FAMILY MIGRATION AND WOMEN'S LABOUR MARKET STATUS IN BRITAIN: THE EFFECTS OF STATE DEPENDENCE AND GEOGRAPHY

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Abstract

Numerous studies show that women's labour market status is influenced negatively by family migration, with women who move long distances with their partners being less likely to be employed in the labour market than otherwise equivalent women. Using longitudinal data from the British Household Panel Study (BHPS) we focus on employment status and investigate four issues which have received relatively little attention in this literature. First, instead of relying on the distance moved to distinguish probable employment-related migration, we use data on the reason for moving, which allows us to separate employment-related moves stimulated by the man or the woman from other types of moves. We compare our results with a standard distance-based approach. Second, we consider the role of state dependence by using models which allow us to examine the influence of women's employment status prior to moving. This has been ignored in previous studies. Third, we compare the results for the majority of women who have a lower occupational status than their partner with those who have a higher occupational score, with the expectation that moving will have an insignificant effect on women in the latter category. Fourth, we test whether women's labour market status is influenced by the metropolitan/non-metropolitan nature of the destination and the employment rate for women in the local area relative to the national average. Our results show, first, that moving for the sake of the man's job does indeed have a dramatic negative effect on the woman's subsequent employment status, although moving for other reasons also increases the risk of women being out of work. Second, we find that these effects are apparent for most women who were employed in the year prior to the move. However, women who were not employed in the previous year tended to benefit from family migration, even if the move was stimulated by the man's job. State dependence is a crucial issue. Third, we show that while women with a higher occupational status than their partner are generally more likely to be employed than other women, the negative effects of family migration still exist. Fourth, while the metropolitan/non-metropolitan nature of the geographical location has no effect on women's employment status, women are less likely to be unemployed or economically inactive in areas where women's employment rates are high relative to the national average.

Keywords: Family migration; employment status; random-effects dynamic probit models; state dependence; geography; British Household Panel Study (BHPS).

Introduction

The influence of ‘family migration’, or the long-distance move of partnered individuals, on employment status has attracted a considerable literature over the past 30 or so years. The general consensus is that families are more likely to move in support of the man’s career and that women’s employment status is likely to suffer as a result. According to human capital theory, families weigh the benefits of gains associated with a move on behalf of one person’s career against the negative effects of disrupting the partner’s employment (Becker 1974). Although a genderless theory, moves tend to be made to support the man’s career more often than the women’s and, as a result, women are more likely to be ‘trailing spouses’ or ‘tied migrants’.

Early studies confirmed this effect (Lichter 1980, 1982, Long 1974, Mincer 1978, Morrison and Lichter 1988, Sjaastad 1962, Spitze 1984) and, more recently, Boyle *et al.* (1999a) used comparative, cross-national data for GB and the US at the beginning of the 1990s and showed that women were more likely to be out of work following family migration (Boyle *et al.* 1999b, 2001). This result was remarkably consistent in GB and the US (Boyle *et al.* 2002), even controlling for motherhood status (Boyle *et al.* 2003) and the relative occupational status of the partners (Boyle *et al.* 1999c). Other studies confirm these broad conclusions (Bailey and Cooke 1998, Cooke 2001, Shihadeh 1991, Jacobsen and Levin 1997, 2000, Smits 1999), although others cast doubt over the strength of these findings. For example, Clark and Withers (2002) argue that while women’s labour force participation is disrupted by family migration, these effects tend to be short-lived in the US, although Clark and Huang (2006) show that the disruptive effects are longer in Britain.

This paper extends this body of research in four distinct ways. First, studies in the past have tended to use the distance moved, or the fact that a move crosses an administrative boundary, as a surrogate measure indicating that the family has probably moved for employment-related reasons. Shorter distance moves are usually assumed to be more likely to be housing related. We use longitudinal data drawn from the British Household Panel Study (BHPS) which includes information about the reason for the move, and we are able to compare this to a more common approach based on distance moved. In particular, we are able to separate employment-related moves, stimulated by the man, the woman or both, from other types of moves.

Second, we consider the role of state dependence which has been ignored in family migration studies to date. Heckman (2001: 706) notes that:

> “a frequently noted empirical regularity in the analysis of unemployment data is that those who were unemployed in the past or have worked in the past are more likely to be unemployed (or working) in the future”.

and this has been shown to be the case in a number of British studies (e.g. Narendranathan and Elias 1993, Arulampalam 2000). From a human capital perspective, state dependence in unemployment occurs because of the deterioration of existing human capital during an unemployment spell as well as from the non-accumulation of new human capital during this period (Stewart 2005). Ignoring this issue is likely to lead to biased models and under-estimated standard errors. Thus, we fit appropriate panel models which adjust for unobserved heterogeneity and initial conditions. We expect that moving will have a particularly negative effect on women who were employed previously.

Third, we compare the results for the majority of women who have a lower occupational status than their partner with those who have a higher occupational score. Assuming that family migration decisions are gender-neutral, as implied by the human capital model, we would expect that moving should have an insignificant, or possibly even positive, effect on women in the latter category.

Finally, although Mincer (1978) argued early on that the choice of geographical location was likely to play a major role in family decisions about relocation, particularly in relation to the potential employment opportunities they provide, few studies seem to have considered this issue. We focus on two aspects of the geographical context, using a classification of places based on their metropolitan or non-metropolitan status, and a continuous variable which compares women’s employment rates in local areas relative to the national average.

In combination, this study provides a unique insight into the influence of family migration on women’s employment status, using up-to-date longitudinal data which allow us to explore transitions in and out of different states. In the remainder of the paper we describe the data in detail, the regression methods that we use and the results, before discussing the implications of our findings.

Data

The data are drawn from the British Household Panel Study (BHPS) which has been collecting data on a nationally representative sample of households since 1991. We extracted data from waves 1-13 (A-M) on a restricted sample of 4,491 women aged between 16-64 in married or cohabiting partnerships. Women may have joined the panel since 1991, or could have left the panel before 2003. Women whose relationship ended were dropped, but they returned to the sample if they began a new partnership. The sample contains 29,349 observations and a range of individual- and area-level variables were included in the analysis.

Our outcome variable compares women who were unemployed or economically inactive (about one third of the observations) with those in employment. We control for a range of explanatory variables, but we focus particularly on migration status. Two variables are used. The first compares those who moved over 30 kilometres between t-1 and t to those who moved only short distances or who did not migrate at all. This is a common strategy used to distinguish probable employment-related moves from moves stimulated by housing or other reasons and this has been used frequently in the past, in the absence of further information on the reason for the move. Fortunately, the BHPS includes data on the reason for moving and we therefore constructed a second variable which compared non-movers with those who moved for various reasons. Figure 1 plots the average distance moved by age and the reason for the move from the BHPS data. It is clear that, on average, employment related moves tend to be longer than 30kms (the distance cut-off we have chosen). However, moves into college for younger adults and moves for family and environmental reasons especially also tend to be longer than 30 kms for older adults. This figure makes it clear that using an arbitrary distance cut-off to distinguish between employment-related and other types of moves is inadequate. In our analysis, we use the reason for the move to divide migrants into those who moved for the man's job, those who moved for the woman's, or both jobs and those who moved for other reasons. In particular, moving for the man's career is hypothesised to have a negative impact on women's employment status, while we would expect that women who moved for either their job or for both jobs would be more likely to be in employment following the move.

We are also interested in state dependence – an issue that seems to have been ignored in the family migration literature – expecting that women who were out of employment at t-1 are also more likely to be out of work at time t. Thus, we use the lag of employment status in our models, but are mindful that the inclusion of such a variable is likely to lead to bias in conventional panel regression models. Hence we adopt a dynamic modelling approach which addresses this problem.

We also compare the employment outcomes by occupational status to test whether the effects of family migration are consistent for women who have higher and lower occupational score than their partners (see Boyle 1999c). We use the Cambridge score (Prandy 2000) which is a gender-specific occupation-based measure of social stratification. This novel approach to occupational classification stresses the social distance between occupations, by comparing the occupations of married partners and friends. Because it is firmly grounded in actual social behaviour, it has been argued to be more theoretically valid than other measures of social stratification. And, a particularly attractive aspect is that it provides separate scores for men and women's occupations – working in the same occupation does not necessarily command the same prestige for men and women. We compare the outcomes of family migration for women who have higher Cambridge scores than their partners with women who have lower Cambridge scores. We expect that the negative effects of family migration will be greater for the latter group.

Finally, we also consider the geographical context in which family migration is played out. We might imagine that certain locations provide better opportunities for women to find employment following a move than others. First, we contrast London with the other metropolitan areas, and the remainder of southern and northern Britain. We expect that women are more likely to be employed in the labour market in the large urban centres. We consider this issue more directly by also including a measure of women's employment rates in the Local Authority District in which they reside. Thus, we expect women to be more likely to be in work in areas where women's employment rates are high compared to the national average.

Methods

We have panel data with repeated observations for the same woman and our outcome measure is a binary variable distinguishing between partnered women who are, or are not, in employment at time t. We therefore use random effects probit models, appropriate for modelling such panel data. However, we are particularly interested in whether employment status in the previous wave (t-1) has an influence on the outcome at time t (state dependence), but the inclusion of a lagged y variable is known to introduce bias. We therefore fit Heckman-type (1981) random-effects dynamic probit models which control for initial conditions using an approach developed by Stewart (2006):

$$y_{it} = \gamma y_{it-1} + x'_{it}\beta + \alpha_i + u_{it} (i = 1,..., N; t = 2...,T) \qquad (1)$$

where y_{it} is the binary indicator variable for employment, x_{it} is a vector of explanatory variables, α_i are (unobserved) individual-specific random effects, and u_{it} are assumed to be distributed N(0, σ^2_u). The estimation of this model requires an assumption about the initial conditions, y_{i1} and their relationship with α_i. The initial conditions can be assumed to be exogenous if the start of the process coincided with the beginning of the observation period for each person, but this is not the case in these panel data. Because we would expect the initial conditions to be correlated with α_i, we would expect γ to be overestimated, leading to an overstated impact of state dependence (Chay and Hyslop 2000). We therefore adopt an approach to the initial conditions problem which involves a linearised reduced form equation for the initial period:

$$y_{i1} = z'_{i1}\pi + n_i (i = 1,..., N) \qquad (2)$$

where z_{i1} is a vector of exogenous instruments and includes x_{i1}, and n_i is correlated with α_i, but uncorrelated with u_{it} for $t \geq 2$. In our case our instrument variables, which are significant in a simple probit model fitted for t=1, but insignificant in a model for $t \geq 2$, were household type, occupational diversity and occupational penetration, the latter two being variables constructed for the Local Authority District in which the respondent resided at a particular wave. Below, we compare the results from the standard and the dynamic random effects models.

Results

Table 1 demonstrates that there is considerable state dependence in unemployment / economic inactivity in the raw data. Thus, while only 9% of previously employed women who did not move were unemployed or economically inactive, 27% of previously employed women who moved a long distance were out of work. However, the comparable figures for women who were not employed at t-1 were 82% and 73%. In the raw data at least moving appears to increase the likelihood that a woman is employed if she was out of work prior to the move, and this demonstrates the importance of controlling for prior employment characteristics in our analysis. Family migration may not have negative effects for all women – it may even have positive effects for some.

Table 2 presents the results from five models, two of which are standard random-effects probit models and three of which are random-effects dynamic probit models. Models 1, 2 and 3 use the distance-based definition of migration, while models 4 and 5 used the reason for the move. Figures 2-5 provide the calculated probabilities for different population sub-groups for Models 1-4 and Figures 6 and 7 relate to Model 5.

Model 1 and Figure 2 show that women who moved long distances are significantly more likely to be unemployed or economically inactive than non-migrants or those who moved short distances. This confirms the findings from previous studies.

Model 2 includes the interaction with the lagged employment variable. Migration remains positive and significant and the lagged variable is negative and significant. Figure 3 shows that women who were employed at t-1 were considerably less likely to be employed following a long distance move. However, consistent with the results for the raw data (Table 1), migrant women who were unemployed at t-1 were less likely to be out of work than women who were unemployed at t-1 and did not migrate a long distance. However, this model is a standard random-effects probit and it is likely to be biased because of the incorporation of the lagged y variable. Model 3 and Figure 4 provides the same model fitted as a random-effects dynamic probit. Although the probability of unemployment or economic inactivity has fallen for women who were not working at t-1, the broad conclusions remain the same – migration has a negative effect on those who were in employment at t-1, but a positive effect on those who were not employed at t-1.

Model 4 presents results from a random-effects dynamic probit model which uses the reason for the move and also includes the lagged y variable. The probabilities are calculated in Figure 5. Interestingly, the probability of being out of work is reasonably similar for all women who were employed at t-1, except for those who moved for the man's job. This group had around twice the probability of being out of work compared to non-migrant women who were employed at t-1. On the other hand, for women who were not working at t-1 the probability of being out of work at time t is similar for all women, except those who moved for the woman's / both jobs. This group had a much lower probability of being out of work.

The results for women who were and were not employed at t-1 were very similar for those who moved for the man's or the women's / both jobs, while the results were considerably different for non-migrant women and women who moved for other reasons. Also, of the women who were employed at t-1, non-migrants had the lowest probability of being out of work; for women who were not employed at t-1, non-migrants had the highest probability of being out of work.

Finally, Model 5 and Figures 6 and 7 compare the results for women who had lower Cambridge scores than their partners, with women who had higher scores (about 42% of the women) at t-1 (those with no occupation coded were given score of zero). The results demonstrate some interesting differences in the chances of being out of work. Women who were employed at t-1 had virtually identical probabilities of unemployment or economic inactivity regardless of whether their Cambridge score was higher or lower than their partner's. This pattern was similar to that in Figure 5, with women who moved for the man's job being much more likely to be out of work. However, for women who were out of work at t-1, the results are significantly different depending on whether they had a higher or lower Cambridge score than their partner. Those with a higher score had a considerably lower probability of being out of work, regardless of their migrant status. Even so, the relative pattern across migration categories remained similar and the interaction between migrant status and Cambridge score was not significant.

We were also interested in the results for structural, geographical variables. In every model the region variable, which distinguished between London, the other Metropolitan areas and the rest of the South and North was insignificant. On the other hand, the women's local employment rate was significant in every model – women are more likely to be working if they live in areas where women's employment rates are high, even controlling for a wide range of individual-level factors. Note that the interaction between migration and this geographical variable was not significant (results not provided).

Discussion

We considered four broad questions in this analysis. First, we compared the use of a standard distance moved variable with a reason for moving variable. While the results for the former confirmed previous studies, the latter provides a much clearer impression of the effects of family migration on women's employment status. Initial results (not displayed) suggested, as expected, that moving for the man's job had a negative effect and moving for the woman's or both jobs had a positive effect on employment.

Our second aim, however, was to consider the influence of state dependence and we found it had a significant role. Employed women seem to suffer from migration, especially if it is for the man's job. Women who are out of work are more likely to be employed following migration, especially if they moved for their own, or for both partners' jobs. Out of work women at t-1 who moved for their partner's job were actually more likely to be employed at t than women who did not move at all. Thus, the impacts of family migration are most obvious for employed women.

Third, we compared the results for the majority of women who have a lower occupational status than their partner with those who have a higher occupational score, with the expectation that moving will have an insignificant effect on women in the latter category. In fact, employed women who moved for the man's job were more likely to be out of work, regardless of whether their Cambridge score was higher or lower than their partners. This suggests that for employed women at least, the gender-neutrality of the human capital model does not seem appropriate and a gender roles perspective may well be more valid.

Finally, we tested whether women's labour market status was influenced by the metropolitan/non-metropolitan nature of the destination and the employment rate for women in the local area relative to the national average. We found that geography does influence employment rates, even controlling for a wide range of individual and household factors – women are more likely to be employed if they live in areas where women's employment rates tend to be higher.

Overall, these results provide convincing evidence of a family migration effect, but demonstrate that state dependence is a crucial factor. Women who were employed prior to moving suffer much more from family migration than women who were out of work prior to the move. This finding is important as around one third of the women in this sample were not employed at any particular wave.

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Table 1 Employment status at time t by employment status at t-1 and migration status (%)

Migration status	Employment status (t-1)	Employment status (t)	
		Employed	Not employed
Non-migrant / short distance migrant	Employed	91	9
	Not employed	18	82
Long distance migrant (30km+)	Employed	73	27
	Not employed	27	73

Table 2 Model parameters

Categories	Model 1 standard	Model 2 standard	Model 3 dynamic	Model 4 dynamic	Model 5 dynamic
Employment status (t-1)					
Employed (base)					
Not employed		1.754*	1.394*	1.445*	1.583*
Migration (30km+)					
Non-migrant / short distance migrant (base)					
Long distance mover	0.521*	0.895*	0.934*		
Reason for move					
Non-migrant (base)					
Moved woman's or both jobs				0.342	0.317
Moved man's job				1.107*	1.073*
Moved other reasons				0.380*	0.373*
Interaction Employment status (t-1) & Migration					
Not employed, long distance mover		-1.048*	-1.059*		
Interaction Employment status (t-1) & Reason for move					
Not employed, moved woman's or both jobs				-1.550*	-1.503*
Not employed, moved man's job				-1.332*	-1.227*
Not employed, moved other reasons				-0.496*	-0.439*
Age group					
16-24 (base)					
25-34	-0.345*	-0.193*	-0.181*	-0.184*	-0.139
35-44	-0.399*	-0.224*	-0.216*	-0.214*	-0.170*
45-54	-0.018	0.071	0.091	0.096	0.137
55-64	1.241*	0.816*	0.996*	0.995*	1.016*
Qualifications					
No higher qualifications (base)					
HND / HNC / Teaching qualification	-0.706*	-0.316*	-0.418*	-0.418*	-0.364*
Higher degree	-1.276*	-0.479*	-0.718*	-0.700*	-0.618*
Marital status					
Married (base)					
Cohabiting	0.031	-0.001	-0.016	-0.030	-0.035
Car ownership					
None (base)					
1	-0.497*	-0.278*	-0.317*	-0.319*	-0.333*
2+	-0.757*	-0.426*	-0.476*	-0.480*	-0.493*
Youngest child					
No children (base)					
0-4 years	1.131*	0.628*	0.745*	0.743*	0.780*
5-10 years	0.184*	0.114	0.134	0.138*	0.144
11-15 years	-0.325*	-0.161*	-0.183*	-0.177*	-0.158
Non-dependent children	-0.473*	-0.165*	-0.195*	-0.188*	-0.162*
Household size					
Two people (base)					
Three people	0.466*	0.230*	0.268*	0.265*	0.250*
Four or more people	0.525*	0.322*	0.364*	0.367*	0.326*
Region					
London (base)					
Other metropolitan	0.088	0.004	0.055	0.065	0.058
Rest of South East	0.206*	0.052	0.102	0.115	0.126
Rest of North	0.116	-0.001	0.063	0.073	0.063
Employment rate					
Women's rate < 75 percentile men's (base)					
Women's rate >= 75 percentile men's	-0.251*	-0.140*	-0.192*	-0.185*	-0.196*
Cambridge score (t-1)					
Women's Cambridge score < man's (base)					
Women's Cambridge score > man's					-0.052
Interaction Employment status (t-1) & Cambridge score					
Not employed, women's Cambridge score > man's					-1.148*
Constant	-0.279*	-1.014*	-0.923*	-0.958*	-0.993*

* P < 0.05

Figure 1 Migration distance by age and reason for move (bold horizontal line represents 30 kms)

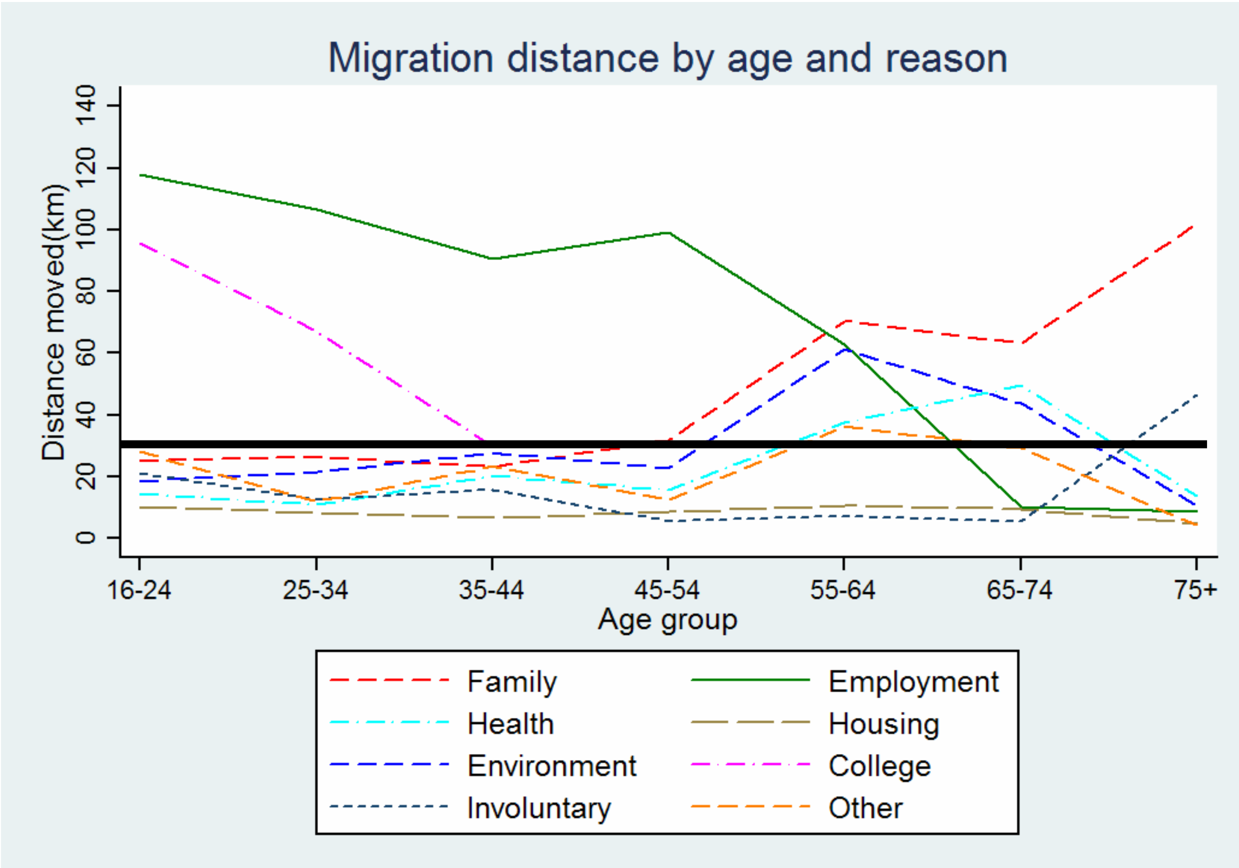

Figure 2 Probability of unemployment or economic inactivity by distance of move: random-effects probit model (Model 1, Table 2)

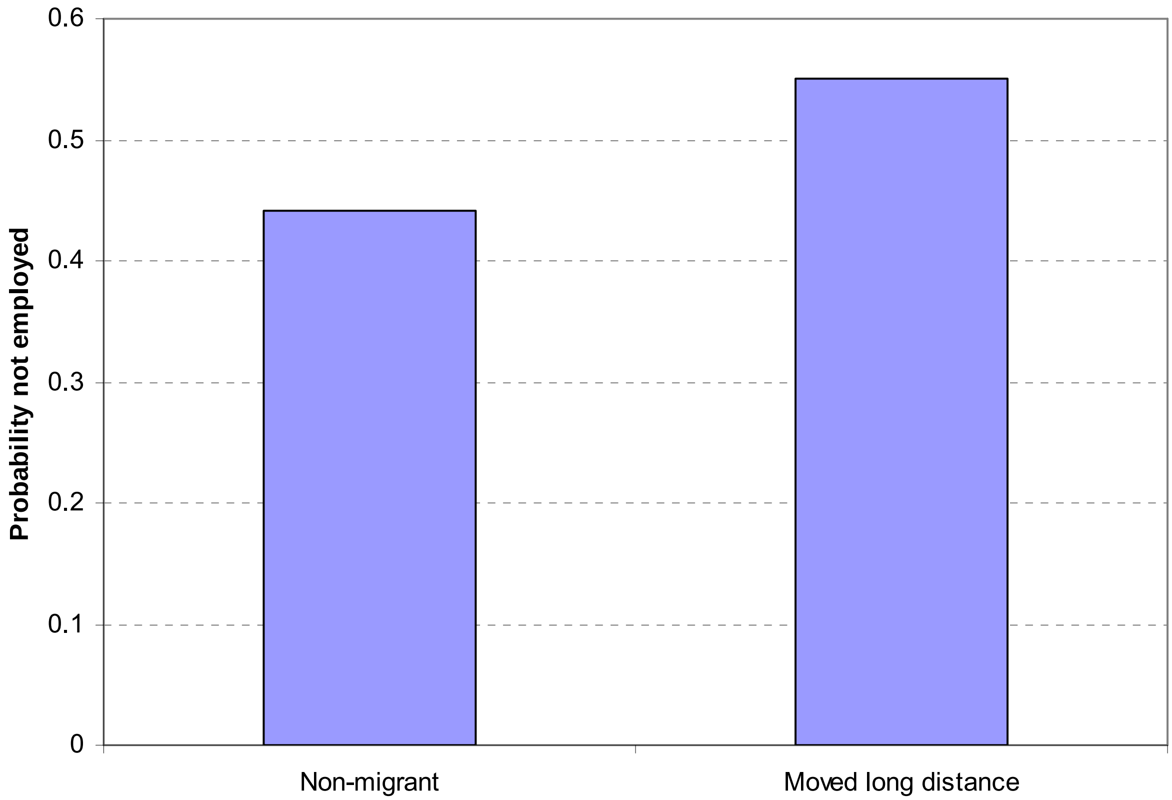

Figure 3 Probability of unemployment or economic inactivity by distance of move: random-effects probit model (Model 2, Table 2)

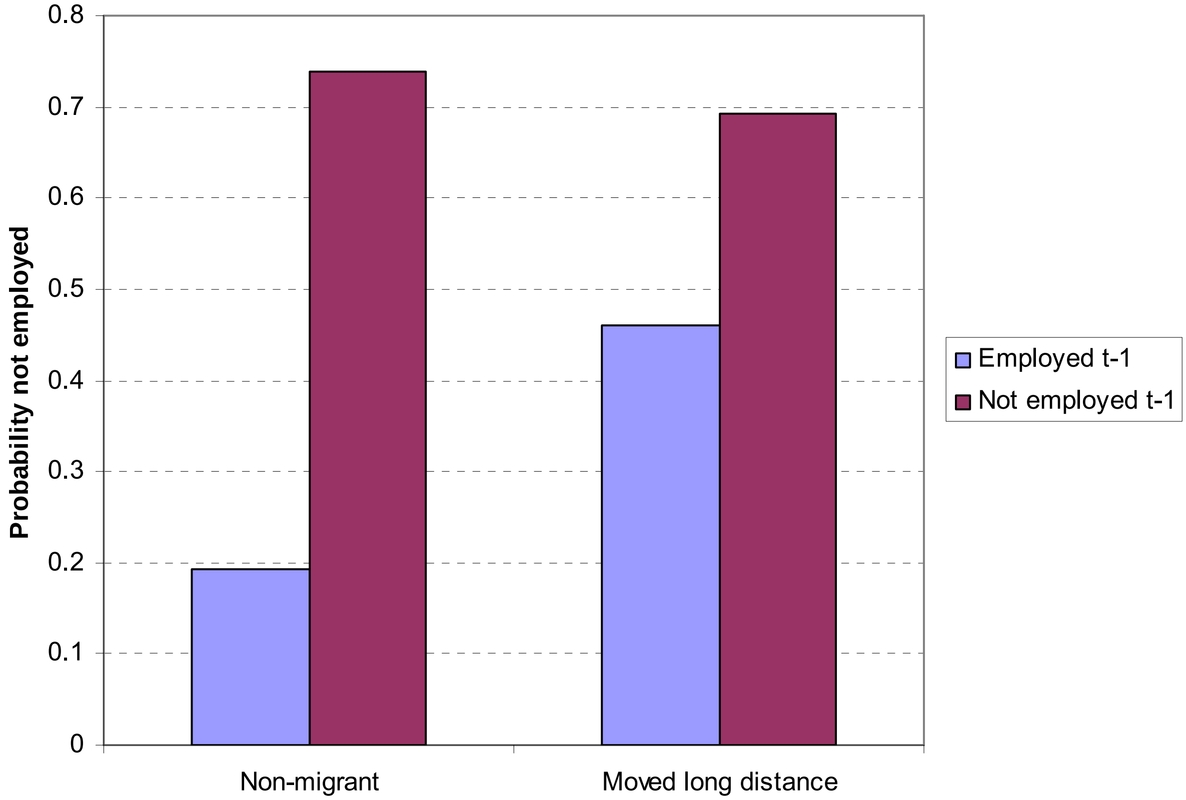

Figure 4 Probability of unemployment or economic inactivity by distance of move: random-effects dynamic probit model (Model 3, Table 2)

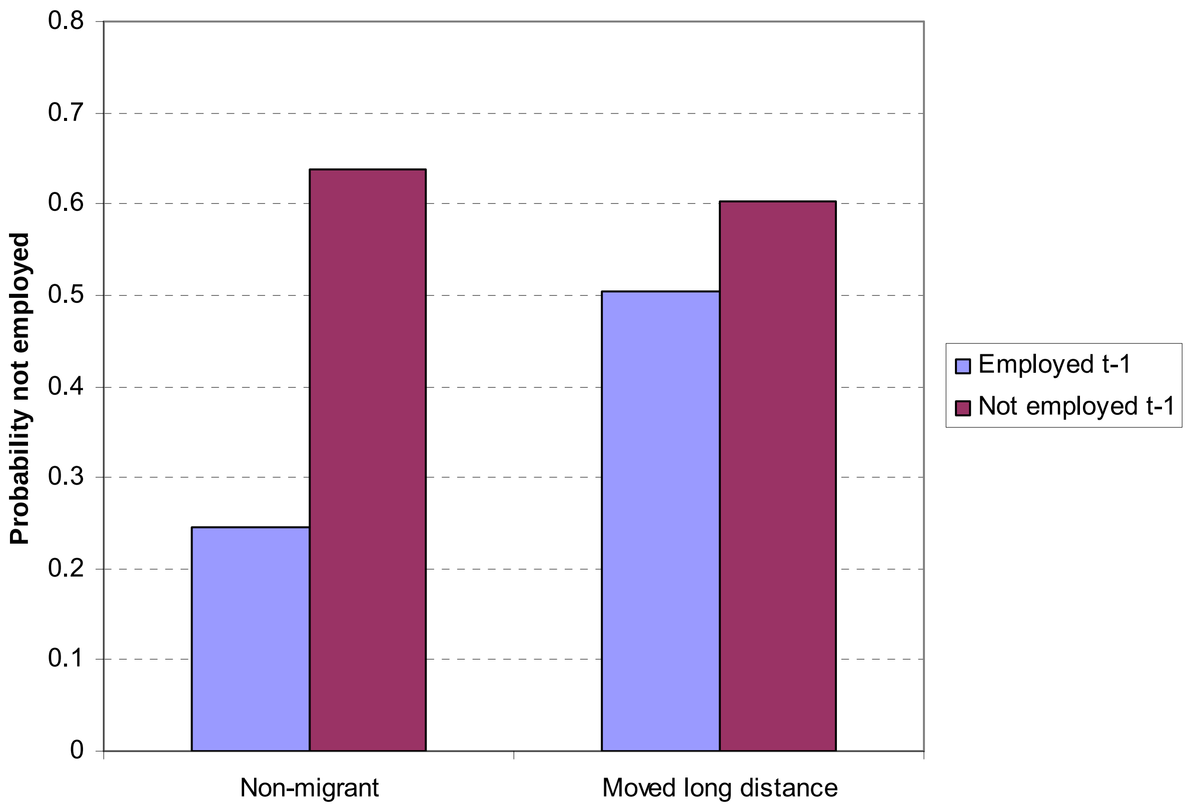

Figure 5 Probability of unemployment or economic inactivity by reason for move and lagged employment status (t-1) : random-effects dynamic probit model (Model 4, Table 2)

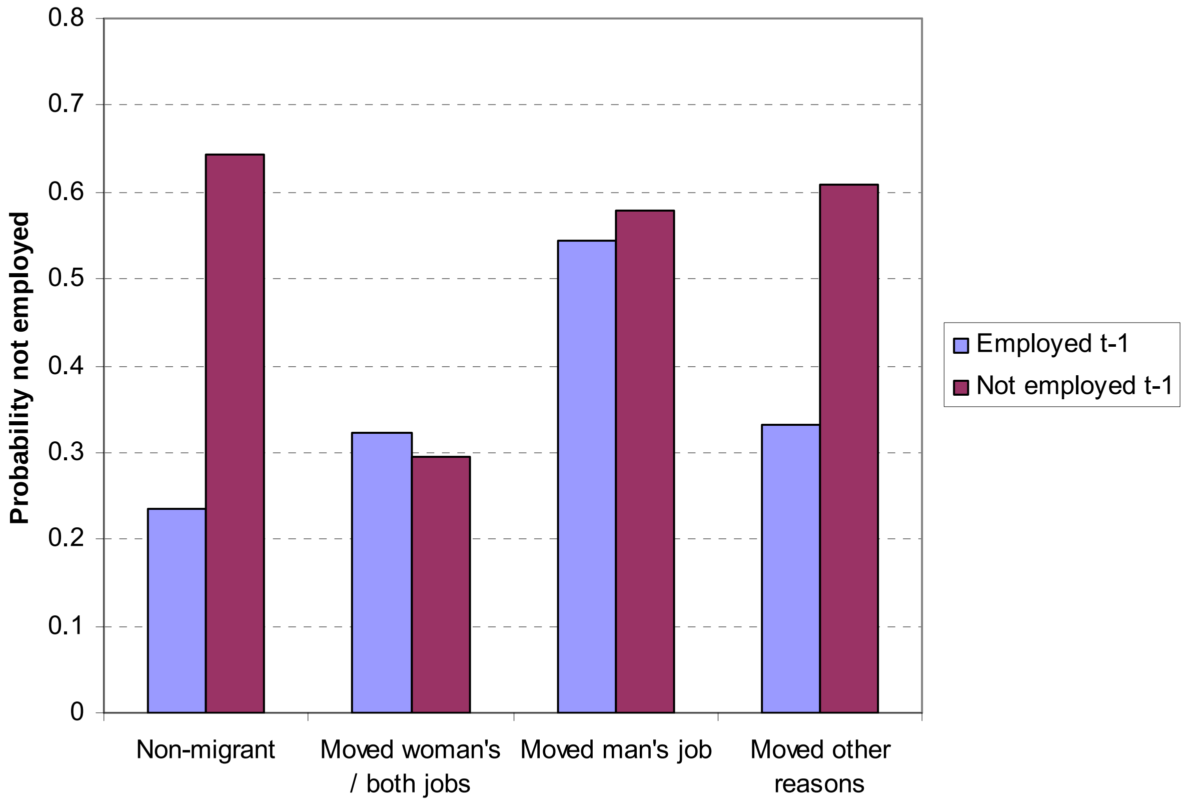

Figure 6 Probability of unemployment or economic inactivity by reason for move and lagged employment status (t-1), for women with lower lagged Cambridge scores (t-1) than partner: random-effects dynamic probit model (Model 5, Table 2)

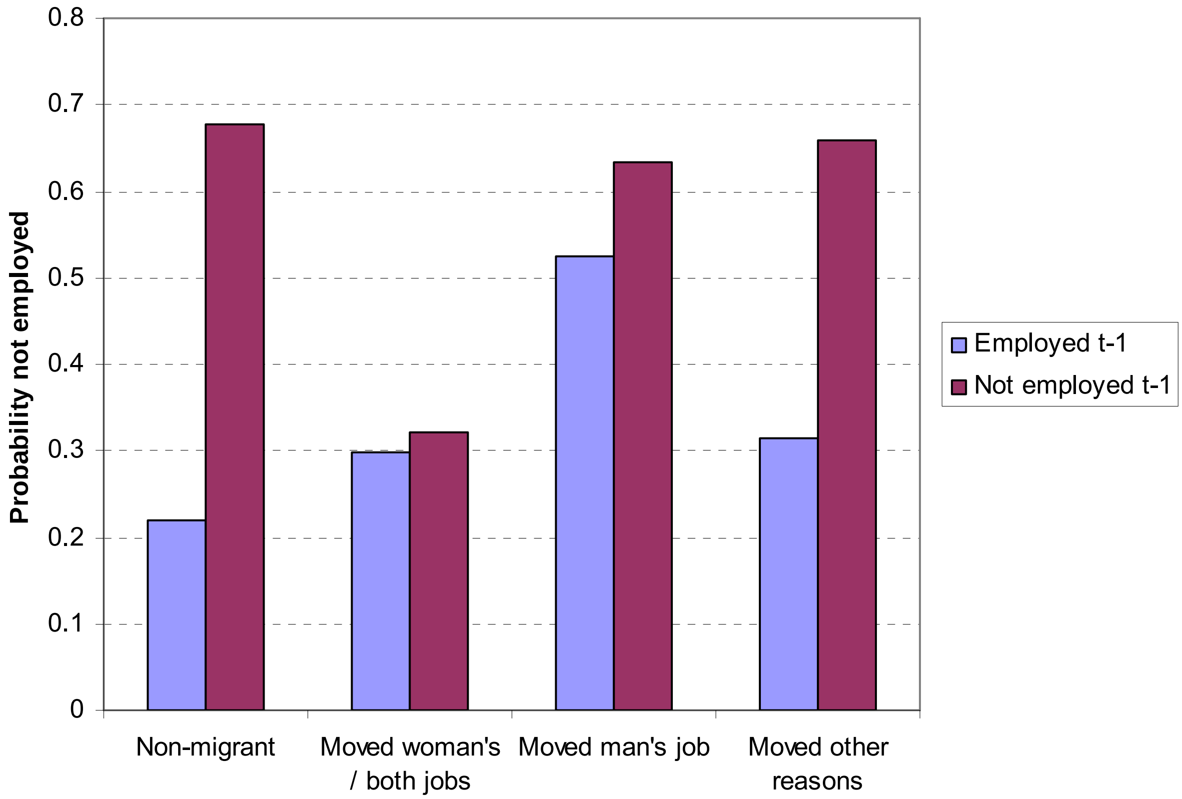

Figure 7 Probability of unemployment or economic inactivity by reason for move and lagged employment status (t-1), for women with higher lagged Cambridge scores (t-1) than partner: random-effects dynamic probit model (Model 5, Table 2)