (3)	
P-value	**0.000**	
Cragg-Donald Wald F Stat	**17.96**	
(Stock-Yogo 5% maximal IV relative bias)	(13.91)	
Sargan stat Chi-sq (df)	5.032 (2)	4.289 (2)
P-value	0.081	0.117

Note: Other regressors include child and mother characteristics as outlined in Table 2. Standard errors in parentheses. Bold and italic figures indicate statistical significance at the 5% and 10% level respectively.

Table 7B IV Estimates Second Stages), Above Median Income, N=250

Outcomes	Leaving School at 16	5+ GCSEs
Equiv. CS Receipt (£100/week)	*-0.455*	**1.282**
	(0.283)	(0.386)
Equiv. stepfather's income (£100/week)	0.012	0.031
	(0.037)	(0.050)
Equiv. other (mother's) income (£100/week)	0.048	-0.029
	(0.039)	(0.047)
Anderson canon corr LR statistic Chi-sq (df)	22.109 (3)	
P-value	**0.000**	
Cragg-Donald Wald F Stat	7.632	
(Stock-Yogo 20% maximal IV relative bias)	(6.46)	
Sargan stat Chi-sq (df)	0.431 (2)	1.424 (2)
P-value	0.806	0.491

Note: Other regressors include child and mother characteristics as outlined in Table 2. Standard errors in parentheses. Bold and italic figures indicate statistical significance at the 5% and 10% level respectively.

Table 8 IV Estimates Second Stages, Separated Mothers Only, N=395

Outcomes	Leaving School at 16	5+ GCSEs
Equiv. CS Receipt (£100/week)	**-0.496**	**1.318**
	(0.235)	(0.271)
Equiv. stepfather's income (£100/week)	-0.027	0.053
	(0.030)	(0.035)
Equiv. other (mother's) income (£100/week)	0.027	0.001
	(0.033)	(0.038)
Anderson canon corr LR statistic Chi-sq (df)	55.712 (3)	
P-value	**0.000**	
Cragg-Donald Wald F Stat	**20.85**	
(Stock-Yogo 5% maximal IV relative size)	(13.91)	
Sargan stat Chi-sq (df)	2.128 (2)	0.543 (2)
P-value	0.345	0.762

Note: Other regressors include child and mother characteristics as outlined in Table 2. Bold and italic figures indicate statistical significance at the 5% and 10% level respectively.