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Marital disruption and economic well-being: a comparative analysis

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Summary. 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. We implement several measures of well-being of monetary and multidimensional nature by using data from the European Community Household Panel. Another issue in this literature concerns selection bias of divorcing couples. We tackle this issue by using a propensity score matching technique combined with a difference-in-differences estimator. Results confirm the importance of the definition of well-being. We find a strong gender bias when using monetary measures but a considerably lower bias, and for some countries non-existent, when using non-monetary indices.

Keywords: Deprivation indices; European Community Household Panel; Marital disruption; Poverty; Propensity score matching; Relative income

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 of continuing concern. Most of the debate around the economic consequences of divorce is focused 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 USA and Europe showed that women experiencing a divorce tend to suffer a substantial loss of income, whereas men's economic circumstances seem quite unaffected or even improve slightly in some cases (Burkhauser *et al.*, 1991; Fritzell, 1990; Jarvis and Jenkins,

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1999; Manting and Bouman, 2004; Poortman, 2000; Smock, 1993, 1994). The reasons behind this pattern are many. One is that women tend to have a lower attachment to the labour market 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 why many divorced women suffer financially.

An equally consistent finding is strong differences between countries in terms of the economic penalty that is 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 which is characterized by poorer welfare provision. Andreß *et al.* (2004), comparing Belgium, Germany, Italy, UK and Sweden by analysing the three main providers of individual welfare, the family, the market and the state, showed that the configuration of these providers to a large extent determines the economic outcome of marital dissolution. It was shown that because of the 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 from that of 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 child care infrastructure, a system which enables Swedish mothers to work full time to a much greater extent than in 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 life style of divorced men. Thus it may seem difficult 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 criticized 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 a loss of income 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 recognized that well-being itself has many more dimensions, which are 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 composition of a household, 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 by using a range of 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 who are not. 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 that they would experience if they were to split from their partner (Becker, 1991). One way to tackle this issue is to implement a *propensity score matching* technique which nets out the effect of separation from the confounding effects that are 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 that is caused by unobserved variables, provided that these are time invariant. Heckman *et al.* (1998) showed that the assumption underlying the DD–PSM estimator is less restrictive than the simple conditional independence assumption (CIA) on which propensity score matching 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 by using data from the European Community Household Panel, which offers unique scope for comparability at the European level (Eurostat, 1999). Uunk (2004) showed 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 show that differences in terms of economic strains that are 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 that are associated with marital disruption. Finally, we recognize the importance of children of the couple, so we make separate estimates for couples with children only. These estimates are compared with the cases where we include couples with *and* without children.

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

2. Definition and measurement of well-being

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% of the median net household equivalized income. Individuals within the household are deemed poor if their 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 the median net equivalized household income is chosen, but many use alternative poverty thresholds of 50% and 70% (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 from those of children. Additionally, there are economies of scale, meaning (for example) that two adults can live together less expensively 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 the total household income by this equivalence scale. We apply the commonly used Organisation for Economic Co-operation and Development modified equivalence scale, which gives a weight of 1 for the first adult, 0.5 for adults other than the household head and 0.3 for children. The original Organisation for Economic Co-operation and Development scale was based on empirical studies during the 1980s 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 Organisation for Economic Co-operation and Development scale did not properly take into account economies of scale. For this reason the 'modified Organisation for Economic Co-operation and Development' scale was proposed and officially adopted by Eurostat as the common scale in the European Community Household Panel.

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 characterizes 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 that is 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). Many researchers 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 utilizes 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 people who are near to the chosen line tends to be overemphasized, whereas that of people who are far from that line is largely ignored (Verma and Betti, 2005).

The degree of income poverty that is associated with each individual as specified below was first proposed by Betti and Verma (1999), who termed it 'fuzzy monetary'. The approach was later officially adopted by Eurostat (2002). The approach can be explained as follows. Let y_i be the net equivalized household income that is associated with individual i. An income index is defined as the sum of the incomes of individuals who are richer than the individual concerned

(the numerator), divided by the sum of the incomes of all n individuals in the sample (the denominator):

$$V_i = \sum_j y_j | y_j > y_i \Big/ \sum_{j=1}^{n} y_j. \tag{1}$$

This index can be seen as the share of the total income that belongs to individuals who are 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 individual with the highest income. The degree of income poverty FM_i is defined as a monotonically increasing function of the income index V_i:

$$\mathrm{FM}_i = V_i^{\alpha / H_c} \tag{2}$$

where H_c is the proportion of individuals who fall below the country poverty threshold, defined as 60% of the median net equivalized household income. The parameter α 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. This is imposed to facilitate better a comparison with the conventional approach.

2.3. A multidimensional and comparative perspective: the deprivation index measures

The relative income measure FM_i overcomes the simplistic categorization of the population into poor and non-poor. 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 in income: some can experience a dramatic rise in monthly expenses (e.g. 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 who are involved.

Just as in the fuzzy monetary approach that was 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 (Table 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 index of deprivation but also the indices of deprivation for each dimension of well-being. Approaches of this kind applied to poverty analysis of European countries are becoming increasingly common (Eurostat, 2002; Aassve *et al.*, 2005). The five dimensions $\delta = 1, 2, \ldots, 5$ 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; 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 $m = 1, \ldots, M$ 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

Table 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 television
A video recorder
A microwave oven
A dish-washer
A telephone
3. Lacking housing facilities—these concern the absence of basic housing facilities (so basic that one can presume that 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 or not enough light
Pollution, grime or other environmental problems caused by traffic or industry
Vandalism or crime in the area

$$d_{ik} = \frac{M_k - m_{ik}}{M_k - 1}, \qquad 1 \leqslant m_{ik} \leqslant M_k. \tag{3}$$

The third step involves determining weights to be assigned to each item of the index of deprivation. The weighting procedure that we propose here is a variant of the procedure that was 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 between individuals in the population, i.e. 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 individuals affected (Filippone *et al.*, 2001)—are given a larger weight. Secondly, 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 δ 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 have a video recorder unless they have 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 that is plagued by rot in window-frames or floors is also more likely to report damp

walls, floors and foundations (see Table 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 $S_{\delta,\iota}$ only considering the items belonging to dimension δ:

$$S_{\delta,i} = \sum_k w_k(1 - d_{ik}) \Big/ \sum_k w_k \tag{4}$$

where w_k are the weights that were defined above in the third step. Note that equation (4) 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 that were defined above:

$$S_i = \frac{1}{5} \sum_{\delta=1}^{5} S_{\delta,i}. \tag{5}$$

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' that is assigned to all individuals who are less deprived than the person concerned. It varies from 1 for the most deprived to 0 for the least deprived individual:

$$\mathrm{FS}_i = \sum_j S_j | S_j > S_i \Big/ \sum_{j=1}^{n} S_j. \tag{6}$$

The same formulation is applied within each of the five dimensions 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 is a set of comparable large scale longitudinal studies implemented by the European Union. The first wave of the European Community Household Panel was collected in 1994 for the original countries in the survey: Germany, Denmark, the Netherlands, Belgium, Luxembourg, France, UK, Republic of 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 its small sample size, Sweden because the data do not form a panel and Germany is dropped because the information that is necessary to construct the indices of deprivation is not available. Eight waves of the European Community Household Panel were collected in total, the last being in 2001. We aggregate data according to the welfare regime clusters that were defined by Esping-Andersen (1990, 1999) and Trifiletti (1999). The clusters are as follows: *liberal countries* (the UK and the Republic of 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 which is defined by separation or a divorce. Normally a marital dissolution is initiated by a separation and followed by a formal divorce. The European Community Household Panel 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, we 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 self-reported martial status to define the time when a marital split occurred.

In this analysis a couple are 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 effect of marital split on poverty status we exclude those couples who were 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 that is used. By using propensity score matching (to be explained in the 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 before 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, i.e. couples who are experiencing a marital separation may be qualitatively different from couples who are not separating. For example, women who are strongly dependent on their partner's income are probably less likely to separate as they are aware of the strong economic distress that they would experience in the case that they split from their partner (Becker, 1991). Here we tackle this issue by implementing a propensity score matching technique. Applications of this kind are growing in the literature (see, among others, Blundell *et al.* (2005), Lechner (2002) and Dehejia and Wahba (1999)).

In our setting we assume that each individual i has two potential outcomes: Y_{1i} in the case that he or she experiences a marital split (the treatment) and Y_{0i} in the case that he or 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 effect 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 which is 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 the treated* ATET and is expressed as

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

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 the observed difference between the treated and control groups, which assumes that

$$\text{ATET} = E(Y_{1i} | D_i = 1) - E(Y_{0i} | D_i = 0). \tag{8}$$

But equation (8) assumes that there is *no* selection bias, which means that the group of treated individuals 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 the treated and control groups are systematically different, implying that an estimate which is based on equation (8) would be a biased for the true ATET. It is important to understand the nature of the bias that arises. Heckman *et al.* (1998) proposed to write the bias B as a function of a set of pretreatment observed covariates X:

$$B = \int_{S_{1X}} E(Y_0|X, D=1)\, \mathrm{d}F(X|D=1) - \int_{S_{0X}} E(Y_0|X, D=0)\, \mathrm{d}F(X|D=0) \tag{9}$$

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 that we observe for the treated group ($D=1$) and for the control group ($D=0$). On the basis of equation (9) Heckman *et al.* (1998) derived a decomposition of B into three terms that they referred 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 lie outside 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 misweighting 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 group (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 equation (9) are eliminated by our proposed matching procedure is explained below.

The matching method is based on the critical assumption termed the CIA, stating that treatment status is random conditionally on a given set of X. The CIA is formally expressed as

$$Y_0 \perp D|X \tag{10}$$

and means that conditionally on X the potential outcome in the case of non-treatment (i.e. Y_0) is independent of the treatment status. Whereas expression (10) imposes full independence, identification of ATET requires a less restrictive condition. As Smith and Todd (2005) pointed out it suffices that we impose *mean* independence, i.e. $E(Y_{0i}|X_i, D_i=0) = E(Y_{0i}|X_i, D_i=1)$. Thus ATET can be written as

$$\begin{aligned}
\text{ATET} &= E(Y_{1i} - Y_{0i}|D_i=1) = E_{X|D=1}\{E(Y_{1i} - Y_{0i}|X_i, D_i=1)\} \\
&= E_{X|D=1}\{E(Y_{1i}|X_i, D_i=1) - E_{X|D=1}\{E(Y_{0i}|X_i, D_i=1)\} \\
&= E_{X|D=1}\{E(Y_{1i}|X_i, D_i=1)\} - E_{X|D=1}\{E(Y_{0i}|X_i, D_i=0)\}.
\end{aligned} \tag{11}$$

The first two lines in equation (11) expand the ATET that was defined in equation (7), whereas in the third line we apply the CIA in its less restrictive form by substituting the second (unidentified) term $E_{X|D=1}\{E(Y_{0i}|X_i, D_i=1)\}$ by $E_{X|D=1}\{E(Y_{0i}|X_i, D_i=0)\}$. In equation (11) ATET is now fully identified and is a direct consequence of the 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 equation (11). Instead of matching on the basis of X we 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: $p(X_i) = \Pr(D_i = 1|X_i)$. 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

$$\text{ATET} = E_{p(X_i)}[E\{Y_{1i}|D_i = 1, p(X_i)\} - E\{Y_{0i}|D_i = 0, p(X_i)\}]. \tag{12}$$

Several matching procedures can be used to estimate equation (12) (see, for example, Becker and Ichino (2002) and Smith and Todd (2005)), but all can be seen as generated by the formula

$$\widehat{\text{ATET}} = \frac{1}{n_1} \sum_{i \in \{D_i = 1\}} \left(Y_{1i} - \sum_{j \in \{D_j = 0\}} W_{ij} Y_{0j} \right) \tag{13}$$

where the weight w_{ij} is defined according to the matching method that is 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 1 for unit j that is closest; otherwise the weight is 0 (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 unit). If so we use all these k controls with weight $1/k$.

All three sources of bias in equation (9) 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 reweighted according to the value of $p(X)$, leading to balance of X between treated and control units. B_3 is the only component of equation (9) that is not eliminated by matching but is assumed to be 0 by the CIA.

The matching procedure that was described above can be implemented on cross-sectional observations by recording both those treated (i.e. those who experienced a marital split) and the controls (those who did not split). However, our source of data 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 can compare the mean change of well-being from one time period t to another, $t + 1$, of treated units, with the mean change of well-being for the same time period for controls. It means that we can define a difference-in-difference estimator as follows:

$$\begin{aligned} \text{DD} &= E(Y_{1i}^{t+1} - Y_{1i}^{t}) - E(Y_{0i}^{t+1} - Y_{0i}^{t}) \\ &= E(\Delta_{1i}) - E(\Delta_{0i}). \end{aligned} \tag{14}$$

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 more clearly we can define the pointwise 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 0, the critical identifying assumption for the DD-estimator is that

$$B^{t+1}(X) - B^t(X) = 0. \tag{15}$$

In this way the CIA is relaxed since we no longer assume that $B(X)$ is 0, rather we only assume that its value does not change from wave t to wave $t + 1$. In other words, provided that unobserved heterogeneity is time fixed, its effect will be netted out by taking the 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 that are correlated with both the outcome and the treatment in the model estimating the propensity score makes equation (15) more likely to hold. The final estimator of the effect of marital split on well-being is given by

$$\text{DD}-\text{PSM} = E_{p(X_i)}[E\{\Delta_{1i}|D_i = 1, p(X_i)\} - E\{\Delta_{0i}|D_i = 0, p(X_i)\}]. \tag{16}$$

Estimation of equation (16) is done by using the estimator that is defined in equation (13), where the values of Δ_{1i} and Δ_{0i} replace 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 before the event (measured both in terms of income and in terms of deprivation), education and employment status (Table 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 those treated and the matched controls. If this is so then all covariates in X are balanced and satisfy what is termed the balancing property (Augurzky and Schmidt, 2001). 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 although their propensity scores are 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 a t-test for equality of means for each covariate X, before and after matching. This is likely to suffice, considering

Table 2. Estimation of the propensity score of marital disruption: matching variables

Wave dummy variables
Age
Number of children
Country dummy variables
Log(household 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 household

that almost all the variables in our application are binary. Accepting the null hypothesis means that control units are not different from the treated units except for the treatment status. Other tests have been suggested (see, for instance, Sianesi (2004) and Dehejia and Wahba (2002)) but there is no clear consensus about which is the most powerful. Becker and Ichino (2002) argued 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 in the propensity score model. The balancing property is satisfied in all 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 that we adopt, using the module that was 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 by using the usual standard normal distribution approximation.

4. Results

4.1. Entering poverty

Taking the sample of individuals who were defined as non-poor in the previous period, Table 3 presents the effects of experiencing a divorce or separation event on entering poverty by using different poverty thresholds. Note that the estimate refers to what is called

Table 3. Average treatment effect on the poverty entry rate at different poverty thresholds by gender, presence of children and welfare regime

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

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 with men. This is so independent of countries and the poverty threshold that is used. Moreover, the effects are largely consistent with welfare regime theory. Especially with the 50% threshold, the ranking of country groups is perfectly in line with 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% of median income definition, 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 the poverty line that is used. Note that the sample mainly consists of individuals from the UK, as the number of separations and divorce is quite low in the Republic of Ireland. As expected the effect for men is far lower and significant only in the conservative group (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 in the Mediterranean countries. When we consider only couples with children the effect of marital disruption is even stronger: for women in liberal countries the rise of the 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.

4.2. Fuzzy monetary indicator

The results that are reported in Table 4 are the estimates of the 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 dissolution of marriage, whereas a negative effect means a rise. Therefore in liberal countries, for instance, women tend to experience a strong decline whereas the men's ranking remains approximately the same after the separation or divorce. Whereas the results for liberal countries are consistent with the estimated poverty entry

Table 4. Average treatment effect of marital dissolution on relative income

Countries	*Results for males*				*Results for 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

rates that are presented in Table 3, the situation is less straightforward for the other countries. In contrast with the poverty entry rate for men, where some significant effects are noted, declines for men are generally small and statistically insignificant in all groups. 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 countries 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 with all women (with or without children). This is largely in line with results of divorce on the poverty entry rate that are reported in Table 3. Mediterranean and conservative countries show no relevant difference between the whole sample and women with children. This is again consistent with results in Table 3.

4.3. Indices of deprivation

We now move to the effect of marital dissolution on total household deprivation. We consider first the change in total deprivation index due to separation from the 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 the treated for the overall index of deprivation, the basic life style deprivation index and the secondary life style 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.

4.3.1. Total deprivation

The results that are reported in Table 5 show a somewhat different picture from 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 effect among the Conservative countries. Importantly, the effects are now significant also for men and, though the magnitude of the effects is always lower than for women, there is less of a gender gap.

The effect for men in the liberal group is particularly striking. Compared with poverty entry and change in relative income (Tables 3 and 4), the effect on deprivation for men is considerably

Table 5. Average treatment effect of marital dissolution on the index of deprivation

Countries	*Results for males*				*Results for 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

stronger, and not too different from that for women. The gender bias in liberal countries is considerably smaller in so far as well-being 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 that we reported for poverty entry. As with the liberal countries, conservative countries now show quite a 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 conutries, 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 effect in the conservative countries.

4.3.2. Basic life style deprivation

If we focus on the first dimension of deprivation, i.e. deprivation on basic life style, we find results shown in Table 6 that are relatively consistent with results for the total deprivation index. Again the liberal group shows the strongest effect both for men and for women, but this time the effect for women is about twice as high compared with that for 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 badly off compared with 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 influence the effect for men negatively: 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 in conservative countries.

4.3.3. Secondary life style deprivation

Finally, in Table 7 we look at the effects of marital disruption on the level of deprivation concerning secondary life style 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

Table 6. Average treatment effect of marital dissolution on the basic life style deprivation index

Countries	*Results for males*				*Results for 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

Table 7. Average treatment effect of marital dissolution on the secondary life style deprivation index

Countries	*Results for males*				*Results for 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

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 the 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 life style deprivation. The role of children is particularly interesting for liberal countries. The estimates for couples with children reveal that for men the effect remains the same, whereas it is lower for women, though the estimate is not significant at the 5% level. A substantial drop is registered also for women in social democratic countries combined with an increase for men in social democratic countries. Conversely we observe a small increase in the effect for women in the other two country groups, whereas there is no relevant change for men in these groups.

5. Concluding remarks

The present work is concerned with the economic consequences of marital disruption for both 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 or separation. Some researchers (McManus and DiPrete, 2001) have challenged this evidence, suggesting that the gender bias is less strong than is 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 a 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 life style 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 can 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 the poverty threshold is an important issue. Results differ considerably depending on whether we use a 50%, 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. indices of deprivation) 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 that is estimated for men and women (this is so in liberal countries when using the overall index of deprivation and the secondary life style 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, in contrast, 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 virtually unaltered.

However, in terms of deprivation, men do suffer significantly. Many of the items that are used to compute the index of deprivation refer to characteristics of the dwelling. If it is the case that men normally must leave the dwelling following a divorce, they will, in the short run at least, lose 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 life style 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 have no children.

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