The impact of cohabitation and divorce on partners' labour force participation: comparing Britain to Flanders

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Abstract

In this paper we look into the possible impact on labour force participation of two demographic variables that have undergone considerable changes in the past few decades: divorce and cohabitation. More specifically we analyse the labour force participation probabilities of men and women currently living with a partner and study the impact of a previous divorce or separation and current non-marital cohabitation. We use seven waves of data of the BHPS and the PSBH to compare British results with results on Flemish individuals. Estimates suggest that cohabitation implies significantly higher labour force participation for women, especially for the older cohorts. A divorce experience is generally found to be insignificant, except for British men who are less likely to be in the labour force after experiencing a divorce or separation than without this experience.

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Non-technical summary

In this paper we investigate the labour supply of British and Flemish households, looking at it from the particular angle of recent changes in the frequency of divorce and cohabitation. We focus on the medium term effect of divorce and study the effect of a divorce experience on 're-partnered' individuals and compare their labour choices to partners who have not lived through such an experience. Analogously, we study people who are currently cohabiting and compare them to legally married individuals.

We hypothesise that individuals will take their separation risk into account when deciding on their labour force participation. Hence, we conjecture that both divorced persons and cohabiting individuals will show higher than average labour market participation because they know that they run higher risks of experiencing a separation in the future, which will make them the sole income provider of their household. However, we also assume that the risk discounting effect will range from minor to non-existent for men and younger women, because they can hardly increase their already high degree of labour force participation.

In the empirical part of this paper, we use seven early waves of the BHPS and PSBH (British Household Panel Study and Panel Study of Belgian Households) and observe, in the first place, a quite pronounced increase of labour force participation among cohabiting women in both Flanders and Britain, compared to married women. Additionally, the effect is diminishing in age, i.e. in the youngest cohort hardly any effect is noted. Equally in line with our hypothesis, men's behaviour is not affected by the legal status of their union. Concerning divorce, our hypotheses are not corroborated for women. For men the results are more in line with the hypotheses, but British men somewhat stretch the hypothesis, showing lower degrees of labour force attachment with divorce experience than without it.

At present, we can conclude that cohabitation tends to increase female labour force participation, but that this effect is diminishing over time. Regarding divorce, our results suggests that the widely observed effect on female labour force participation tends to disappear once divorced women start living with a new partner.

Table of contents

1 Introduction 1

2 The impact of marital status and history on the spouses' paid labour propensity: previous empirical results and hypotheses. 2

2.1 The impact of past divorce on present labour decisions 3

2.2 Cohabitation and labour choices 7

2.3 Cohabitation and divorce 9

3 An empirical enquiry among British and Flemish couples 9

3.1 The sources and selection of data 10

3.2 The empirical specification of household labour decisions 12

3.3 Cohabitation and divorce in the BHPS and PSBH 14

3.4 The dependent variable: labour force participation 16

3.5 The empirical content of the independent variables 19

3.6 Estimation procedure 22

4 Estimates, simulations and discussion 24

4.1 The basic specification 24

4.2 Marital status and history: basic versus interacted specification 28

4.3 A decomposition exercise 32

5 Conclusions 35

6 References 37

7 Annexes 41

1 Introduction

In this paper we investigate the labour supply of British and Flemish households, looking at it from the particular angle of recent changes in the frequency of divorce and cohabitation. In the past, divorce and labour supply have been frequently linked in studies on the immediate consequences of divorce, especially for lone parents. Moreover, the causal direction in the relation between divorce and labour supply has attracted many scholars' attention, leading to empirical findings that alternatively support unidirectional, bi-directional and latent variable explanations.

This paper differs from the previous literature in both respects. First, we focus on a medium term effect of divorce and, secondly, we have no causal aspirations. Rather than looking into the increased probabilities of labour force entry caused by divorce (anticipation) or comparing the divorce probabilities of working and non-working individuals, we will study the effect of a divorce experience on 're-partnered' individuals and compare their labour choices with those of partners who have not lived through such an experience. Analogously, we will study people who are currently cohabiting and compare them to legally married individuals.

The underlying hypothesis is the following: if the rising degree of cohabitation and divorce alters the relations in existing couples and changes the bargaining position of wives and husbands, it may well affect the outcome of marital bargaining relating to the partners' involvement in the labour market. This, in turn, would influence the total labour supply and hence is of interest for labour policy makers.

We develop our analytical story along three stages. First, we will discuss previous empirical findings and derive some working hypotheses about the relation between current cohabitation, a past experience of divorce and current labour choices. We will argue that both cohabitation and divorce can be expected to lead to a higher than average female labour force participation, but do not have an effect on men. Inevitably, we will touch upon the ongoing discussion on the causal directions in the decision triad marriage/divorce, labour and children. However, we will not elaborate a real causal analysis, because the purpose of this paper is to look into the labour supply consequences of the current and predicted rise in divorce and cohabitation.

In the second chapter, we will develop an estimation strategy to get an idea of the empirical value of our hypotheses. Therefore, we analyse the data of the Belgian and British household panel studies, using seven early waves of both of these comprehensive data collections. We use the rich dataset to construct fine indicators of most of the variables mentioned in our hypotheses. Furthermore, we estimate reduced form specifications that are consistent with the basic ideas of collective household models.

Finally, we will devote a chapter to the discussion of the estimates and some further simulation and decomposition exercises. The latter highlight the most relevant results and focus on comparisons between both male and female respondents and Flemish and British respondents. In both these comparisons, an intent is made to differentiate between the impact of behavioural differences and differences in population characteristics on the household's decision on labour force participation.

2 The impact of marital status and history on the spouses' paid labour propensity: previous empirical results and hypotheses.

Though there are obviously a wide array of reasons to look for market work and postmodernists have argued that non-monetary motivations are becoming increasingly important in Western societies where basic needs are almost universally covered, the provision of income has remained a very important function of market labour. This observation is routinely made when analysing survey questions on labour motivation, but can also be inferred from the almost universally significant coefficient of income variables in estimated labour supply functions.

With the provision of income as a cornerstone of labour motivation, a household dimension is added for persons living with a partner. In a society without marital break-up and with a high degree of job stability, spouses could afford to rely on the income of their partner and devote their efforts to non-market labour. However, the last three decades have been characterised by both a higher (and growing) number of marital separations and a lesser degree of labour stability than what had become usual during the preceding decades. Therefore, we will argue that people's lifetime perspective has dramatically changed and people have reacted to the altered risks by changing their labour market behaviour. With a higher probability of separation and spouse's unemployment, the risk of becoming the sole breadwinner of the household has also risen. For individuals who were previously economically inactive (mainly women), it becomes increasingly rational to look for risk insurance through paid work.

Of course, many other valuable explanations for the increasing labour market participation of women have been put forward. Among others, growing individualism, changing attitudes towards the household labour distribution, new ideas about fatherhood, the diminishing educational gap and public policies towards equal opportunities and non-discrimination are commonly cited. We will come back to some of these at a later stage, but for now we will focus on the risk insurance motive. More specifically, we will discuss the effect of a past divorce or separation experience on current labour choices and look into the similar relation for people who are currently cohabiting.

2.1 The impact of past divorce on present labour decisions

Beck and Hartmann (1999) evoke Andreas Diekmann's concept of a 'divorce spiral'. Largerly within the line of reasoning of Becker (1991), Diekmann argues that the growing number of divorces has become a self-enforcing process. "The observation of growing marital instability motivates people to invest less in marriage specific capital –especially children-, which would in turn have diminished the risk of separation. Moreover, the growing number of divorced people improves the prospects of finding a similarly aged, new partner after a separation. Furthermore, the growing proportion of marital break-ups has enhanced the social acceptability of a divorce, thereby lowering the normative pressure to carry on with a marriage. These three developments all ease the decision to end a marriage. … Finally, a last aspect of the 'divorce spiral' is the mutual influence of female labour force participation and divorce risk. On the one hand, female labour force participation increases the risk of marital dissolution and the feeling of marital instability. On the other hand, the observation of a higher divorce risk increases the inclination to be economically active." (Beck and Hartmann,1999:655-656, author's translation)

Most empirical work on the relation between divorces and labour force participation starts with analogous statements on mutual influencing. As an example from the economic literature, Lundberg and Rose state: "determining the direction of the causality between divorce and female labour force participation" (…) is "one of the classic 'chicken and egg' problems in labour economics" (1999:21). On the one hand, working women are frequently observed to be more likely to divorce. This may be explained by purely monetary considerations (she can afford it) and/or by commonly unobserved characteristics like individualism that may account for both a higher inclination towards paid work and a higher risk of divorce. On the other hand, it is empirically undisputed that divorce goes together with growing labour force participation among women. This is not so surprising since most women go through a certain period of economic hardship after divorce. An example of the latter is provided by John Ermisch' analysis of lone parenthood in the UK. The author observes "a rise in the overall labour force participation rate of women who become lone mothers, and (…) this occurs from the time of the end of the first marriage, with a particular sharp rise during the first year after becoming a lone mother." (Ermisch,1991:116).

Other studies have added more details to this picture. Nakamura and Nakamura (1996) confirm the general results with US data (divorce increases labour), but note that a considerable part of the divorced women anticipate their divorce and start working some time before the divorce. Johnson and Skinner (1986) analyse labour supply and divorce in a simultaneous framework and find little evidence of an effect of labour supply on divorce probability[1]. Van der Klaauw (1996) develops a structural model of female labour supply and marital status decisions and finds that the marital status decision is actually endogenous in the labour supply decision of US women. However, he also observes that endogeneity bias is generally small, except for certain groups like young unmarried women (Van der Klaauw,1996:225).

Beck and Hartmann (1999) analyse the German situation[2] and observe, among older cohorts, a clear influence of female full-time working on the probability of divorce. Interestingly, part-time working does not seem to have any influence. Additionally, the effect disappears in younger cohorts and is generally weakened by divorce anticipation. The authors' conjecture is that the effect will disappear in time with the movement of cohorts, given that female full-time employment is no longer a distinctive characteristic. On the other hand, both actual divorce

[1] As in Nakamura and Nakamura (1996) some evidence of anticipatory behaviour of women is found. This is not surprising, since both papers use PSID data. Additionally, both articles indicate that most of the increase in labour supply is due to women with little work experience prior to the divorce.

[2] We only report on their results for the former German Federal Republic. Results for the former German Democratic Republic are highly divergent for largely historic reasons. See Beck and Hartmann (1999) for more details.

and the feeling of marital instability increase the likelihood of labour force participation. Again, the effect is weakened, but not entirely eliminated, by anticipatory behaviour.

Unfortunately, none of these studies has specifically focused on the group of reconstituted couples and most do not even regard couples in general. As stated before, an increase in labour force participation after a divorce is completely logical because of the need for income. This is especially true when observing the relatively frequent problems with maintenance payments, that make reliance on the former partner unfeasible, and the reinforced labour link of welfare state provisions, through which governments stress the pivotal role of market labour in every individual's income trajectory and limit prospects of long-term reliance on benefits.

However little relevant to the labour decisions after the start of a new partnership, these observations definitely do not contradict a continued higher degree of labour force participation after the start of a new relationship, which can be expected for several reasons. In the first place, the actual experience of divorce will, even more so than the perception of marital instability, increase the person's (perception of) risk of future separation or divorce. In fact, apart from the subjective risk perception, research repeatedly linked previous divorce to a higher current divorce probability. Thus, seen from a risk insurance perspective, it is completely rational for individuals to continue in paid work even after entering a new relationship. Secondly, there may be some kind of "habit formation". Women, working at home before the divorce, may have grown used or even attached to their paid job if they started working because of the divorce[3].

From these observations, we can derive a first hypothesis for empirical testing:

1. *Partners who have experienced a separation or divorce before entering their current relationship show higher degrees of labour market participation than partners without this experience.*

Moreover, risk insurance motives may also be at work regarding the partner's past. Logically, risk calculus does not only incorporate the personal experience, but also adds the partner's history. Therefore, we can formulate as a second hypothesis:

[3] Earlier studies have indicated that childbirth is one of the major causes of labour market transition. As a simplifying hypothesis, we suppose that the occurrence of this event is independent of the transition to a reconstituted couple.

2. *Persons who live with partners who have experienced divorce or separation in the past, show higher degrees of labour market participation than persons living in a divorce-free household.*

However, some caution applies when interpreting the former hypotheses. As stated at the beginning of this chapter, many reasons explain (part of) the rise in (female) labour force participation during the past decades. Some groups have shown extraordinary rises in labour force participation, irrespective of their personal divorce experience[4]. Examples of these groups are the cohorts of younger women in general and young women with a high level of education more specifically.

Some caution applies also when analysing the behaviour of men. Indeed, the already existing high participation rate of men leaves little room for further increase. Conversely, new patterns of fatherhood may become particularly relevant to fathers who have divorced the mother of their children. For fathers, the recent tendency to grant both parents partial custody may confront them with a balancing act between their traditional role as main income-provider and the additional care-time demanded by the division of parental responsibilities[5]. Divorced fathers may well seize the opportunity of enhanced labour force participation of their current partner (for all the reasons mentioned) to reduce their own labour time.

As further hypotheses, we would therefore state:

3. *The effect of a divorce experience in the household is, ceteris paribus, smaller in younger cohorts than in older cohorts.*
4. *For men, the effect of a divorce experience on their labour force participation is non-existent.*
5. *Fathers with a personal divorce experience are more responsive to indicators of the need for care-time than non-divorced fathers.*

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[4] Though Diekmann (cited in Beck and Hartmann,1999:657) attributes at least part of this rise to rational discounting of the increased risk of separation and divorce.

[5] In their analysis of repartnering in Great Britain, Lampard and Peggs cite a divorced father's opinion: "I don't see them every day of the week any more, so when I do see them, they're all mine, I want to have all that time." (1999:460)

2.2 Cohabitation and labour choices

While less perceived as a social problem, cohabitation has, just like divorce, become much more widespread in recent decades. This enormous increase in Western Europe and North America has taken many different forms, however. In Scandinavia, cohabitation is becoming a full-fledged alternative to marriage. In other parts of Europe and in North America, it has to a great extent remained a sort of trial marriage[6]. Though many cohabitation spells end in separation, most marriages are now preceded by a period of cohabitation with the current spouse.

At first sight, cohabiting partners can be expected to be more self-oriented, not least because the threat of becoming the sole source of income in a one-person household is much more palpable than for married partners[7]. Additionally, it should be stressed that a break-up of cohabitation does not entitle the economically weaker partner to maintenance payments. Thus, the economic risk of specialising in home work is greater in a cohabiting relationship than in marriage (Bernasco and Giesen,1997) and forward-looking individuals may be expected to look for a paid job.

As with divorce, the gender balance is not immediately clear. Since both partners face a higher risk of separation than when married, their rational response might be to increase their labour force participation. However, men again face the limits of their already high degree of labour force participation. Therefore male partners may prefer to seize the opportunity given by higher labour force participation of their partners (and the consequently higher level of household income) to reduce their labour time, gaining leisure time or time to spend with their children. This does not imply, however, that men would retreat from the labour market, as this would engender too great a risk. As the economically weaker partner[8], the female spouse may be perfectly happy with this arrangement, because it allows her to reduce her dependence on the income provided by her male partner.

[6] See Ermisch and Francesconi (1999) for a detailed analysis of cohabiting in Great Britain.
[7] Separation is consistently found to be more likely for cohabitation than for formal marriage (e.g. Kravdal,1999)
[8] In most cases, women are responsible for less than half of the household's earnings and, given educational homogamy and the gender pay gap, will not earn more than their partner even when both are working full time.

It remains to be seen, however, whether these elements of difference really have some impact on actual behaviour. With cohabiting as a phase before marriage, cohabitation is much more prevalent among the younger cohorts, where the pattern of dual careers has become quite dominant, irrespective of the marital status. The balance of power may differ between young cohabiting partners and their married counterparts, but in practice the impact of the inclination towards paid work of young women and men dominates the decision process in both cases.

As a second mitigating argument, we could presume that, when cohabitation lasts a long time, partners will develop a mutual trust that is fairly comparable to the feelings many married partners have towards each other and the initial risk differential may disappear. Over time, only ‘habit formation’ would then explain a higher than average labour force participation.

As with divorce, some authors reverse the direction of causality. According to Schmidt (1992) partners in a dual-career household prefer cohabitation to marriage because they want to safeguard their bargaining position and marriage lowers the level of their ‘threat point’. Ressler and Waters (1995) see the growing female labour force participation as the single most important explanation for the growing cohabitation rate in the USA. According to their view, the demand for marital flexibility, that underlies the choice for cohabitation, is to be related to the growing career drive of women who want to be able to leave their partner if one of the partners needs to relocate for his or her job. Empirical support for their hypothesis is weak, however.

Bernasco and Giesen (1997) obtain the opposite results for the Netherlands. The latter authors explicitly model the ‘chicken and egg’ problem and simultaneously estimate the choice for a type of relationship and labour force participation. Results indicate that labour choices do not determine the family type, but conversely marriage effectively depresses the labour force participation of women, or, in other words, cohabiting women tend to participate more actively in the labour force.

Overall, we would formulate the following hypotheses:

6. *Women who are cohabiting with their partners are more likely to be economically active than their married counterparts.*

7. *Men who are cohabiting with their partners tend to reduce their labour time, compared to their married counterparts, though no effect is expected on their degree of labour force participation.*
8. *Both effects are smaller to non-existent in younger cohorts.*
9. *The duration of the cohabiting union tends to reduce the effect of cohabitation on labour force participation.*

2.3 Cohabitation and divorce

While cohabitation has replaced marriage as the dominant mode of first partnership, Ermisch and Francesconi also note a strong increase in the proportion of cohabiting unions within the group of first partnerships after the dissolution of marriage. Comparing different sources, the authors conclude that in Great Britain currently more than 75% of women repartnering after marital dissolution enter cohabitation (1999:7).

Unfortunately, to date no information is available on dissolution risk differentials of cohabiting unions whether they were preceded by marriage or not. Therefore, we will not formulate any additional hypotheses on the labour choices of currently cohabiting respondents depending on their marital history.

3 An empirical enquiry among British and Flemish couples

In this chapter we will test the hypotheses of the former chapter on data from two large sets of panel data. A first section will describe the data sources: the British Household Panel Survey (BHPS) and the Panel Study on Belgian Households (PSBH). Secondly, we will discuss our estimation method and provide descriptive data on the variables included. The third section will present estimates and in the final fourth section we will discuss the results elaborating some simulations.

3.1 The sources and selection of data

For the Belgian region of Flanders[9], we use data from the first seven waves of the Panel Study on Belgian Households (PSBH). The first wave of this ongoing study was designed as a nationally representative sample of the population of Belgium living in private households in the spring of 1992. For Britain we rely on the data of the second to eighth wave of the British Household Panel Survey (BHPS). As in the Flemish case, the observation period thus spans 1992 to 1998. Completely analogous to the Belgian panel, the first wave of the BHPS was designed as a nationally representative sample of the population of Great Britain living in private households in 1991. In both cases, the original sample respondents have been followed (even if they split off from their original households) and they, and their adult co-residents, interviewed at approximately one year intervals subsequently[10]. From wave three onwards the PSBH is integrated in the European Community Household Panel (ECHP). BHPS is providing the British data for the ECHP from its seventh wave[11] on.

For the purpose of this paper, we selected a sub-sample using several selection criteria. In the first place the sample was limited to people of working age. Hence, individuals had to have reached the age of 19 (Flanders) or 16 (Britain) and not passed the age of 60 (Flanders) or 64 (Britain) at the observation moment to be included in the sub-sample. For Flanders, we used 60 as the upper age limit though the legal retirement age for men is 65, because women's retirement age is 60 and especially because the actual age at retirement is 60 or lower for almost all employees. In Britain, employment seems to go on until the legal retirement age to a much larger extent. Therefore we maintained 64 as the upper limit.

Most importantly, the sample was restricted to observations of individuals who were living with a partner at the moment of the interview. Our theoretical model relies on information on both the respondent and his or her partner to predict the individual's labour force participation. Hence, a selection of "partnered" individuals is needed. However, a selection of individuals who lived with a partner for the full period of seven consecutive years would introduce a serious selection bias. This type of restriction would mean that persons who lived without a

[9] The sub-sample was restricted to Flemish observations, because crucial questions on marital history are only available for this region of Belgium.
[10] See Taylor et al. (1998) for more details on the following rules.
[11] Currently, the integration of earlier waves of BHPS in ECHP is being realised.

partner for some years during the observation period would be excluded and therefore the variable on divorce experience would reduce to a divorce experience followed by at least seven years of partnership. Therefore, we included all observations of "partnered" individuals, irrespective of their situation at the earlier or later interview. This way, we constructed a so-called "unbalanced panel", in which most respondents do not report seven times.

In fact, uneven selection may be attributed to various causes. First, persons may not have reached the lower selection age (19/16) or may have passed the upper selection age (60/64) in some, but not all waves. Secondly, information on one of the variables of the model may be missing for the respondent or his/her partner. In panel studies the 'normal' non-response to questions related to income is aggravated by panel attrition. Unfortunately, our model causes missing data to have a double effect. The observation of the respondent is lost, but also the observation of the respondent's partner, since no estimation is possible on cases as soon as one variable is coded missing. In our databases, there are quite a number of cases with full information on the respondent, but without information on the partner, resulting in the loss of the complete observation.

Furthermore, we limited our sample to respondents living in nuclear family type households. This way, we attempted to minimise potential bias of our estimates through the presence of other household members at working age, apart from the partners in the couple[12]. We conjectured that household labour decisions in extended family type households or households with non-family members may be structurally different, but given the predominant position of nuclear family households, we chose not to develop specific analyses for the former type of households and confine our analysis to the latter.

The final panel database contains 9881 Flemish observations of 2969 individuals, averaging 3.34 interviews per individual. The British database contains 27847 observation on 6060 individuals, resulting in an average of 4.60 interviews per individual. This clearly more favourable result for Britain can also be deduced from the frequency of the number of full interviews per respondent, shown in Table 1. Response is equally distributed among the sexes

[12] This selection does not exclude all potential bias, however. Our sample will definitely contain some households with adult children. We consider them as 'marginal' to their parents' labour decision, because they will be, at most, considered temporary sources of income and, in most cases, parents tend to allow their children to save most of their income.

with respectively 49.8% and 49.4% observations of women (Flanders/Britain) and 50.2% and 50.6% of men. Additionally, the response pattern does not seem to differ according to the year of the interview.

Table 1 Distribution of Respondents according to the Number of Observations Included (waves)

	Flemish respondents		British respondents	
Number of Waves	Number	Percentage	Number	Percentage
1	1140	38.4	746	12.3
2	309	10.4	1033	17.0
3	243	8.2	467	7.7
4	212	7.1	461	7.6
5	299	10.1	498	8.2
6	311	10.5	685	11.3
7	455	15.3	2170	35.8
Total	2969	100.0	6060	100.0

3.2 *The empirical specification of household labour decisions*

We will estimate the labour force participation of man and women separately using the following reduced form equation:

$$lfp_i = \alpha_{i1} w_f + \alpha_{i2} w_m + \alpha_{i3} y_f + \alpha_{i4} y_m + \alpha_{i5} y_h + \alpha_{ix} X + \varepsilon_i \qquad (i = m,f)$$

The symbols in specification (5) read as follows: lfp_i labour force participation of person *i*, w_f wage of the female partner, w_m wage of the male partner, y_m non-labour income of the male partner, y_f non-labour income of the female partner, y_h non-labour income of the household, X a vector of covariates and ε_i the usual error term.

In the specification male and female behaviour are modelled symmetrically. Both men and women are supposed to account for both their own wage and their partner's wage (w_f and w_m) when deciding about labour force participation lfp_i.

It is readily understood that the empirical specification (5) does not impose restrictions of the "male chauvinist" type, where women are treated as "followers" taking into account the labour decisions of their partner, while the men are supposed not to take the women's decision into account. Restrictions of the type $\alpha_1=0$ for husbands (i=m), have been decisively rejected in all recent studies of household labour supply.

Non-labour income is introduced in three variables: male non-labour income (y_m), female non-labour income (y_f) and general household non-labour income (y_h). This distinction is made because the level of the threat points is crucially dependent on the income that the partners control (i.e. even when choosing the outside option). A typical example of female non-labour income in Flanders is the child allowance, because it is always given to the mother[13]. Clearly, the separation of non-labour income into three categories adds a 'collective household model'-flavour to the specification. Earlier tests of collective versus unitary household models have frequently focused on (and rejected) the 'pooling' restriction of the unitary model which supposes $\alpha_3 = \alpha_4 = \alpha_5$, since non-labour income is pooled in the unified decision process of the unitary model.

Of course, the use of a symmetrical structural description of the decision about labour force participation does by no means imply that the actual behaviour of men and women is comparable. First, coefficients on the earlier mentioned variables can differ. Secondly, men and women are known to react very differently to several covariates that are summarised in the vector X in equation (5).

To start with, the usual socio-demographic control variables are added to X: age, age squared and education dummies.

Furthermore, the number and age of children is well-known to be of crucial importance to the labour decisions of the household, especially of the mother. Lundberg (1988), for example, observed a very different pattern among couples with and without children and with young children. Applying the specification of Booth, Jenkins and García-Serrano (1999:180) to Flanders, we represent the children's influence by four dummies: "*one child*", "*two children*", "*three or more children*" and finally "*youngest child below three years old*". The non-linear specification of the impact of the number of children relies on earlier evidence found by Duysters and Vanherck (1997) in their study of the labour division in Flemish couples. More specifically, the authors observed a sharp fall in the labour force participation of women with the arrival of the third child. Additionally the age of the youngest child is introduced for its crucial influence on the demand for care and/or day-care (Browning,1992). In Belgium, a very high proportion of children enters nursery school at age two and a half. By age three, 98% of the children has entered nursery school and day-care is less of a problem (European Commission,1996). Following the same reasoning for Britain, we define young children as children below four years old, because nursery school attendance starts somewhat later in Britain, reaching 94% of four year olds (European Commission,1997a).

3.3 Cohabitation and divorce in the BHPS and PSBH

In both studies cohabitation is clearly distinguished from marriage for all persons who live with a partner at the moment of the interview. The divorce variable is deduced from the marital history questions in waves 2 and 8 and wave 7 for Britain and Flanders, respectively. Additionally, this information is matched with changes in marital status observed across the panel interviews.

Passing on to the actual distribution of our demographic categories of interest, it should be noted first that both cohabitation and personal divorce experience are well represented in the samples. In Flanders they represent respectively 8.2 and 7.0 % of the observations (810 and 690 cases). In Britain cohabiting people account for 10.0 % of the observations and respondents with a personal divorce experience for 18.8 % (2721 and 5113 cases).

Additionally, table 2 shows that cases are equally spread among men and women. Finally, it can be noted that a divorce experience is more common among cohabiting man and women than among married persons. This result is consistent with the earlier observation of Ermisch and Francesconi (1999) that the majority of partnerships after divorce are cohabiting unions and extends it to Flanders.

[13] Except when there is no mother (of course) or in the rare cases mothers have been deprived of their custodial rights.

Table 2 Distribution of Divorce Experience and Cohabitation

	Women Number	%		Men Number	%	
Flanders						
Married						
Neither partner previously divorced	4219	93.0		4225	93.1	
Spouse with divorce experience, respondent not	122	2.7		82	1.8	
Respondent with divorce experience, spouse not	85	1.9		121	2.7	
Both partners with divorce experience	109	2.4		108	2.4	
Total	4535	100.0	92.1	4536	100.0	91.5
Cohabiting						
Neither partner previously divorced	197	50.5		231	55.0	
Spouse with divorce experience, respondent not	79	20.3		36	8.6	
Respondent with divorce experience, spouse not	37	9.5		75	17.9	
Both partners with divorce experience	77	19.7		78	18.6	
Total	390	100.0	7.9	420	100.0	8.5
With personal divorce experience	308	6.3		382	7.7	
With any kind of divorce experience	509	10.3		500	10.1	
Total	4925	100.0	100.0	4956	100.0	100.0
Britain						
Married						
Neither partner previously divorced	9805	79.9		9975	79.7	
Spouse with divorce experience, respondent not	792	6.5		827	6.6	
Respondent with divorce experience, spouse not	804	6.6		809	6.5	
Both partners with divorce experience	866	7.1		897	7.2	
Total	12267	100.0	89.2	12508	100.0	88.7
Cohabiting						
Neither partner previously divorced	684	46.2		760	47.7	
Spouse with divorce experience, respondent not	197	13.3		232	14.6	
Respondent with divorce experience, spouse not	215	14.5		214	13.4	
Both partners with divorce experience	384	25.9		386	24.2	
Total	1480	100.0	10.8	1592	100.0	11.3
With personal divorce experience	2269	16.5		2306	16.4	
With any kind of divorce experience	3258	23.7		3365	23.9	
Total	13747	100.0	100.0	14100	100.0	100.0

Sample: All observations in the respective panel databases (see text)

3.4 The dependent variable: labour force participation

When Booth, Jenkins and García-Serrano chose "*actual paid work"* as their dependent variable, they argued: "When comparing men and women, it is important to focus on work versus non-work rather than, say, employment versus unemployment, because of the problems of distinguishing unemployment and inactivity especially for women" (1999:169). The latter problem is certainly relevant in the Belgian setting. Belgium has among the lowest female working rates in the OECD. This observation is commonly related to the generosity of the Belgian unemployment benefit system (Cantillon,1999), which has allowed women with young children to continue to receive benefits even when not (immediately) available to the labour market[14].

Therefore we cannot use the simple measure of labour force participation: "having paid work or being unemployed". Nevertheless, we will take a somewhat more moderate position than Booth, Jenkins and García-Serrano (1999). We will use the strict unemployment definition of the ILO: "unemployed, actively looking for a job and available to the labour market within 15 days". People who participate in the labour force are then either in a paid job or ILO-style unemployed.

Table 3 illustrates the underlying employment status for the Flemish observations of the seventh wave (1998) included in our sub-sample. Table 4 reflects analogous information on the eight wave (1998) of the British data.

The results show that, within the group of labour force participants, working persons constitute by far the most important group. Moreover, it should be noted that in both samples a considerable group of women, who are actively looking for a job and are hence considered ILO-style unemployed, declare "housekeeping" as their major occupation, while none of the men in a comparable situation did so[15]. This observation provides a clear signal that it is not advisable to use self-declared labour market status without further consideration. Additionally, the point made by Booth, Jenkins and García-Serrano (1999) can also be illustrated with our data. If all self-declared unemployed persons were included without checking the ILO

[14] Recent changes in the legislation have partially rectified this problem and pressure by the European Union is likely to cause more cuts in the future, but full correction is politically unfeasible.

requirements, the number of unemployed persons in the Flemish sub-sample would roughly triple. The increase is markedly higher for women (+247 %) than for men (+181 %)[16]. In the British sample the increase is less pronounced, but still considerably higher for women than for men (+156 and +91% respectively).

Table 3 Flemish labour force participants by employment status and gender

	Male		Female	
	#	%	#	%
Working in a regular job	800	*98.1*	616	*93.8*
Having an occasional job or a job of less than 15 hrs/week	1	*0.1*	8	*1.2*
Self-declared unemployed fulfilling ILO-requirements	11	*1.3*	17	*2.6*
Other (training, waiting for a job to start, housework)	4	*0.5*	16	*2.4*
Total	816	*100.0*	657	*100.0*
Overall labour force participation		92.4		75.2

Sample: Prime age individuals living with a partner at the observation moment of wave 7

Table 4 British labour force participants by employment status and gender

	Male		Female	
	#	%	#	%
Employed	1680	*97.1*	1359	*94.6*
Self-declared unemployed fulfilling ILO-requirements	35	*2.0*	9	*0.6*
Other (training, waiting for a job to start, housework)	15	*0.9*	69	*4.8*
Total	1730	*100.0*	1437	*100.0*
Overall labour force participation				72.1

Sample: Prime age individuals living with a partner at the observation moment of wave 8

Table 5 shows the mean of the dependent variable for the complete samples of men and women and for some specific demographic types. At this initial level of description, only very general observations can be made. First, and not surprisingly, male LFP is consistently higher than female LFP. Secondly, divorce experience in its widest definition (anyone of the partners in the couple) does not seem to be of influence and cohabitation is only important to women. All other effects have to be interpreted with great caution because the comparison groups may well differ on decisive characteristics like age. For example, the generalised decrease of labour

[15] Flanders none, Britain one in fifty (35+15) (compared to 38 women of 78 (69+9)).

[16] Unfortunately our data do not allow us to check the official unemployment status of the respondents. Earlier studies and our data suggest that even more women would be included if registered unemployment were the criterion. It can be expected that some women are officially unemployed (receiving unemployment benefits), but are not actively looking for a job or not able to accept a job within 15 days and hence declare ‘housekeeping’ as their major occupation.

force participation among cohabitors due to a mutual divorce experience can probably be explained by differences in the mean age of the groups

Table 5 Mean of Labour Force Participation by Gender and Marital Status

	Flanders		Britain	
Note: selected types of previous table	Women	Men	Women	Men
Married with neither of the spouses previously divorced	.70 (.46)	.92 (.27)	.72 (.45)	.88 (.32)
Married with mutual divorce experience	.64 (.48)	.93 (.26)	.66 (.47)	.78 (.41)
Cohabiting neither of the partners previously divorced	.88 (.33)	.97 (.18)	.84 (.36)	.91 (.28)
Cohabiting with mutual divorce experience	.79 (.41)	.91 (.29)	.77 (.42)	.88 (.32)
Any kind of cohabiting	.86 (.35)	.95 (.21)	.81 (.39)	.90 (.30)
Any kind of divorce experience	.73 (.44)	.92 (.27)	.73 (.45)	.85 (.35)
Total	.71 (.45)	.93 (.26)	.72 (.45)	.88 (.33)

Sample: All panel observations (see Table 2 for numbers in categories)
Standard Deviation between brackets

Table 6 provides some longitudinal insight, reflecting the pattern of labour force participation of the respondents across the waves they participated in. In both samples, stable choices are clearly dominant, with continued presence in the labour market as modal choice among all groups. Other common traits are the lower level of stability among women and a relatively large group of women who have not entered the labour market on any of the observation moments[17].

[17] Comparing British and Flemish results, one might wonder to what extent the distribution of the panel participation biases the results presented. Clearly, the higher persistence of British respondents in the panel gives them more chances to experience a transition. To get a preliminary idea of the magnitude of this bias, we derived a comparable table selecting only respondents who participated the full seven waves in the panel. This procedure generates relatively minor changes that do not alter the British/Flemish comparison (only proportion of continuous choices shown): Britain women 11.9 and 56.4, men 4.8 and 78.9; Flanders women 15.5 and 67.3, men 1.7 and 91.7 Most strikingly the proportion of respondents without any work spell is reduced.

Table 6 Distribution of Labour Force Participation indexes by Gender

(percentages)

Flanders			
	Women	Men	Total
No LFP	25.8	7.9	16.9
LFP below 50%	3.7	1.3	2.5
LFP 50 to 99%	7.8	3.1	5.4
LFP 100%	62.7	87.7	75.2
Number of respondents	1484	1485	2969

Britain			
	Women	Men	Total
No LFP	18.3	9.0	13.6
LFP below 50%	8.2	3.3	5.6
LFP 50 to 99%	15.1	9.6	12.2
LFP 100%	58.5	78.3	68.5
Number of respondents	3010	3050	6060

LFP indices reflect the number of Lfp=1 observations divided by the overall number of observations of the respondent and consequently range from 0 to 100%
Sample: All respondents in the sample (irrespective of the number of waves present)[18]

3.5 The empirical content of the independent variables

In paragraph 3.2 we introduced the empirical specification and the independent variables on a theoretical basis. Yet, much still remains to be said about the actual implementation of this specification. The following table (Table 7) presents the means of all independent variables. The socio-demographic variables age and age squared (not shown) need little explanation. The age squared variable is added to introduce the commonly observed curvilinear pattern of labour force participation when plotted along the age axis.

The educational level is represented by three dummy variables for respectively lower secondary education, higher secondary education and higher education. In Britain the

[18] This general table may be somewhat misleading, because of the inclusion of "singleton respondents" that inflates the proportion of stable choices. Bias is not enormous, however. A corresponding table for all Flemish respondents with at least two observations in our data-set gives 22.1 and 59.2% for the stable choices of women and 5.7 and 87.3% for the corresponding LFP-indices for men. LFP-stability is still the case for more than 80% of both men and women. With overall stability at a lower level, British man and women with stable choices account for respectively 7.5 and 77.9 % (men) and 16.3 and 57.4 % (women). As is the Flemish case, only a small decrease of the proportion of the stable choices is noted.

respective categories correspond to O-level (1), A-level and non-university higher education (2) and degree or teaching qualification (3).

Four dummies for the number and age of the children were already discussed in paragraph 3.2. CHILD1 represents households with one child[19], CHILD2 with two children, CHILD3 with three children and CHILDLOW stands for the presence of at least one young child.

The COHABIT variable is a dummy variable coded 1 for individuals who are cohabiting with their partner and 0 for married individuals. DIVORCE refers to the experience of a divorce or separation in the past.

Table 7 Means of the Independent Variables

		Flanders			Britain	
		Women	Men		Women	Men
AGE		39.34	41.44		40.28	42.4
COHORT1		.39	.32		.39	.31
COHORT2		.32	.33		.27	.28
COHORT3		.22	.25		.24	.26
EDUC1		.22	.21		.25	.18
EDUC2		.32	.31		.27	.38
EDUC3		.33	.34		.13	.14
CHILD1		.23	.23		.19	.19
CHILD2		.26	.26		.21	.21
CHILD3		.12	.12		.10	.10
CHILDLOW		.20	.20		.20	.20
DIVORCE		.06	.08		.17	.16
COHABIT		.08	.08		.11	.11
JDIVORCE		.08	.06		.16	.17
IWAGE	100 BEF	9.09	11.89	£	6.03	8.41
JWAGE	100 BEF	11.87	9.12	£	8.44	6.03
INONLAB	1000 BEF	8.77	4.96	£	77.64	92.50
JNONLAB	1000 BEF	4.99	8.72	£	95.44	78.59
HNONLAB	1000 BEF	1.40	1.80	£	31.14	30.97

Sample: All panel observations

The wage variables IWAGE (wage of the respondent) and JWAGE (wage of the respondent's partner) are instrumented variables. To all persons we assigned a predicted wage, which was calculated for every wave separately applying the usual two-stage Heckman correction

[19] All persons younger than 16 are treated as children in the household.

procedure for selection bias[20]. For non-workers and self-employed people, this wage represents their potential hourly wage when working as an employee, while for employees it engenders their wage, corrected for the likely bias caused by their current labour market participation.

The variables for non-labour income are merely indicators. In Flanders, the figures reflect the monthly average income of the year preceding the interview. For Britain we used the 'estimated usual income for the month of september preceding the interview' as the indicator. This is a derived variable, but usually derivation meant nothing more than conversion to a monthly rate.

In both datasets, INONLAB and JNONLAB (with I for respondent and J for the partner) include social security income that can be expected to last in the future, since respondents are expected to take these amounts into account when deciding about their labour force participation. Consequently, sickness benefits for short periods of illness and birth allowances were excluded from these categories. Additionally, the net income from maintenance payments was included. This resulted in a small proportion of negative values for (predominantly) men. For Flemish observations, the male predominance in maintenance payments and the formerly mentioned female predominance in child benefits, explains the marked difference in INONLAB between men and women. In Britain, however, the relation seems to be the reverse, with average male non-labour income higher than the female counterpart. This result has two major explanations. First, the higher age limit of our British sample leads to the inclusion of quite a number of pension payments, which mostly concern men and are significantly higher for men than for women. Secondly, child benefits are relatively higher in Flanders than in Britain, as shown in **Table 8**. The last line of **Table 8** adds to this difference in benefit levels the slightly higher number of children per household in Flanders as compared to Britain. The child benefit was calculated for an artificial household, consisting of the mean levels of the CHILD dummies in Table 7, assuming that all children were younger than six years old. Thus even without taking the age supplements in account, an average child benefit in Flanders would be about three times higher than in Britain.

[20] The results of these seven estimation procedures are not reported here, but are available from the author on simple request. The probit equation for working as an employee included AGE, AGE2, SEX, CHILD1, CHILD2, CHILD3, EDUC1, EDUC2 and EDUC3 as explanatory variables. The wage equation considered only age, sex and education.

Table 8 Child benefits in the UK and Belgium

	Flanders (=Belgium)	Britain (=UK)
Age limits	0 to 18 years	0 to 16 years
All amounts ECU (July 1996)		
1st child	67	58
2nd child	124	47
Subsequent children	186	47
Aged 6 – 11	+23	No age
Aged 12 – 15	+36	supplements
16 and over (not 1st child)	+44	
Benefit for mean household (see Table 7)	70	26

Note: family credit and other supplementary benefits for specific population groups are not included in this table (e.g. unemployed persons, lone parents, disabled children, orphans)

Source: European Commission (1997b)

Table 9 Percentage of Zero Observations on the Non-labour Income Variables

	Flanders		Britain	
	Women	Men	Women	Men
INONLAB	23.4	51.1	47.8	80.7
JNONLAB	51.1	23.6	80.2	47.6
HNONLAB	86.3	82.3	82.5	82.6

Sample: All panel observations

HNONLAB represents the household's self-reported income from rent and investments. As can be appreciated in Table 9, most households did not report any income in this category.

For reasons of comparability, all amounts of money were adjusted to the 1996 price levels using the official index of consumption prices. Flemish respondents reported all amounts net of taxes, in Britain the reported wages were gross amounts.

3.6 Estimation procedure

Given the panel nature of our sample and the bivariate characteristic of the dependent variable, we chose a panel estimator for limited dependent variables (logit/probit). Following Hsiao (1986) we used a random effects estimator because our sample is of considerable size and, hence, individual effects are likely to conform to the randomness hypothesis.

Following the standard textbook solution, we first estimated our specification using the random effects probit estimator as implemented by Butler and Moffitt (Greene,2000:838-839).

However, the finite sample properties of this estimator are not well known (Guilkey and Murphy,1993) and the estimator is known to be potentially unreliable with large values of 'ρ' and if there is little longitudinal variation for a large proportion of the sample[21] (Booth, Jenkins and García-Serrano,1999). With the latter being particularly true for our data (see Table 6) and estimated values of 'ρ' that range from .73 to .84 , we performed the quadrature check procedure of STATA. This procedure re-estimates the model for 8 and 16 quadrature points instead of the standard 12. In both alternative cases the point estimates of the coefficients diverged from the original estimates by one to more than hundred percent[22]. We therefore decided not to continue analysis with these estimates.

As an alternative, we relied on the corrected pooled probit estimator as proposed by Guilkey and Murphy (1993). Even without the reliability problems of the Butler and Moffitt estimator, we would eventually have chosen this estimator, because it proved to be a better predictor of the actual labour force participation in all cases (see Annex 1).

Additional to the estimator comparison, we experimented with different specifications of the wage variable: a simple linear implementation (iwage), the common logarithmic indicator (ilnwage) and a parabolic formula (iwage + iwage^2). The latter proved to be the best predictor for all sub-samples, irrespective of the estimator (random effects probit or corrected pooled probit). Consequently, we report the estimates of the latter specification[23].

Finally, it should be noted that we excluded the indicators of the educational level from the estimated specification to avoid identification problems introduced by the instrumented wage. In chapter 3.5 we explained how we instrumented an expected wage indicator. Unfortunately at this stage of our analysis, we could not rely on variables other than the classic variables needed for the labour force participation equation. Introducing wage and all its predictors would make the wage variable implicitly redundant.

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[21] The value of earlier estimates of labour force participation in Flanders (Ghysels,1999) diverged greatly with varying numbers of quadrature points, indicating unreliable point estimates.

[22] The results reported here refer to the estimates of British women. Analogous results were found for Flemish men.

[23] Fit measures can be compared to the corresponding measures in the Annexes, where fit measures of the estimates with the natural logarithm of own wage are reported.

4 Estimates, simulations and discussion

In this chapter, we will present the results of the estimation described in the former chapter and relate these results to the hypotheses of chapter 2. The first section discusses the basic hypotheses using the specification of section 3.2. In section 4.2 we will enlarge this basic specification with interaction variables to allow for cohort changes, which in turn lets us analyse the cohort effects of hypotheses # 3 and # 8. In a final section, we will discuss the differences between men and women and between British and Flemish individuals.

4.1 The basic specification

Table 10 reports point estimates and corrected robust standard errors for the four sub-samples. To facilitate interpretation we have also calculated "predicted participation probabilities" for a limited number of typical scenarios (Table 11).

As shown at the bottom of Table 10, the basic specifications provide reasonably good predictors for the actual labour force participation. Generally, the male specifications perform better than the female specifications. Interestingly, this is not to be attributed only to the predominance of lfp=1 observations (labour force participation) among men, because the estimators perform equally better for lfp=0 observations (no labour force participation).

The signs of the coefficients on the standard covariates are as expected from the literature and are in most cases highly significant. Age, for example, has a curvilinear relation with labour force participation. Furthermore, non-labour income and the wage of the partner have the expected negative sign, where significant.

The number and age of children has a negative effect on mothers' labour force participation, though more so in Britain than in Flanders. In the latter the effect is only significant for mothers with three or more children and/or a young child. As Table 11 indicates, the starting position should not be forgotten, however. British women without children are considerably more likely to be in the labour market than their Flemish counterparts. Therefore reductions in the labour force participation for one or two children merely level the initial British premium. Moreover, British women with large families (three or more children) tend to take the lead again, with a higher predicted participation probability than Flemish women.

Table 10 Estimates of the corrected pooled probit regressions

	Flanders Men	Flanders Women	Britain Women	Britain Men
Age	.1422**	.0766**	.0958***	.0739***
	(.0517)	(.0376)	(.0196)	(.0236)
Age2	-.0022***	-.0017***	-.0015***	-.0011***
	(.0006)	(.0005)	(.0002)	(.0003)
Child1	.1874	.0782	-.2392***	-.0877
	(.1404)	(.1018)	(.0679)	(.0915)
Child2	.5236**	-.1361	-.3974***	-.0064
	(.1981)	(.1111)	(.0716)	(.0937)
Child3	.0504	-.4515***	-.5106***	-.1931*
	(.2530)	(.1437)	(.0878)	(.1082)
Childlow	-.1841	-.1788**	-.6577***	-.0603
	(.1560)	(.0844)	(.0554)	(.0730)
Iwage	-.0516	-.0341	.3972***	.2968*
	(.1056)	(.0436)	(.0765)	(.1630)
Iwage2	.0061	.0075***	-.0198***	-.0089
	(.0048)	(.0025)	(.0057)	(.0099)
Jwage	-.0003	-.0052	-.0105	.0082
	(.0233)	(.0174)	(.0153)	(.0226)
Inonlab	-.0457***	-.0210***	-.0026***	-.0024***
	(.0052)	(.0033)	(.0002)	(.0001)
Jnonlab	.0075	-.0053*	-.0008***	-.0012***
	(.0062)	(.0027)	(.0001)	(.0002)
Hnonlab	.0180	.0005	-.0015***	-.0021***
	(.0154)	(.0027)	(.0002)	(.0002)
Divorce	-.1850	-.1506	.0608	-.2605***
	(.3498)	(.1616)	(.0716)	(.0914)
Cohabit	.0852	.2826**	.2671***	-.1529*
	(.2270)	(.1411)	(.0792)	(.0864)
Jdivorce	-.0578	.1102	-.0852	.0367
	(.2855)	(.1745)	(.0696)	(.0907)
Constant	-.1127	.3815	-1.3975***	-.7719***
	(1.1910)	(.7162)	(.3611)	(.6037)
N (observations)	4956	4925	13747	14100
-Log likelihood	640.10	2401.56	6292.44	3015.15
Root mean sqrd error	.1812	.3996	.3839	.2469
Correct predictions	95.94	77.06	79.01	91.48
Correct predictions for non-participation	53.12	38.18	40.54	46.37
Cramer's λ	.4890	.2227	.2568	.4173

*Coefficients are probit estimates Significance levels: *** > 99.5 % ** > 95 % * > 90 %, Robust standard errors(Huber-White) with cluster correction in parentheses*

Table 11 Predicted participation probabilities

	Flemish Men	Women	British Women	Men
References:				
Observed labour force participation	.926	.711	.725	.878
Mean predicted participation probability	.927	.712	.727	.882
No children	.922	.741	.808	.887
One child	.935	.761	.754	.877
One young child	.922	.715	.564	.870
Two children	.954	.706	.714	.887
Three children	.926	.613	.682	.864
Own wage plus one standard deviation	.943	.794	.783	.908
Own wage plus one standard deviation and no children	.940	.816	.853	.912
Own wage plus one standard deviation and three children, at least one young	.931	.662	.547	.887
Married without divorce experience	.928	.708	.718	.890
Divorce experience	.914	.665	.734	.858
Currently cohabiting	.934	.781	.783	.872
Currently cohabiting with divorce experience	.921	.744	.796	.836

All labour force participation probabilities are sample averages of predictions evaluated at every observation using the estimates of the basic specification (Greene,2000:816)

The only exception to this rule is the impact of young children. While the presence of a young child tends to depress the labour force participation of British women very strongly, this is less the case for Flemish women. Consequently, women with a small child are more active in the labour market in Flanders than in Britain. No doubt, this is partially to be attributed to institutional differences like possibilities (unintentionally) offered by the unemployment benefit system in Flanders (see section 3.4) or the almost universal attendance of (semi-)public nursery schools among Flemish three year olds, which both make it easier for mothers not to retire from the labour market (or return to it more quickly).

Fathers' behaviour differs even more between the countries. While Flemish fathers tend to work more when having children, British fathers seem more inclined to reduce their paid work commitment. Both results are consistent with previous findings as summarised in, respectively, Ghysels (1999) and Booth, Jenkins and García-Serrano (1999).

The effect of the respondent's wage is difficult to decipher from the estimates of Table 10, because of the quadratic specification. Before turning to the actual interpretation, it should be noted that the wage is significant in all cases, including the case of Flemish men, when both coefficients are jointly tested[24]. Table 11 allows us to compare across the sub-samples the impact of a one standard deviation increase of "own wage". Flemish women are the only group, which clearly tends to increase labour force participation. In all other sub-samples the impact is relatively moderate, but always positive. Interestingly, among Flemish women the positive impact of a one standard deviation increase of "own wage" on the predicted participation probability is stronger than the negative impact of a young child on the latter (.613 vs. .662), while the opposite is true for British women (.682 vs. .547).

Graph 1 provides a further illustration of the wage effect and its quadratic relation with labour force participation. As in Table 11, the graph presents sample means of changes in one variable (i.e. the instrumented wage, see chapter 3.5) evaluated at all observations using the estimates of the basic specification. We predicted labour force participation for nine points ranging from the respondent's wage minus two standard deviations to the wage plus two standard deviations. In this relevant space[25], wages exhibit a constantly positive slope for all observations.

When comparing Britain to Flanders, we should not forget that wages are net for the Flemish observations and gross for the British. This may partially explain for the different forms of the curves between Flanders and Britain. In net terms, an increase of one standard deviation will be bigger in the lower area than it is in the higher area. Hence, the slope of the British curve can be expected to be less steep in the lower area and steeper in the higher wage area if it would be based on net wages. In that way, it would approach the Flemish result more than it does with the current data.

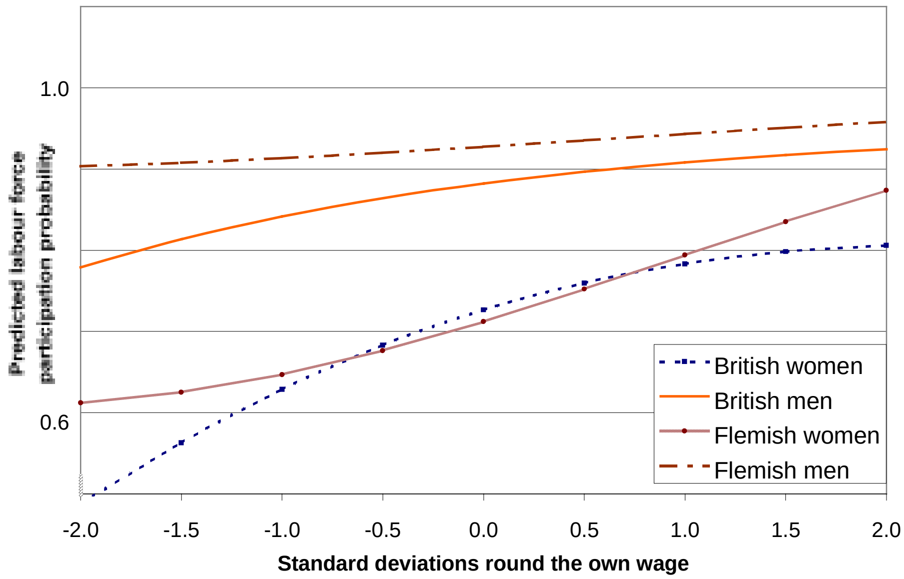
Considering the gender gap, it should be noted that the instrumented wage variable is strongly linked with the respondents' educational level. Hence the graph somewhat confirms earlier

[24] Wald tests on the hypotheses that both coefficients would jointly be zero are consistently rejected. For Flemish men the corresponding value of the χ^2 indicates a joint significance level of 99.73% (χ^2 = 11.83 with 2 d.f.). For British men the χ^2 value is 45.96

[25] See Table 17 in the Annexes for indications of the relevance of this interval. Dispersion is generally larger for the Flemish data than for the British data, but the two standard deviation space covers about 90% of the observations.

observations that the gender gap is smaller among the highly skilled than among the lowly skilled.

Graph 1 Predicted labour force participation varying along the wage axis



4.2 Marital status and history: basic versus interacted specification

Turning to the variables of interest for this paper, the results in Table 10 do not look particularly promising for the divorce variable. Contrary to our hypotheses (#1 and #2), a divorce experience does not increase the labour force participation of women and the divorce experience of the male partner has no effect either. For men the results are somewhat more in line with the hypothesis (#4) with Flemish men insensitive to divorce experience, but British men somewhat stretch the hypothesis, showing lower degrees of labour force attachment with divorce experience than without it.

The estimates of the cohabitation coefficients largely confirm our hypotheses (#6 and #7) with small to non-existent effects for men and significantly positive effects for women. Flemish women seem to be particularly responsive to the additional risks that cohabitation implies, since they increase their labour force participation by about 7 percentage points when living in a cohabiting union compared to a regular marriage.

Of course these results are prone to be biased by cohort effects. As explained in chapter 2.2, people's attitudes towards marriage and cohabitation have changed considerably during the past fifty years. Though some of this bias may be controlled for by the age variables, we believe additional cohort indicators may further clarify the picture. Moreover, cohort indicators are needed to address the hypotheses #3 and #8, which explicitly refer to cohort effects.

Because we have no hypotheses on cohort effects apart from those related to the marital status and history, we introduce cohort effects through three levels of interaction. We divided the sample in four groups: respondents born after 1959 (COHORT1), those born from 1950 to 1959 (COHORT2), from 1940 to 1949 (COHORT3) and finally the individuals born before 1940 (contrast category)[26]. We then interacted the cohort dummies with the DIVORCE and COHABIT variables. This way, we added six new variables to the specification and obtained new corrected pooled probit estimates for the interacted specification[27].

Before discussing the results of these new estimates, we should have a look at Table 12. This table shows the occurrence of cohabitation and divorce in the different cohorts. Given the recent increase in cohabitation and the typical pre-marital nature of cohabitation in Flanders and Britain, it is not surprising to see that cohabitation is far more common in the youngest cohort as compared to the other cohorts.

Table 12 Cohort decomposition of cohabitation and divorce experience

Within group percentages	Flanders Women	Men	Britain Women	Men
Currently Cohabiting				
Cohort1	13.0	13.6	18.8	21.4
Cohort2	5.9	7.7	6.8	9.2
Cohort3	2.9	5.6	5.5	5.9
Cohort4	4.4	2.4	2.6	3.4
With Personal Divorce Experience				
Cohort1	5.5	3.8	12.7	8.1
Cohort2	7.6	11.0	21.3	20.9
Cohort3	6.4	2.8	19.4	21.3
Cohort4	3.5	6.9	11.4	16.5

For a definition of cohort samples see text

[26] See Table 19 in Annexes for the actual distribution of the observations among the cohort groups.
[27] While estimating the specification for Flemish men, Stata dropped the interaction variable relating to divorce in the first cohort because all observations with these characteristics were active in the labour market. We report results omitting the coefficient.

Divorce is more equally spread, with lower levels in the youngest cohorts because of the obvious duration effect (marriage has not even occurred for many members of the youngest cohort) and lower levels in the oldest cohort because of cultural changes in the post-war cohorts.

In Table 13 we report the estimates of the interacted specifications. We do not repeat all estimates because the addition of the interaction variables does hardly change the original estimates of the other variables. All significant coefficients remain so and no other become significant. To make interpretation easier we report marginal effects, which indicate percentage point changes in the predicted labour force probabilities.

As the bottom lines in Table 13 indicate, the fit of the estimates hardly improves by the addition of the interaction variables. On the other hand, Wald tests support joint significance of the variables related to cohabitation and, in the case of British men, the divorce variables.

The results for Flemish and British women coincide with our hypothesis (# 8). In the youngest cohort the joint coefficient of cohabitation is very small and non-significant. It grows a little in the second cohort to reach its highest level in the third cohort. Contrary to our hypothesis, the coefficient for the fourth cohort is negative, though insignificant. The picture for men is even less clear. British cohabiting men tend to be less active in the labour market than their married counterparts, except in the eldest cohort. This effect is most marked (and significant) in the youngest cohort. For Flemish cohabiting men labour force participation probabilities are somewhat higher in the second cohort, but exhibit a surprisingly large drop in the third cohort.

The estimates of divorce experience variables hardly change when introducing cohort effects. Only for British men a significant divorce effect is noted and it is continuously negative, though with some variation along the cohorts.

Summarising, we would conclude that a past divorce does not seem to exert any influence on present labour force participation among repartnered individuals. This empirical observation contradicts our hypotheses, especially for women. Additionally, the only significant effect found has an unexpected sign. British men with divorce experience are less likely to be in the labour market than their counterparts without a divorce experience. Future research should reveal whether this effect has anything to do with the hypothesised enhanced fatherhood feelings of divorced men.

Table 13 Marginal effects of the corrected pooled probit regressions

	Flanders				Britain			
	Men		Women		Women		Men	
	Basic	Inter	Basic	Inter	Basic	Inter	Basic	Inter
Cohabit	.0038	.0060	.0833*	-.1498	.0757*	-.0459	-.0191*	.0337
	(.0094)	(.0162)	(.0379)	(.1347)	(.0205)	(.1586)	(.0117)	(.0240)
Cohabit*Cohort1		.0091		.1626		.1030		-.1119*
		(.0155)		(.0740)		(.1124)		(.0854)
Cohabit*Cohort2		.0195*		.1838*		.1282		-.0814
		(.0034)		(.0601)		(.0948)		(.0831)
Cohabit*Cohort3		-.1918*		.2023*		.1325		-.0849
		(.1539)		(.0537)		(.0981)		(.0877)
Joint significance		*99.99*		*91.55*		*99.57*		*93.65*
Divorce	-.0103	-.0130	-.0504	.1299	.0184	.0516	-.0341*	-.0364
	(.0230)	(.0231)	(.0561)	(.0977)	(.0214)	(.0597)	(.0136)	(.0283)
Divorce*Cohort1		*dropped*		-.1073		-.0314		.0082
				(.1787)		(.0773)		(.0233)
Divorce*Cohort2		-.0269		-.2048		-.0213		-.0367
		(.0484)		(.1986)		(.0761)		(.0337)
Divorce*Cohort3		.0177*		-.4742*		-.0764		.0297
		(.0033)		(.1815)		(.0860)		(.0165)
Joint significance				*87.73*		*30.93*		*99.99*
Jdivorce	-.0029	-.0038	.0343	.0434	-.0267	-.0264	.0041	-.0013
	(.0150)	(.0133)	(.0525)	(.0533)	(.0223)	(.0225)	(.0099)	(.0100)
N (observations)	4956	4896	4925	4925	13747	13747	14100	14100
-Log likelihood	640.1	626.9	2401.6	2391.0	6292.4	6288.1	3015.1	2992.7
Root mean sqrd.error	.1812	.1811	.3996	.3986	.3839	.3838	.2469	.2456
Correct predictions	95.94	96.00	77.06	77.12	79.01	79.05	91.48	91.67
Correct predictions for non-participation	53.12	53.93	38.18	38.81	40.54	40.68	46.37	47.53
Cramer's λ	.4890	.4964	.2227	.2268	.2568	.2572	.4173	.4222

Coefficients are marginal effects calculated at the sample means comparing 0 to 1 states of the variable considered (Greene,2000:817) Robust standard errors corrected for clustering between parentheses (calculated using the delta method, Greene,2000:824) Joint significance refers to a Wald test of coefficients being jointly zero and is derived of the χ^2 distribution with 4 degrees of freedom
** indicates significance of the individual coefficient at > 90%*

Regarding cohabitation, our hypotheses seem to hold true for women. Both Flemish and British cohabiting women are more likely to be in the labour market than their married counterparts. Moreover, the effect is diminishing in age. As expected, the relation is non-existent in the youngest cohort where cohabitation has become the norm, but is significant in the second and third cohort, where cohabitation can be expected to engender a type of selection. For men the results are somewhat puzzling. Among British men we observed no relation between cohabitation and labour force participation activity, except for the youngest cohort where a slightly negative relation was detected. A possible explanation to be verified in future research, is an added cohabitation effect for fathers. The results for Flemish men are quite striking because of considerable cohort differences and relatively strong and significant effects, especially among the third cohort, where cohabitation is predicted to lead to a 19 percent drop in the labour force participation probability.

4.3 A decomposition exercise

Previous tables and Table 11 in particular have provided ample evidence of the differences in labour force participation between the four sub-samples we are studying here. In general, Flemish men are more likely to be in the labour market than their British counterparts and this result is not altered when studying more specific groups. At first sight, women may seem to have very similar labour force participation probabilities in Britain and Flanders, but a closer look at specific situations also reveals marked differences between these two populations.

To get a clearer picture of the nature of the differences, we adopted the decomposition formula for binary variable regressions proposed by Gomulka and Stern (1990). This formula allows us to decompose the difference in labour force participation probability into a first component reflecting differences in the estimated coefficients and a second component reflecting differences in the observed characteristics of the respondents.

Table 14 list results for comparisons in both directions. As in previous studies[28], the male female differential is mainly to be attributed to differences in the coefficients. Both in Flanders and Britain more than two thirds of the observed labour force participation gap is to be explained by so-called behavioural differences.

[28] See for example: Booth, Jenkins and García-Serrano (1999:184-185). The latter authors analyse differentials in the probability of being in work and observe an even higher proportion of the gap being explained by differences in the coefficients.

Table 14 Decomposition of differences in predicted participation probabilities

	Total Gap	Coefficients #	Coefficients %	Characteristics #	Characteristics %	Mean PPP [a] $\overline{P}(\hat{\alpha}^r, X^l)$
Flanders						
Men-women	.2153	.1481	68.79	.0672	31.21	.77939
Women-men	-.2153	-.2077	96.45	-.0076	3.55	.91985
Britain						
Men-women	.1555	.1075	69.10	.0481	30.90	.77468
Women-men	-.1555	-.1208	77.63	-.0348	22.37	.84737
Men						
Britain-Flanders	-.0453	.0809	-178.45	-.1262	278.45	.80128
Flanders-Britain	.0453	-.0597	-131.61	.1050	231.61	.98715
Women						
Britain-Flanders	.0144	.3497	2424.14	-.3352	-2324.14	.37695
Flanders-Britain	-.0144	-.1516	1051.22	.1372	-951.22	.86382

[a] *Gomulka and Stern (1990) elaborate the following decomposition formula for predicted labour force participation probabilities (their equation #8):*

$$\hat{y}^1 - \hat{y}^0 = \left\{\overline{P}(\hat{\alpha}^1, X^1) - \overline{P}(\hat{\alpha}^0, X^1)\right\} + \left\{\overline{P}(\hat{\alpha}^0, X^1) - \overline{P}(\hat{\alpha}^0, X^0)\right\}$$

We adapted their formula replacing the indexes 1 and 0 by l (left) and r (right)
For the non-mixed mean predicted participation probabilities, see Table 11

The British-Flemish decomposition supports our earlier conjecture that the apparent similarity of British and Flemish individuals is only hiding underlying differences in both characteristics and coefficients. Indeed, the decomposition indicates that the actually observed differences are the results of considerably larger gaps due to both sources, which run in opposite directions and offset each other. This is most clearly seen among women. Applying Flemish coefficients to British women would seriously reduce the latter's labour force participation, but the opposite applies to Flemish women, who would be active in the labour market to a much higher degree than they actually are, if they would "behave" like British women.

Table 15 provides another illustration of gender and cross-country differences showing predicted labour force participation probabilities for some hypothetical individuals. Basically, we compare three age groups with their typical family composition and median values for the monetary variables. For every age group, we elaborated a type with rather Flemish monetary characteristics (F-types) and a type that represents the British situation (B-types). These differences refer to both the wage variables and the non-labour income, here limited to the child benefit[29]. As a fourth scenario, we added a typical cohabiting person. This way, horizontal comparisons in the table highlight differences in coefficients and vertical comparison between F and B alternatives show the effect of differences in (typified) characteristics.

Table 15 Simulations of the likelihood of labour force participation

	Flanders		Britain	
	Men	Women	Women	Men
Type 1: Married, Age 30, 2 young children				
F: woman's wage £ 4 and non-labour income £ 250, man's wage £ 5	.998	.799	.356	.911
B: woman's wage £ 5 and non-labour income £ 140	.999	.891	.549	.954
Type 2: Married, Age 40, 2 children between 6 and 11				
F: woman's wage £ 5 and non-labour income £ 310	.999	.790	.592	.932
B: woman's wage £ 6 and non-labour income £ 140, man's wage £ 8	1.000	.909	.799	.979
Type 3: Married, Age 50, 1 child of 16				
F: man's wage £ 8, woman's non-labour income £ 90	.997	.804	.761	.959
B: man's wage £ 10, woman's non-labour income £ 80	1.000	.801	.769	.978
Type 4: Cohabiting, Age 30, 1 young child,				
F: woman's wage £ 4 and non-labour income £ 90, man's wage £ 5	.994	.935	.681	.904
B: woman's wages £ 5 and non-labour income £ 80	.997	.964	.759	.936

Unless otherwise mentioned all types share the following characteristics:
all non-labour income categories at £ 0 and both wages at £ 6
Simulations are based on estimates of the basic specification

The table allows several observations. Regarding men, Table 15 confirms first our earlier observation that they, and especially Flemish men, are relatively insensitive to changes in the variables of our specification. Secondly, men show consistently high participation probabilities, with Flemish men somewhat above their British counterparts.

[29] We used the data of **Table 8** to calculate the typical child benefit for the child configuration of every type (we converted the ECU-amounts of **Table 8** by the December 1996 exchange rate of £1=1.3192 ECU)

Considering women, there are many more fluctuations. Apart from the effect of children that we documented earlier, women's labour force participation seems to vary quite strongly with the monetary covariates. For Flemish women, an increase of the own wage with £1 always produces higher participation probabilities, adding to the work incentive of the decrease of non-labour income implied by the F to B shift. The same observation applies to British women, but to an even greater degree. While a typical F to B shift increases the Flemish labour force participation probability by about ten percentage points, it produces an increase of twenty percentage points among British women. Moreover, a comparison of type 1 to type 4 shifts suggests that this large increases are mainly to be attributed to the effect of the non-labour income, because in the absence of a big change in the latter category, no great alterations of the participation probability are observed either.

Summarising, we could maintain our conclusion of Table 14. Both differences in the characteristics of the samples and behavioural differences are clearly present. For women, these effects generally work in opposite directions, thus producing the overall similarity in female labour force participation observed in Table 5.

5 Conclusions

In this paper we studied the labour force participation of individuals living with a partner. Within this population we differentiated between people with and without a divorce experience and between persons who are legally married and those who are cohabiting. We hypothesised that individuals would discount for their separation risk when deciding on their labour force participation. Hence, we conjectured that both divorced persons and cohabiting individuals would show higher than average labour market participation because they know that they run higher risks of experiencing a separation in the future, which will make them the sole income provider of their household.

However, we also assumed that the risk discounting effect would be minor to non-existent for population groups who tend to be highly attached to the labour market in general. Close to the margin only a very small additional effect can be expected. This is particularly the case for men, but also for younger women.

In effect, we observed a quite pronounced increase of labour force participation among cohabiting women in both Flanders and Britain, compared to married women. Additionally, the effect was seen to be diminishing in age, i.e. in the youngest cohort hardly any effect was noted. Equally in line with our hypothesis, men's behaviour is not affected by the legal status of their union.

Regarding divorce, the hypotheses are not so successfully confirmed in our data. For women, a divorce experience does not increase labour force participation. For men the results are somewhat more in line with the hypotheses with Flemish men insensitive to divorce experience. However, British men somewhat stretch the hypothesis, showing lower degrees of labour force attachment with divorce experience than without it.

When comparing the different groups across the gender divide, the usual result is confirmed giving a larger weight to differences in the coefficients ('behavioural differences') than to differences in the characteristics of men and women when explaining their different degrees of labour force participation. Cross-national comparisons offer less clear results, but argue against quick generalisations. For example, it is shown that the comparable degree of labour force participation among Flemish and British women is the result of a different process starting from different characteristics going through different types of behaviour.

More work is clearly needed. In future research, we will address some of the working hypotheses that were not put in an empirical perspective in this paper, more specifically those regarding fathers' hours of work and the impact of children on divorced fathers' time use. Additionally it may be interesting to add more historical perspective to the analysis using data on both the marital and the labour history of the respondents. The intertwining of both may shed light on the questions of causal direction we entirely forego in the empirical part of this paper. If the likelihood of divorce is strongly linked to an individual's labour force status, our estimates may show some endogeneity bias, which we should control for in the future.

For now, we can conclude that cohabitation tends to increase female labour force participation, but that this effect is diminishing over time. At the other hand, a divorce is shown not to have any effect on women's labour once they have started living with a new partner.

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7 Annexes

Table 16 An overview of estimator performance

	-Log likelihood[a]	Root mean sqrd. Error	Percent correctly predicted (.50 cut-off) All	Lfp=0	Cramer's λ
Flemish women					
Pooled probit with corrected standard errors	2418.89	.4013	76.71	37.26	.2161
Random effects probit (Butler and Moffitt)	1623.61	.4368	75.70	28.91	.2387
Flemish men					
Pooled probit with corrected standard errors	642.98	.1817	96.00	53.93	.4876
Random effects probit (Butler and Moffitt)	518.45	.1838	95.72	47.15	.4458
British women					
Pooled probit with corrected standard errors	6297.58	.3840	79.06	40.33	.2561
Random effects probit (Butler and Moffitt)	4443.47	.4073	78.21	33.86	.2802
British men					
Pooled probit with corrected standard errors	3016.03	.2469	91.47	46.31	.4172
Random effects probit (Butler and Moffitt)	2334.72	.2652	91.18	40.45	.3928

[a]Log likelihood values are given as basic reference, but are not comparable because the pooled probit with clusters does not produce real maximum likelihood results (pseudo-maximum likelihood)
Samples: all observations in the respective panel databases (see text)
The results refer to the basic specification with the natural logarithm of the own wage(ilnwage)

Table 17 Basic descriptive statistics of the instrumented own wage (IWAGE)

	Flemish Men	Flemish Women	British Men	British Women
Mean	11.89	9.09	8.41	6.03
Standard deviation	2.62	2.83	1.76	1.78
Percentage of sample within range				
[mean -1 s.d., mean +1 s.d.]	67.66	67.47	74.50	71.71
[mean -2 s.d., mean +2 s.d.]	96.61	95.65	92.35	91.19

Sample: all panel observations

Table 18 Descriptive statistics of the standardised monetary variables

	Flanders Men	Flanders Women	Britain Women	Britain Men
IWAGE	5.16	3.94	6.03	8.41
	(1.14)	(1.22)	(1.78)	(1.76)
INONLAB	93.24	164.72	77.64	92.50
	(272.46)	(203.50)	(143.79)	(254.49)
JWAGE	3.96	5.15	8.45	6.03
	(1.23)	(1.14)	(1.76)	(1.77)
JNONLAB	163.81	93.74	95.44	78.59
	(202.93)	(273.45)	(259.27)	(145.22)
HNONLAB	33.80	26.27	31.14	30.97
	(181.86)	(186.02)	(104.66)	(104.20)

Standard deviations in parentheses This table complements Table 7 All sample observations All amounts in Belgian Franc were converted to Pound Sterling using the December 1996 exchange rate (£ 1 = 53.244 BEF) Additionally Flemish wages were converted to monthly amounts (divided by 4.33)

Table 19 The cohort distribution of the samples

	Flanders Men		Flanders Women		Britain Women		Britain Men	
Cohorts	N	%	N	%	N	%	N	%
1: after 1959	1565	*31.6*	1933	*39.2*	5392	*39.2*	4393	*31.2*
2: 1950-1959	1636	*33.0*	1587	*32.2*	3665	*26.7*	3993	*28.3*
3: 1940-1949	1248	*25.1*	1090	*22.1*	3324	*24.2*	3626	*25.7*
4: before 1940	507	*10.2*	315	*6.4*	1366	*9.9*	2088	*14.8*
	4965	*100.0*	4925	*100.0*	13747	*100.0*	14100	*100.0*