

# Marital Disruption and Economic Well-being: A Comparative Analysis

Arnstein Aassve\* - Gianni Betti\*\* - Stefano Mazzuco\*\*\* - Letizia Mencarini\*\*\*\*

\*Corresponding author, Institute for Social and Economic Research, University of Essex,  
Address: Wivenhoe Park, Colchester, Essex CO4 3SQ UK  
Phone ++441206873151, [aaassve@essex.ac.uk](mailto:aaassve@essex.ac.uk)

\*\*Department of Quantitative Methods  
Address: University of Siena, Italy, P.zza S. Francesco, 8 - 53100 - Siena  
Phone ++390577235084 Fax +39/0577/232626, [beti2@unisi.it](mailto:beti2@unisi.it)

\*\*\*Department of Statistics, University of Padua, Italy  
Address: Department of Statistics, Via Cesare Battisti, 241, 35121, Padua  
Phone ++390498274192, Fax ++390498274170, [mazzuco@stat.unipd.it](mailto:mazzuco@stat.unipd.it)

\*\*\*\* Faculty of Political Science, University of Florence, Italy.  
Address: Department of Statistics, viale Morgagni 59, 50134 Florence,  
Phone ++390554237268, Fax ++390554223265, [mencarin@ds.unifi.it](mailto:mencarin@ds.unifi.it)

## Abstract

Though there is a considerable literature concerned with the economic consequences of marital breakdown, there is still substantial disagreement in terms of its magnitude. One of the major problems underlying this debate is how economic well-being is defined. In this work we implement several measures of well-being of monetary and multidimensional nature using data from the European Community Household Panel. Another issue in this literature concerns selection bias of divorcing couples. We tackle this issue using a propensity score matching technique combined with a Difference-in-Differences estimator. Results confirm the importance of well-being definition. We find a strong gender bias when using monetary measures but a considerably lower one, and for some countries non existent, when using non-monetary indices.

[Word count: 115]

Keywords: marital disruption, poverty, deprivation indices, relative income, ECHP, propensity score matching

# 1. Introduction

Household structures across Europe are changing and evolving. A particular feature of modern family patterns is the significant increase in marital breakdowns. As a result the number of children living in single parent households, most of which are female-headed, has also increased. Though the issue of divorce and marital breakdowns is not new in most countries, it is an issue of continuing concern. Most of the debate around the economic consequences of divorce is focussed on gender inequalities, and the most consistent finding from the literature is a rather sharp gender difference in terms of financial outcomes following a marital disruption. Early longitudinal research from the US and Europe showed that women experiencing a divorce tend to suffer a substantial loss of income, whereas men's economic circumstances seem rather unaffected or even improve slightly in some cases (Burkhauser *et al.*, 1991; Fritzell, 1990; Jarvis and Jenkins, 1999; Manting and Bouman 2004, Poortman, 2000; Smock 1993, 1994). The reasons behind this pattern are many. One is that women tend to have lower labour market attachment, and therefore face lower earnings. Another reason is that children tend to stay with the mother following a divorce, in many cases imposing a major strain on the single female-headed household. Finally, lack of state support is another reason for why many divorced women suffer financially.

An equally consistent finding is strong country differences in terms of the economic penalty associated with a marital dissolution (Burkhauser *et al.*, 1991; Duncan and Hoffman, 1985; Finnie, 1993; Fritzell, 1990; Jarvis and Jenkins, 1999; Smock, 1993, 1994; Smock *et al.*, 1999; Poortman, 2000). The general pattern is that divorced women in Scandinavian countries, with their generous welfare provision, are much better off than divorced women in the UK, a country characterised by poorer welfare provision. Andreß *et al.* (2004) comparing Belgium, Germany, Italy, UK, and Sweden analysing the three main providers of individual welfare: 1) the family, 2) the market and 3) the state, show that the configuration of these providers to a large extent determines the economic outcome of marital dissolution. It is shown that due to limited welfare provision UK mothers are particularly vulnerable and considerably more dependent on the labour market as a means of maintaining a reasonable level of economic self sufficiency. As expected the UK setting is quite different to Scandinavian countries, but also different with respect to Continental countries such as Germany. The social democratic welfare regime is not only generous in terms of levels, but also provides strong support in terms of extensive childcare infrastructure, a system which enables Swedish mothers to work full-time to a much greater extent than other European countries, and especially the UK. However, there is no clear consensus on these findings,

especially concerning the issue of gender differences. Many maintain that the gender bias is overestimated and that the actual trend constitutes an increasing number of men who are subject to economic strain following separation (McManus and DiPrete, 2001). Indeed, there are many reasons to believe that men also experience economic problems following separation: alimony payments and the need to find another dwelling (usually the conjugal house is assigned to the woman especially if there are children) may relevantly and negatively alter the lifestyle of divorced men. Thus it may seem hard to believe that men are better-off after marital dissolution.

One of the key problems underlying this debate is the definition and measurement of the rather vague concept of “economic well-being”. Many use income or poverty status as an overall indicator of economic well-being, but these measures suffer from many drawbacks. Poverty status as a measure of well-being is criticised because it divides the population into a simple poor/non poor dichotomy, based on sometimes arbitrarily chosen thresholds (Cheli and Lemmi, 1995). Of course, the dichotomy is easily overcome by using income as a measure of economic well-being. However, this measure is problematic as it is difficult to assess to what extent an income loss brings about a real drop in living standards, especially in a comparative perspective. Moreover both income and poverty status are only monetary measures of well-being, whereas it is well recognised that well-being itself has many more dimensions, often non monetary in nature (Atkinson, 2003; Bourguignon and Chakravarty, 2003). Another drawback is that poverty status and income depend on the choice of equivalence scale, which is essentially an adjustment for household composition, acknowledging that within a household there are economies of scale in expenditure. Given that a marital breakdown inevitably modifies the household composition, the equivalence scale becomes of great consequence. However, it is not clear which equivalence to use, especially in comparative analysis. Thus, it is beneficial to consider measures of well-being in which the use of equivalence scales is not imperative. Differently from previous studies, we analyse here the effect of marital separation on economic well-being using a range of different measures. We show that the estimates and conclusions differ depending on which measure is being applied.

Another key issue in assessing the role of marital dissolution on economic well-being concerns selection bias. This is driven by the fact that couples experiencing a marital separation may be qualitatively different from couples not doing so. For example, women who are strongly dependent on their partner’s income might be less likely to separate from them as they are aware of the strong economic distress they would experience were they to split from their partner (Becker, 1991). One way to tackle this issue is to implement a *propensity score matching* (PSM) technique which nets out the impact of separation from the

confounding effects driven by other observed covariates. Obviously, many other *unobserved* covariates may influence the estimate of the effect of marital dissolution. As a result we combine the propensity score matching approach with a Difference-in-Differences estimator (DD-PSM) as suggested by Heckman *et al.* (1998). In this way we control for the effect caused by unobserved variables, provided these are time-invariant. Heckman *et al.* (1998) show that the assumption underlying the DD-PSM estimator is less restrictive than the simple Conditional Independence Assumption (CIA) on which PSM estimates rely, but of course, in so far as unobserved factors vary over time, the approach will not fully eliminate such bias.

The analysis is implemented using data from the European Community Household Panel (ECHP), which offers a unique scope for comparability at the European level (Eurostat, 1999). Uunk (2004) shows that welfare state arrangements tend to influence the economic consequences of divorce for women. Income-related arrangements alleviate the economic strains most, followed by employment-related arrangements. His findings underpin the importance of welfare regimes, and shows that differences in terms of economic strains associated with divorce, are not simply a result of country differences. Taking advantage of his work we also analyse the consequence of marital disruption under different welfare regimes, using the well known country classification of Esping-Andersen (1990; 1999). The analysis provides information about the possible effects of different family policies in European countries, with respect to consequences associated with marital disruption. Finally, we recognise the importance of the presence of children in the couple, so we make separate estimates for couples with children only. These estimates are compared to the cases where we include couples with *and* without children.

The paper is organised as follows: section 2 explains how we measure economic well-being, section 3 give details of data and estimation strategy, section 4 presents the results and section 5 concludes.

## **2. Well-being definition and measurement**

### **2.1 Measuring well-being: the conventional approach**

A simple approach to measuring an individual's well-being is to consider whether he or she is poor or not. The poverty threshold is normally defined as 60 percent of median net household equivalised income. Individuals within the household are deemed poor if the income falls below this threshold. Poverty is consequently a relative measure, and takes into account the individual's position in the income distribution relative to others within his or her own

country. An important feature of this approach is that it overcomes the fact that countries will differ in terms of per capita incomes and their purchasing power parity. A drawback, however, is that it is not clear what constitutes an appropriate poverty threshold. Often 60% of median net equivalised household income is chosen, but many use alternative poverty thresholds of 50 and 70 percent (Whelan *et al* 2001).

When assessing *economic* well-being, any measure of household income must be adjusted to reflect the needs of the people living within the household. Larger households need more income than smaller households to attain the same standard of living; adults have different needs than children. Additionally, there are economies of scale, meaning (for example) that two adults can live together more inexpensively than they could if living separately. Correction for household composition is conventionally done by calculating an equivalence scale, which is a number reflecting the needs of the household. The adjustment is done by dividing total household income by this equivalence scale. We apply the commonly used OECD modified equivalence scale, which gives a weight of one for the first adult, 0.5 for other adults than the household head, and 0.3 for children. The original OECD scale was based on empirical studies during the eighties and proved a “common” and simple scale for western countries. In this case the first adult was given a weight of 1, other adults 0.7, and children given a weight of 0.5. Hagenaars et al. (1994) revised this scale by means of new empirical findings, noting that the original OECD scale did not properly take into account economies of scales. For this reason the "modified-OECD" scale was proposed and officially adopted by Eurostat as the common scale in the ECHP.

## **2.2 Well-being as a matter of degree: the relative income measure**

Dividing the population into a simple dichotomy of “poor” and “non-poor” is clearly unsatisfactory. An individual’s well-being is not a single attribute that characterises an individual or household in terms of its presence or absence (Cheli and Lemmi, 1995). Instead we propose a measure treating poverty as a matter of degree: in principle all individuals are subject to poverty, but to varying levels. That level, say 1 for the poorest to 0 for the richest, is determined by the individual's rank in the income distribution, and the individual's share in the total income received by the population. The state of poverty is thus seen in the form of “fuzzy sets” to which all members of the population belong, but to varying degrees (membership functions). A number of authors have used the concepts of fuzzy sets in the analysis of poverty and living conditions (Lemmi and Betti, 2006) and the present contribution represents a continuation and further development of the work of Cerioli and Zani (1990), Cheli and Lemmi (1995) and Betti and Verma (1999).

There are several advantages of treating poverty in this way. Most important is that it utilises the whole distribution directly as a measure of economic well-being, as opposed to dividing the population by a dichotomous category and hence avoids the specification of a poverty line. Equally important is the potential of this approach in studying poverty (or more generally, deprivation in multiple dimensions) in the longitudinal context. The conventional approach measures mobility simply in terms of movements across some designated poverty line, and does not reflect the actual magnitude of the changes affecting individuals at all points in the distribution. Consequently, the degree of mobility of persons near to the chosen line tends to be over-emphasised, while that of persons far from that line is largely ignored (Verma and Betti, 2005).

The degree of income poverty associated with each individual as specified below was first proposed by Betti and Verma (1999), who termed it as “Fuzzy Monetary” (FM). The approach was later officially adopted by Eurostat (Giorgi and Verma, 2002). The approach can be explained as follows. Let  $y_i$  be the net equivalised household income associated with individual  $i$ . An income index is defined as the sum of the incomes of individuals richer than the individual concerned (numerator), divided by the sum of the incomes of all  $n$  individuals in the sample (denominator):

$$V_i = \frac{\sum_j y_j | y_j > y_i}{\sum_{j=1}^n y_j} \quad (1)$$

This index can be seen as the share of the total income that belongs to individuals richer than the individual concerned. It is easy to see that this index is (almost) 1 for the individual with the lowest income and equals 0 for the one with the highest income. The degree of income poverty ( $FM_i$ ) is defined as a monotonically increasing function of the income index  $V_i$  defined as:

$$FM_i = (V_i)^{\alpha/H_c} \quad (1b)$$

where  $H_c$  is the proportion of individuals who fall below the country poverty threshold, defined as 60% of median net equivalised household income. The parameter  $\alpha$  is determined such that for the European population as a whole, the mean of the index  $FM_i$  is equal to the proportion of poor individuals. The latter is imposed to better facilitate comparison with the conventional approach.

### 2.3 A multi-dimensional and comparative perspective: the deprivation index measures

The relative income measure  $FM_i$  overcomes the simplistic poor and non-poor categorisation of the population. However relative income considers deprivation only in its monetary dimension, disregarding other non-monetary aspects which may be important for individual well-being. This calls for a measure which considers deprivation in its multiple dimensions (Tsui, 1985). Certainly, in our application of consequences of marital disruption, we expect that individuals' experiences of well-being go beyond a simple drop of income: some can experience a dramatic rise in monthly expenses (for example alimony payments) with a substantial change of life-styles. Moreover, a marital disruption is likely to change, sometimes dramatically, the housing situation of the individuals involved.

Just as in the FM approach described above, we define here the concept of multiple deprivation as a matter of degree. In doing so we select a list of items indicating non-monetary deprivation in the households (see the Appendix 1). These items often take the form of simple "yes/no" dichotomies (such as the presence or absence of enforced lack of certain goods or facilities), whereas other items may involve more than two ordered categories, reflecting different *degrees* of deprivation.

At the first step these items are grouped into five different dimensions of deprivation. Thus we want to analyse not only an overall deprivation index but also the deprivation indices for each dimension of well-being. Approaches of this kind applied to poverty analysis of European countries are becoming increasingly common (Eurostat 2000; Aassve et al 2005). The five dimensions  $\delta = 1, 2, \dots, 5$  are identified from factor analysis and are as follows: (1) basic non-monetary deprivation; (2) secondary non-monetary deprivation; (3) lack of housing facilities; (4) housing deterioration; and (5) environmental problems (see Whelan *et al.* 2001 for details). The second step consists of creating a deprivation score for every item. Consider the general case of item  $k$  with  $m=1$  to  $M$  ordered categories, with  $m=1$  representing the most deprived and  $m=M$  the least deprived situation. Let  $m_{ik}$  be the category to which individual  $i$  belongs with respect to item  $k$ . As in Cerioli and Zani (1990) we assume that the rank of the categories represents an equally-spaced metric variable, and adopt the deprivation score:

$$d_{ik} = \frac{M_k - m_{ik}}{M_k - 1}, \quad 1 \leq m_{ik} \leq M_k \quad (2)$$

The third step involves determining weights to be assigned to each item of the deprivation index. The weighting procedure we propose here is a variant of the procedure developed by Betti and Verma (1999) and incorporates crucial dimensions of how the items are distributed in the population. Firstly, the weight is determined by the item's power to differentiate among individuals in the population, that is, by its dispersion. We take this into

account by letting the weight be directly proportional to the coefficient of variation of deprivation score  $d_{ik}$ . Thus items that affect only small proportions of the population – which can be expected to be considered more critical for the affected individuals (Filippone et al, 2001) - are given a larger weight. Secondly, in order to avoid redundancy, it is necessary to limit the influence of those characteristics that are highly correlated with the others within each index for the five dimensions. This means that the weight of item  $k$  in deprivation dimension  $\delta$  is taken as the inverse of an average measure of its correlation with all the variables in that dimension. There are many examples where items within a dimension can be correlated. One is the two items relating to possession of a television and a video recorder. It is unlikely that a household will possess a video recorder unless they possess a television set as well, thus inducing a positive correlation. Similarly, different items describing the conditions of the dwelling may also be correlated. For instance, a dwelling plagued by rot in window frames or floors is also more likely to report damp walls, floors and foundations (see Appendix 1 for a detailed description of the items). However, a household reporting both items should not be counted as being two times worse off than a household reporting none of these items. The final weight is proportional to the product of the two factors: the coefficient of variation of the deprivation score, and the inverse of the average of the correlations.

In the fourth step we define, for each dimension and for each individual  $i$ , the deprivation score  $S_{\delta,i}$  only considering the items belonging to dimension  $\delta$ :

$$S_{\delta,i} = \frac{\sum_k w_k (1 - d_{ik})}{\sum_k w_k} \quad (3)$$

where  $w_k$  are the weights defined above in the third step. Note that (3) defines a “positive” score indicating *lack* of deprivation. We can also consider an overall deprivation score which is a simple average of the dimension-specific scores define above:

$$S_i = \frac{1}{5} \sum_{\delta=1}^5 S_{\delta,i} \quad (4)$$

The final step is to create the non-monetary indicators of deprivation. As in the monetary approach, we define the individual’s degree to non-monetary deprivation (FS) as the *share* of the total "non-deprivation" assigned to all individuals less deprived than the person concerned. It varies from 1 for the most deprived, to 0 for the least deprived individual:

$$FS_i = \frac{\sum_j S_j \mid S_j > S_i}{\sum_{j=1}^n S_j} \quad (5)$$

The same formulation is applied within each of the five dimensions in order to derive the corresponding degrees of deprivation.

### 3. Data and estimation strategy

#### 3.1 Data and definition of marital breakdown

The European Community Household Panel (ECHP) is a set of comparable large-scale longitudinal studies implemented by the European Union. The first wave of the ECHP was collected in 1994 for the original countries in the survey: Germany, Denmark, the Netherlands, Belgium, Luxembourg, France, UK, Ireland, Italy, Greece, Spain and Portugal. Three countries were late joiners to the project: Austria joined in 1995, Finland in 1996 and Sweden in 1997. All countries except Luxembourg, Sweden and Germany are included in the analysis; Luxembourg is omitted because of small sample size, Sweden because the data do not form a panel and Germany is dropped because the information necessary to construct the deprivation indices is not available. Eight waves of the ECHP were collected in total, the last collected in 2001. We aggregate data according the welfare regime clusters defined by Esping-Andersen (1990, 1999) and Trifiletti (1999). The clusters are as follows: *Liberal countries* (United Kingdom and Ireland), *Social Democratic countries* (Finland and Denmark), *Conservative countries* (Belgium, Netherlands, France, and Austria), and *Mediterranean countries* (Italy, Spain, Portugal, and Greece).

The event of interest is marital dissolution that is defined by separation or a divorce. Normally a marital dissolution is initiated by a separation and followed by a formal divorce. The ECHP reports self reported marital status where separation and divorce are mutually exclusive categories. However, laws and regulations on separation and divorce vary across European countries. Thus, the duration between separation and the divorce will not only differ between individuals within a country, they will also differ systematically between countries. A problem arises from the fact that individuals might put different meaning to the term separation. In general, one would expect that a separation implies that the couple move to separate dwellings. However, in a significant amount of cases a *physical* separation does not take place until the couple is formally divorced. This means that many couples report separation despite living in the same dwelling, sometimes for several time periods. Consequently it is problematic to use the self reported marital status to define the time when a marital split occurred.

In this analysis a couple is not formally recorded as separated unless they also reported that they lived in separate households. We make this distinction since in this situation they cannot benefit from economies of scale of the household, nor can they share the burdens of rearing children. In estimating the impact of marital split on poverty status we exclude those couples who are already poor before the split. By 'poor' we mean those individuals whose equivalent household income is below the poverty line. Consequently the samples differ according to the poverty line used. By using propensity score matching (to be explained in next section) we estimate the differential *poverty entry rate* among separated and not separated individuals, i.e. the difference of the percentages of individuals entering poverty in the two groups. Including also those defined as poor prior to divorce will make interpretation more difficult since this would potentially mix individuals entering poverty with those exiting poverty. Note however, that when estimating the effect of marital split on relative income or on deprivation indices, the *complete* sample is used.

### ***3.2 Propensity score matching***

In estimating the effect of marital disruption on economic well-being, we face the potential problem of selection bias. That is, couples experiencing a marital separation may be qualitatively different from couples not separating. For example, women who are strongly dependent on partner's income are probably less likely to separate as they are aware of the strong economic distress they would experience in the case they split from their partner (Becker, 1991). Here we tackle this issue by implementing a propensity score matching (PSM) technique. Applications of this kind are growing in literature (see, among others, Blundell *et al.* 2005; Lechner, 2002; Dehejia and Wahba, 1999).

In our setting we assume that each individual  $i$  has two potential outcomes,  $Y_{1i}$  in the case he or she experiences a marital split (the treatment) and  $Y_{0i}$  in the case he/she does not (the controls). In the case of a marital split  $Y_{0i}$  or conversely  $Y_{1i}$  in the case of no split, is also referred to as the 'counterfactual'. The causal impact is given by the comparison between  $Y_{1i}$  and  $Y_{0i}$ . Obviously, only one of these two outcomes is observable for each individual making a direct comparison impossible, a problem often referred to as the "fundamental problem of causal inference" (Holland, 1986).

Let  $D_i$  be the treatment variable taking the value 1 if individual  $i$  receives the treatment (marital split) and 0 otherwise. One characteristic of interest is referred to as the *average treatment effect on treated* (ATET) and is expressed as:

$$\text{ATET} = E(Y_{1i} - Y_{0i} | D_i = 1) = E(Y_{1i} | D_i = 1) - E(Y_{0i} | D_i = 1) \quad (6)$$

In order to estimate ATET we need to identify  $E(Y_{0i} | D_i = 1)$ , which can be done by imposing assumptions on the selection process. An easy solution is to use a naïve estimator of ATET consisting of observed difference between treated and control groups, which assumes

$$\text{ATET} = E(Y_{1i} | D_i = 1) - E(Y_{0i} | D_i = 0) \quad (7)$$

But (7) assumes that there is *no* selection bias, which means that the group of treated is randomly selected from the total population so that in all other relevant respects apart from receiving the treatment the two groups may be regarded as comparable. It is well known that in observational studies this assumption is overly strong and treated and control groups are systematically different, implying that an estimate based on (7) would be a biased for true ATET. It is important to understand the nature of the bias that arises. Heckman *et al.* (1998) propose writing the bias  $B$  as a function of a set of pre-treatment observed covariates  $X$ :

$$B = \int_{S_{1X}} E(Y_0 | X, D = 1) dF(X | D = 1) - \int_{S_{0X}} E(Y_0 | X, D = 0) dF(X | D = 0) \quad (8)$$

where  $S_{1X}$  and  $S_{0X}$  are the supports of  $X$  for  $D=1$  and  $D=0$  respectively. These are the sets of values of  $X$  we observe for the treated group ( $D=1$ ) and for the control group ( $D=0$ ). Based on (8) Heckman *et al.* (1998) derive a decomposition of  $B$  into three terms they refer to as  $B_1$ ,  $B_2$ , and  $B_3$ .  $B_1$  arises when the supports of the observable  $X$  for the treated and the control group  $S_{1X}$  and  $S_{0X}$  are not overlapping, i.e. among the treated group we observe values of  $X$  that are not observed in the control group and vice versa. This implies that for treated individuals whose values of  $X$  lies out of the common area of  $S_{1X}$  and  $S_{0X}$  (the *common support*) we are unable to find an equivalent (i.e. same values of  $X$ ) individual in the control group to match with. Term  $B_2$  depends on mis-weighting of  $E(Y_{0i} | D_i = 0)$  in the common support. It arises when the distribution of  $X$  is different between the treatment and the control group. Finally, term  $B_3$  refers to the bias that arises if the distribution of *unobserved variables* is different between the treated and untreated (see Heckman *et al.* 1998, for a more detailed discussion on  $B_1$ ,  $B_2$ , and  $B_3$ ). The way in which the biases in (8) are eliminated by our proposed matching procedure is explained below.

The matching method is based on the critical assumption termed *conditional independence assumption* (CIA) stating that treatment status is random conditional on a given set of  $X$ . The CIA is formally expressed as:

$$Y_0 \perp D \mid X \quad (9)$$

and means that conditional on  $X$  the potential outcome in case of non-treatment (i.e.  $Y_0$ ) is independent of the treatment status. Whereas (9) impose full independence, identification of ATET requires a less restrictive condition. As Smith and Todd (2005) point out it suffices that we impose *mean* independence, i.e.  $E(Y_{0i} \mid X_i, D_i=0) = E(Y_{0i} \mid X_i, D_i=1)$ . Thus the ATET can be written as:

$$\begin{aligned} \text{ATET} &= E(Y_{1i} - Y_{0i} \mid D_i = 1) = E_{X \mid D=1} \{E(Y_{1i} - Y_{0i} \mid X_i, D_i=1)\} = \\ &E_{X \mid D=1} \{E(Y_{1i} \mid X_i, D_i=1) - E_{X \mid D=1} \{E(Y_{0i} \mid X_i, D_i=1)\}\} = \\ &E_{X \mid D=1} \{E(Y_{1i} \mid X_i, D_i=1) - E_{X \mid D=1} \{E(Y_{0i} \mid X_i, D_i=0)\}\}. \end{aligned} \quad (10)$$

The first two lines in (10) expand the ATET defined in (6), whereas in the third line we apply CIA in its less restrictive form by substituting the second (unidentified) term  $E_{X \mid D=1} \{E(Y_{0i} \mid X_i, D_i = 1)\}$  by  $E_{X \mid D=1} \{E(Y_{0i} \mid X_i, D_i = 0)\}$ . In (10) the ATET is now fully identified, and is a direct consequence of CIA. Though theoretically appealing, the matching approach is in practice difficult to apply when the dimension of  $X$  is high because of the computational difficulties in estimating the conditional expectations in (10). Instead of matching on the basis of  $X$  one can equivalently match the treated units to control comparison units on the basis of a balancing score. A particular form of this is a ‘‘propensity score’’ (Rosenbaum and Rubin 1983), which is the conditional probability of receiving the treatment given the values of  $X$ :  $p(X_i) = \Pr(D_i = 1 \mid X_i)$ . The propensity score is usually estimated with either a logit or a probit model. This result reduces the dimensionality problem of computing the conditional expectation, as we now only need to condition on a one dimensional variable (i.e. the propensity score) and ATET can be written as:

$$\text{ATET} = E_{p(X_i)} \{[E(Y_{1i} \mid D_i = 1, p(X_i)) - E(Y_{0i} \mid D_i = 0, p(X_i))]\} \quad (11)$$

Several matching procedures can be used to estimate (11) (see, for example Becker and Ichino, 2002; and Smith and Todd, 2005), but all of them can be seen as generated by the following formula:

$$\hat{\text{ATET}} = \frac{1}{n_1} \sum_{i \in \{D_i=1\}} \left[ Y_{1i} - \sum_{j \in \{D_j=0\}} w_{ij} \cdot Y_{0j} \right] \quad (12)$$

where the weight  $w_{ij}$  is defined according to the matching method used and  $n_1$  is the number of treated individuals. Here we implement a *nearest neighbour* matching consisting of pairing every treated unit with the closest control unit in terms of their propensity score. Thus, a treated unit  $i$  is paired with the control unit  $j$  that gives the smallest value of  $|p(X_i) - p(X_j)|$ , meaning that for every  $i$  the weight  $w_{ij}$  is one for unit  $j$  that is closest, otherwise the weight is zero (see Smith and Todd, 2005 and Caliendo and Kopeinig, 2005). Of course in some cases  $k$  control units (with  $k > 1$ ) may satisfy the matching rule (i.e. there are more than one control with the minimum distance from the treated). If so we use all these  $k$  controls with weight  $1/k$ .

All three sources of bias in (8) are now eliminated.  $B_1$  is eliminated by allowing matches only in the common support region, i.e. treated and control units whose values of  $X$  lie outside the common area of  $S_{1X}$  and  $S_{0X}$  are ruled out from estimation of ATET.  $B_2$  is eliminated because the control units are re-weighted according to the value of  $p(X)$ , leading to balance of  $X$  between treated and control units.  $B_3$  is the only component of (8) that is not eliminated by matching but is assumed to be zero by CIA.

The matching procedure described above can be implemented on cross-sectional observations by recording both treated (i.e. those who experienced a marital split) and the controls (those who did not split). However, our data source is longitudinal which means that measurements of the outcomes of interest are available both before and after treatment. This is a highly useful feature since we are able to compare the mean change of well-being from one time period  $t$  to another,  $t+1$ , of treated, with the mean change of well-being for the same time period for controls. It means that we can define a Difference-in-Difference (DD) estimator as follows:

$$DD = E(Y_{1i}^{t+1} - Y_{1i}^t) - E(Y_{0i}^{t+1} - Y_{0i}^t) = E(\Delta_{1i}) - E(\Delta_{0i}) \quad (13)$$

An important advantage of the *DD* estimator is that it allows us to control for selection into the treatment group caused by unobserved variables. To see this clearer we can define the point-wise bias at  $X$  at time  $t$  as  $B^t(X) = E(U_{0i}^t | X_i, D_i = 1) - E(U_{0i}^t | X_i, D_i = 0)$ , where  $U_{0i}$  is the value of unobserved variables (Heckman et al 1998). Whereas the CIA assumes that  $B_3$ , which is a weighted average of  $B(X)$ , is zero, the critical identifying assumption for the *DD* estimator is that:

$$B^{t+1}(X) - B^t(X) = 0. \quad (14)$$

In this way the CIA is relaxed since we no longer assume  $B(X)$  to be zero, rather we only assume its value does not change from wave  $t$  to wave  $t+1$ . In other words, provided unobserved heterogeneity is time-fixed, its effect will be netted out by taking first difference. As a result it has been argued that the DD-PSM estimator is more robust since it eliminates temporarily-invariant sources of bias (e.g. Dehejia and Wahba, 1999, 2002 and Smith and Todd, 2005). Of course even this assumption might be violated if some *time-varying* source of bias among the unobserved variables exists. For this reason the selection of matching variables remains a crucial part of estimation. By including all the variables correlated with both the outcome and the treatment in the model estimating the propensity score makes (14) more likely to hold. The final estimator of the impact of marital split on well-being is given by:

$$DD - PSM = E_{p(X_i)} \{ [E(\Delta_{1i} | D_i = 1, p(X_i)) - E(\Delta_{0i} | D_i = 0, p(X_i))] \} \quad (15)$$

Estimation of (15) is done using the estimator defined in (12), where the values of  $\Delta_{1i}$  and  $\Delta_{0i}$  substitute  $Y_{1i}$  and  $Y_{0i}$ . The DD-PSM estimator is implemented throughout the analysis. However when we estimate the effect of separation on poverty status, DD-PSM and cross-sectional estimators are equivalent given that all those who are poor *before* the marital split are ruled out from analysis. This means in the analysis of marital split on entering poverty  $Y^t=0$  for all individuals (i.e. non-poor individuals are not included in the sample).

In all samples the variables which are suspected to confound the effect of marital split on poverty (or deprivation) are included in the estimation of the propensity score: wave, age, number of children, well-being level prior the event (measured both in terms of income and in terms of deprivation), education and employment status (see Appendix 2). Though estimation results predicting participation into treatment might be of interest, it must be kept in mind that the main purpose of propensity score estimation is to ensure that the distribution of observed covariates is the same between treated and the matched controls. If this is the case then all covariates in  $X$  are balanced and satisfy what is termed the balancing property (Augurzky and Schmidt, 2000). Clearly this needs to be tested, since if the balancing property is not satisfied, we would potentially match units with quite different values of  $X$  despite their propensity scores being close. In this scenario we would need to correct our logit model underlying the propensity score estimation. Following Dehejia and Wahba (2002) and Smith and Todd (2005) we use t-test for equality of means for each covariate  $X$ , before and after matching. This is likely to suffice considering that almost all the variables in our application are binary. Accepting the null hypothesis means that control units are not different from the treated

except for the treatment status. Other tests have been suggested (see, for instance, Sianesi, 2004; Dehejia and Wahba, 2002) but there is no clear consensus on which is the most powerful. Becker and Ichino (2002) argue that using a t-test for equality of means with 0.05 significance level is a relatively conservative approach, especially since this level applies to each single variable used in the propensity score model. The balancing property is satisfied in all of our estimates

The estimation of standard errors of ATET is not a trivial exercise – the main problem being that the estimated variance of ATET should also include the variance due to estimation of the propensity score. The common solution to this problem is bootstrapping (see for example, Lechner, 2002 and Blundell *et al.*, 2005). This is the solution we adopt, using the module developed by Leuven and Sianesi (2003) for STATA. Reported t-values are the ratio between the estimated ATETs and the bootstrapped standard errors. For formal significance testing these ratios may be referred to as critical values using the usual standard Normal approximation.

## 4. Results

### *4.1 Entering Poverty*

Taking the sample of individuals defined as non-poor in previous period, Table 1 presents the effects of experiencing a divorce/separation event on entering poverty using different poverty thresholds. Note that the estimate refers to what is called the average treatment effect on the treated, and reflects therefore the difference between the rate of entering poverty for married couples and individuals experiencing a marital break-up. The results confirm that women are considerably more likely to enter poverty as a result of divorce compared to men. This is the case independent of countries and the poverty threshold used. Moreover, the effects are largely consistent with welfare regime theory. Especially with the 50 percent threshold, the ranking of country groups is perfectly in line with the Welfare Regime theory, the Social-Democratic group having the smallest effect followed in ascending order by the Conservative countries, the Mediterranean, and, finally, the Liberal group that presents the highest effect. However, this ranking does not remain perfectly consistent if we consider higher poverty thresholds. By using the 60 or 70 percent of median income, the effect of marital disruption increases dramatically for the Conservative and Social-Democratic

countries. In fact the Social Democratic countries reach in this case the levels of the Mediterranean group. Thus, divorce clearly affects women in Social Democratic countries as well in that they are considerably more likely to enter “mild” poverty, and they are more likely to do so than divorced women in the Conservative countries. Women in the Liberal countries clearly experience the strongest effect, independent of poverty line used. Note that the sample mainly consists of individuals from the United Kingdom, as the number of separations and divorce is rather low in Ireland. As expected the effect for men is far lower and only in the Conservative group significant (when the poverty line is 60% or 70% of median income). The Liberal countries also have the largest gender difference. This gender difference is slightly larger than Mediterranean countries. When we consider only couples with children the effect of marital disruption is even stronger: for Liberal women the rise of poverty entry rate is beyond 0.5 when the poverty threshold is set at 70% of the median income. For men the figures are not significantly different when we consider only those with children.

**Table 1: Average Treatment effect on poverty entry rate at different poverty thresholds, by gender, presence of children and welfare regime.**

	MALES				FEMALES			
	All Couples		Couples with children		All Couples		Couples with children	
	ATET	t-value	ATET	t-value	ATET	t-value	ATET	t-value
<i>Liberal Countries</i>								
50% threshold	0.030	1.250	0.020	0.623	0.335	7.328	0.365	7.262
60% threshold	0.016	0.518	0.045	1.024	0.389	8.543	0.414	7.017
70% threshold	0.000	0.000	0.011	0.250	0.432	9.086	0.509	8.755
<i>Social Democratic Countries</i>								
50% threshold	0.029	1.372	0.019	0.766	0.110	2.978	0.103	3.418
60% threshold	0.047	1.444	0.071	1.909	0.250	5.987	0.204	4.926
70% threshold	0.057	1.372	0.064	1.451	0.276	6.612	0.296	5.223
<i>Conservative Countries</i>								
50% threshold	0.009	0.819	0.009	0.570	0.123	5.401	0.147	7.058
60% threshold	0.043	2.803	0.024	1.183	0.210	7.593	0.217	8.557
70% threshold	0.065	2.574	0.039	1.555	0.227	8.066	0.242	9.353
<i>Mediterranean Countries</i>								
50% threshold	0.038	1.735	0.057	1.982	0.217	9.008	0.215	6.741
60% threshold	0.045	1.636	0.007	0.225	0.245	7.430	0.284	7.016
70% threshold	0.016	0.540	0.042	1.106	0.295	8.012	0.294	8.445

#### 4.2 Fuzzy monetary indicator

The results reported in Table 2 are the estimates of Average Treatment Effect on the fuzzy monetary indicator, namely the relative income. These estimates reflect a decline or a rise in the terms of ranking of income within a certain country. In other words, a positive effect means a decline in the income ranking due to marriage dissolution, whereas a negative effect means a rise. Therefore in Liberal countries, for instance, women tend to experience a strong decline whereas men's ranking remains approximately the same after the separation\divorce. Whereas the results for Liberal countries are consistent with the estimated poverty entry rates presented in Table 1, the situation is less straightforward for the other countries. In contrast to poverty entry rate for men, where some significant effects are noted, in general for all country groups, declines in Table 2 for men are small and statistically not significant. In Mediterranean countries declines for women are weaker than in Liberal countries but stronger than in the other two country groups. Women from Social Democratic and Conservative countries experience approximately the same decline when we consider the whole sample. However, when considering only couples with children women in the Social Democratic are lower, and lower than in the whole sample for that group. Thus for this group, the effect of divorce or separation on own income ranking is milder if they have children. Interestingly we find a reversed trend for Liberal countries: women with children experience a stronger effect compared to all women (with or without children). This is largely in line with results of divorce effect on poverty entry rate reported in Table 1. Mediterranean and Conservative countries show no relevant difference between the whole sample and women with children. This is again consistent with results in Table 1.

**Table 2: Average Treatment Effect of marital dissolution on relative income.**

	MALES				FEMALES			
	All Couples		Couples with children		All Couples		Couples with children	
	ATET	t-value	ATET	t-value	ATET	t-value	ATET	t-value
Liberal	0.014	0.624	-0.011	-0.508	0.298	9.039	0.346	8.268
Social Democratic	0.022	1.178	0.035	1.716	0.141	6.106	0.077	3.918
Conservative	-0.000	-0.031	-0.001	-0.541	0.144	9.464	0.152	7.528
Mediterranean	0.011	0.472	-0.015	-0.696	0.205	9.561	0.199	7.697

### *4.3 Deprivation indices*

We now move to the effect of marital dissolution on total household deprivation. We consider first the change in total deprivation index due to the separation from spouse. We then consider in more detail the effect of separation on the five dimensions of deprivation as

defined earlier. Here we are showing only the estimated average treatment effect on treated for the overall deprivation index, the basic lifestyle deprivation index, and the secondary lifestyle deprivation index. The estimates for the remaining indices (housing facilities, housing deterioration, and environmental problems) are omitted as in none of these cases did we find significant effects of marital separation.

### Total deprivation

The results reported in Table 3 show a somewhat different picture than the analysis of poverty entry rates. The effect for women from Liberal countries is still the highest, but now the Social Democratic and the Mediterranean groups show quite similar figures for both men and women. We find the lowest impact among the Conservative countries. Importantly, the effects are now significant also for men, and though the magnitude of the effects is always lower than women, there is less of a gender gap.

The effect for men in the Liberal group is particularly striking. Compared to poverty entry and change in relative income (Tables 1 and 2), the effect on deprivation for men is considerably stronger, and not too different from that of women. The gender bias in Liberal countries is considerably smaller in Liberal countries in so far economic wellbeing is measured in terms of overall deprivation. Men in the Conservative countries suffer a significant rise of deprivation after separation as well, but this is consistent with the figures we reported for poverty entry. As with the Liberal countries, Conservative countries now show a quite narrow gender gap. Thus by measuring well-being in terms of total deprivation the geographical pattern of gender differences changes dramatically. Now the Social Democratic and the Mediterranean countries have the largest gender differences out of the four groups. This time the effect of marital split changes somewhat when considering couples with children only: the effect for males is milder in Liberal countries and stronger in Social Democratic ones, whereas it does not change much for the other country groups. For women we observe a smaller effect of having children in Liberal countries and a higher one in the Conservative countries.

**Table 3: Average Treatment Effect of marital dissolution on deprivation index.**

	MALES				FEMALES			
	All Couples		Couples with children		All Couples		Couples with children	
	ATET	t-value	ATET	t-value	ATET	t-value	ATET	t-value
Liberal	0.124	3.100	0.093	2.203	0.138	4.166	0.123	3.094
Social Democratic	0.023	0.723	0.073	2.054	0.106	3.646	0.097	2.736
Conservative	0.041	2.688	0.044	2.341	0.058	3.682	0.075	4.795
Mediterranean	0.034	1.137	0.036	1.112	0.115	4.860	0.105	3.831

### Basic Lifestyle deprivation

If we focus on the first dimension of deprivation, i.e. deprivation on basic lifestyle, we find results shown in Table 4 relatively consistent with results for the total deprivation index. Again the Liberal group shows the strongest effect both for men and women, but this time the effect for women is about twice as high compared to men, reflecting a strong gender bias. The weakest effect is found in Mediterranean countries, and it is particularly low for men. Moreover, for Mediterranean men it does not matter much whether children are present or not. Whereas Mediterranean women suffer considerably more than their male counterparts, they are less bad off compared to women in the other groups. Again for the Social Democratic countries we notice a relatively high effect for women and a significant gender gap. Finally, we register as before a significant effect for men also in the Conservative group.

The presence of children seems to negatively influence the effect for men: apart from Mediterranean countries, almost everywhere the effect of marital split is stronger when we consider only couples with children. Conversely, the effect for women is almost everywhere weaker, with the exception of Conservative countries.

**Table 4: Average Treatment Effect of marital dissolution on basic lifestyle deprivation index.**

	MALES				FEMALES			
	All Couples		Couples with children		All Couples		Couples with children	
	ATET	t-value	ATET	t-value	ATET	t-value	ATET	t-value
Liberal	0.114	2.785	0.136	2.178	0.224	4.541	0.194	3.303
Social Democratic	0.033	0.850	0.100	2.251	0.166	3.646	0.104	2.173
Conservative	0.086	4.840	0.089	3.904	0.127	6.010	0.145	5.881
Mediterranean	0.025	0.809	0.024	0.613	0.126	4.374	0.118	3.988

### Secondary lifestyle deprivation

Finally, in Table 5 we look at the effects of marital disruption on the deprivation level concerning the secondary lifestyle deprivation. Surprisingly we find a rather strong effect for women in the Social Democratic countries and for the estimates concerning all couples they suffer more than women from the other groups, though the estimate for women in Liberal countries is at a similar level. Once we consider women with children only, we see that the Mediterranean women are worst off, but it is interesting that women from the Social Democratic group are worse off than women from both Liberal and the Conservative group. . Another interesting feature of these results is the effect of separation for men in Liberal

countries, which is now quite close to deprivation for women, i.e. the gender gap is reduced when considering secondary lifestyle deprivation. The gender gap remains, however, in the remaining groups. The role of children is particularly interesting for Liberal countries. For men there is no difference in the effect of divorce

Again, if we consider couples with children only, the results change somewhat. Surprisingly the effect for women is no longer significant whereas for men it remains unaltered in Liberal countries. A substantial drop is registered also for Social Democratic women combined with an increase for men in Social Democratic countries. Conversely we observe a small increase of the effect for females in the other two country groups. No relevant change is registered for men in these countries.

**Table 5: Average Treatment Effect of marital dissolution on secondary lifestyle deprivation index.**

	MALES				FEMALES			
	All Couples		Couples with children		All Couples		Couples with children	
	ATET	t-value	ATET	t-value	ATET	t-value	ATET	t-value
Liberal	0.149	3.311	0.148	2.734	0.147	3.067	0.077	1.647
Social Democratic	0.069	2.147	0.119	2.625	0.157	4.976	0.129	3.179
Conservative	0.052	2.750	0.046	1.987	0.086	4.840	0.109	4.938
Mediterranean	0.049	1.578	0.042	1.008	0.134	4.393	0.149	4.826

## 5 Concluding remarks

The present work is concerned with the economic consequences of marital disruption for both the members of the separating couples. Most of the literature on this topic assesses whether there is a large gender bias, with women being exposed to high poverty risks in the aftermath of separation whereas men seem not to experience any dramatic drop in their income and are sometimes even better off after divorce/separation. Some authors (McManus and DiPrete, 2001) challenge this evidence, suggesting that the gender bias is less strong than generally acknowledged, and that men also suffer economically after marital disruption. Here we suggest two issues that are essential to this debate: firstly the conventional measures of well-being (i.e. income and poverty status) are not entirely satisfying. Poverty status creates distinction between “poor” and “non poor”, but it is not clear which poverty line should be considered appropriate and why. Moreover, income and poverty status do not encapsulate all the dimensions underlying poverty and social exclusion - only the monetary one. We may expect that men are not suffering in monetary terms in the aftermath of separation but they

experience an increased deprivation in lifestyle standards all the same because of a rise in expenses due to alimony payments, new dwelling costs, etc. The second issue concerns selection. This is driven by the fact that men and women who are at high risk of entering poverty may be more likely to avoid separation. By using a propensity score matching procedure combined with a Difference-in-Differences estimator we control for such a selection bias.

We expect that by using different measures of well-being, we are able to observe that both men and women experience an economic deprivation after separation, women being more deprived in monetary terms and men in non monetary terms. The results conform largely to our expectations: it is confirmed that the definition of poverty threshold is an important issue. Results differ considerably depending on whether we use a 50%, a 60%, or 70% poverty line. Moreover when we use monetary measures (i.e. poverty status and relative income) it is unquestionable that women suffer a disproportionately larger negative effect than men. Also important is that by using monetary measures, we find that most of the results are consistent with welfare regime theory. However, the non-monetary measures (i.e. deprivation indices) provide a different picture. Women are still found to suffer significantly more than men, but it is also clear that men's level of deprivation also increases, and in some cases there is no significant difference between the ATET estimated for men and women (this is the case in Liberal countries when using the overall deprivation index and the secondary lifestyle deprivation index).

Children play an important role in explaining the gender differences. If there are children in the conjugal dwelling, then mothers are much more likely to be granted custody following a divorce. Thus the divorce event will for many women imply reduced income (poorer access to the husband's income) and a higher relative expenditure. Men on the other hand, are likely to live alone or with parents, and are much less likely to experience poverty and financial strain. Considering couples with children only in the analysis of entering poverty, we notice that in Liberal and Mediterranean countries the gender gap is even larger, in Social Democratic countries it is smaller, and in the Conservative countries it remains more or less unaltered.

However, in terms of deprivation, men do suffer significantly. Many of the items used to compute the deprivation index refer to characteristics of the dwelling. If it is the case that men normally have to leave the dwelling following a divorce, they will, in the short run at least, loose out on many of the goods and services that the household would provide. So although men are not worse off financially, they are worse off in terms of consumer durables

and certain expenditure goods. It also seems likely that the new dwelling is often of poorer quality than the original dwelling, which is consistent with our estimates.

The gender difference is clearly smaller when children are not present in the dwelling. With no children, the effect on lifestyle deprivation among men becomes higher, whereas it is slightly smaller for women. One important factor here is that it is less clear which of the spouses will remain in the conjugal dwelling if the couple has no children.

## References

- Aassve A., Mazzucco S. and Mencarini L. (2005) Childbearing and well-being: a comparative analysis of European welfare regimes. *Journal of European Social Policy*, 15, 283-299.
- Andreß H.J., Borgloh B., Bröckel M., Giesselmann M., Hummelsheim D. (2004) The economic consequences of partnership dissolution A comparative analysis of panel studies from five European countries. Paper presented at the 3rd Conference of the European Research Network on Divorce, December 2 – 4, 2004 in Cologne/Germany.
- Atkinson A.B. (2003) Multidimensional deprivation: contrasting social welfare and counting approaches. *Journal of Economic Inequality*, 1, 51-65.
- Augurzky, B. and Schmidt, C. (2001) The Propensity Score: A Means to an End, Institute for the Study of Labour (IZA) Discussion Paper number 271, March, Bonn.
- Becker G. S. (1991) *A Treatise on Family* Harvard University Press, Cambridge, Mass
- Becker S. and Ichino A. (2002) Estimation of average treatment effects based on propensity scores. *The STATA Journal*.
- Betti G. and Verma V. (1999) Measuring the degree of poverty in a dynamic and comparative context: a multi-dimensional approach using fuzzy set theory. *Proceedings, ICCS-VI*, 11, pp. 289-301, Lahore, Pakistan.
- Blundell, R., Dearden, L. and Sianesi, B. (2005) Evaluating the effect of education on earnings: Models, methods and results from the National Child Development Survey. *Journal of the Royal Statistical Society, Series A*, 168, 3, 473-512.
- Bourguignon F. and Chakravarty S.R. (2003) The measurement of multidimensional poverty. *Journal of Economic Inequality*, 1, 25-49.

- Burkhauser, R. V., Duncan, G. J., Hauser, R. and Berntsen, R. (1991) Wife or Frau, women do worse: a comparison of men and women in the United States and Germany after marital disruption. *Demography*, 28, 353-360.
- Caliendo M. Kopeinig S. (2005) Some Practical Guidance for the Implementation of Propensity Score Matching, IZA Discussion Paper n. 1588, Bonn, Germany.
- Ceroli A. and Zani S. (1990) A Fuzzy Approach to the Measurement of Poverty, in Dagum C. and Zenga M. (eds.), *Income and Wealth Distribution, Inequality and Poverty, Studies in Contemporary Economics*. Berlin: Springer Verlag, pp. 272-284.
- Cheli B. and Lemmi A. (1995) A Totally Fuzzy and Relative Approach to the Multidimensional Analysis of Poverty. *Economic Notes*, 24, pp. 115-134.
- Dehejia R, Wahba, S. (1999) Causal Effects in Nonexperimental Studies: Reevaluating the Evaluation of Training Programs. *Journal of the American Statistical Association*, 94 (448): 1053-1062.
- Dehejia R, Wahba, S. (2002) Propensity Score Matching Methods for Nonexperimental Causal Studies, *The Review of Economics and Statistics*, 84(1), 151-161.
- Duncan, G. J. and Hoffman, S. D. (1985) A reconsideration of the economic consequences of marital dissolution. *Demography* 22: 485-497.
- Esping-Andersen, G. (1990) *The Three Worlds of Welfare Capitalism*. Princeton, NJ: Princeton University Press.
- Esping-Andersen, G. (1999) *The Social Foundations of Post-industrial Economies*. Oxford: Oxford University Press.
- Eurostat. (1999) *ECHP Data Quality*, European Commission, Luxembourg, 108/99.
- Eurostat. (2002) *European Social Statistics: Income, Poverty and Social Exclusion: 2nd Report*. Luxembourg: Office for Official Publications of the European Communities.
- Filippone A., Cheli B. and D'Agostino A. (2001) Addressing the interpretation and the aggregation problem in totally fuzzy and relative poverty measures. ISER Working Paper, number 2001-22, University of Essex.
- Finnie, R. (1993) Women, men and the economic consequences of divorce: evidence from Canadian longitudinal data. *Canadian Review of Sociology and Anthropology* 30, 205-241.
- Fritzell, J. (1990) The dynamics of income distribution: economic mobility in Sweden in comparison with the United States. *Social Science Research* 19: 17-46.

- Giorgi L and Verma V. (2002) *European Social Statistics: Income, Poverty and Social Exclusion: 2nd Report*. Luxembourg: Office for Official Publications of the European Communities.
- Hagenaars A.J.M., de Vos K., Zaidi M.A. (1994) *Poverty statistics in the late 1980s: research based on micro-data*, EUROSTAT, Luxembourg.
- Heckman J.J., Ichimura H., Smith J., Todd P.E. (1998) Matching as an Econometric Evaluation Estimator: Evidence from Evaluating a Job Training Programme. *Review of Economic Studies* 64, 605-654
- Holland P.W. (1986) Statistics and Causal Inference. *Journal of the American Statistical Association* 81, 945-970.
- Jarvis S. and Jenkins S.P. (1999) Marital splits and income changes: evidence from the British Household Panel Survey. *Population Studies*, 53 (2), 237-254.
- Lechner M. (2002) Some practical issues in the evaluation of heterogeneous labour market programmes by matching methods. *Journal of the Royal Statistical Society Series A* 165(1) 59-82.
- Lemmi A. and Betti G. (2006) *Fuzzy set approach to multidimensional poverty measurement*, Springer, New York.
- Leuven, E. and Sianesi, B. (2003) PSMATCH2: Stata module to perform full Mahalanobis and propensity score matching, common support graphing, and covariate imbalance testing. University of Amsterdam and Institute for Fiscal Studies, London.
- Manting D., Bouman A. M. (2004) Short and long term economic consequences of union dissolution: The case of the Netherlands. Paper presented at the 3rd European Conference of the Research Network on divorce (Keulen, December 2004).
- McManus P.A., DiPrete T.A. (2001) Losers and Winners: The Financial Consequences of Separation and Divorce for Men. *American Sociological Review*, 66, 246-268.
- Poortman, A. (2000) Sex differences in the economic consequences of separation: a panel study of the Netherlands. *European Sociological Review* 16: 367-383.
- Rosenbaum P.T. and Rubin, D.B. (1983) The Central Role of the Propensity Score in Observational Studies for Causal Effects. *Biometrika*, 70(1), 41-55.
- Sianesi B. (2004) An Evaluation of the Active Labour Market Programmes in Sweden, *The Review of Economics and Statistics*, 86(1), 133-155.

- Smith J.A., Todd P. (2005), Does Matching Overcome Lalonde's Critique of Nonexperimental Estimators? *Journal of Econometrics*, 125(1-2), 305-353.
- Smock P.J, Manning W.D., Gupta S. (1999) The Effect of Marriage and Divorce on Women's Economic Well-Being. *American Sociological Review*, 64, 794-812.
- Smock, P. J., (1993) The economic costs of marital disruption for young women over the past two decades. *Demography* 30: 353-371.
- Smock, P. J., (1994) Gender and the short-run economic consequences of marital disruption. *Social Forces* 73, 243-262.
- Trifiletti, R. (1999) Southern European Welfare Regimes and the Worsening Position of Women. *Journal of European Social Policy*, 9(1), 49-64.
- Tsui K. (1985) Multidimensional generalisation of the relative and absolute inequality indices: the Atkinson-Kolm-Sen approach. *Journal of Economic Theory*, 67, 251-265.
- Uunk, W. (2004) The Economic Consequences of Divorce for Women in the European Union: The Impact of Welfare State Arrangements. *European Journal of Population*, 20, 251-285.
- Verma V. and Betti G. (2005) Sampling Errors for Measures of Inequality and Poverty. Invited paper in *Classification and Data Analysis 2005 - Book of Short Papers*, pp. 175-179, CLADAG, Parma, 6-8 June 2005.
- Whelan C.T., Layte R., Maitre B. and Nolan B. (2001) Income, deprivation and economic strain: an analysis of the European Community Household Panel. *European Sociological Review*, 17, 357-372.

## Appendix 1: Variables for calculating deprivation indices

### Dimensions and items of non-monetary deprivation

1 Basic non-monetary deprivation – these concern the lack of ability to afford most basic requirements:

Keeping the home (household's principal accommodation) adequately warm.

Paying for a week's annual holiday away from home.

Replacing any worn-out furniture.

Buying new, rather than second hand clothes.

Eating meat chicken or fish every second day, if the household wanted to.

Having friends or family for a drink or meal at least once a month.

Inability to meet payment of scheduled mortgage payments, utility bills or hire purchase instalments.

2 Secondary non-monetary deprivation – these concern enforced lack of widely desired possessions ("enforced" means that the lack of possession is because of lack of resources):

A car or van.

A colour TV.

A video recorder.

A micro wave.

A dishwasher.

A telephone.

3 Lacking housing facilities – these concern the absence of basic housing facilities (so basic that one can presume all households would wish to have them):

A bath or shower.

An indoor flushing toilet.

Hot running water.

4 Housing deterioration – these concern serious problems with accommodation:

Leaky roof.

Damp walls, floors, foundation etc.

Rot in window frames or floors.

5 Environmental problems – these concern problems with the neighbourhood and the environment:

Shortage of space.

Noise from neighbours or outside.

Dwelling too dark/not enough light.

Pollution, grime or other environmental problems caused by traffic or industry.

Vandalism or crime in the area.

## Appendix 2: Estimation of the propensity score of marital disruption: matching variables

WAVE DUMMIES

Age

Number of Children

COUNTRY DUMMIES

Log HH income ( $t-1$ )

Log person income ( $t-1$ )

Deprivation index ( $t-1$ )

EMPLOYMENT STATUS (REFERENCE: EMPLOYED)

Student

Out of labour force

Unemployed

EDUCATION (REFERENCE: LESS THAN SECONDARY LEVEL)

Degree

Secondary

Shortage of space in HH