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Employment Dynamics of Married Women in Europe

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Abstract

We use eight waves from the European Community Household Panel (1994-2001) to analyze the intertemporal labor supply behaviour of married women in six European countries (Netherlands, France, Spain, Italy, Germany and United Kindgom) using dynamic binary choice models with different initial condition solutions and non parametric distributions of unobserved heterogeneity. Results are used to relate cross-country differences in the employment rate to the estimated dynamic regimes. We find that cross-country differences in the employment rate and the persistence of employment transitions of married women are mostly due to composition effects related to education and unobserved characteristics rather than state-dependence effects or the dynamic effect of fertility.

Keywords: intertemporal labor supply, female employment, dynamic binary choice models, initial conditions.
JEL Codes: C23, C25, D91, J22.

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1 Introduction

In 2000, the European Union set a 60% female employment rate target for 2010.[1] Employment rates vary considerably across European countries and only a small fraction of that gap is accounted for by differences in unemployment rates.[2] Particularly for Mediterranean countries, the targets appear relatively ambitious.[3] Still there is room for some optimism. Pissarides et al. (2003) estimate that the momentum created by rising employment rates for recent cohorts of women can help southern European countries close the gap by one-third by 2010. Since this is probably insufficient to meet the targets, more should be known about what explains the differences across countries and what can potentially increase employment of women in Europe.

In this paper, we look at longitudinal micro-evidence on intertemporal employment decisions of married women across 6 European countries (France, Germany, Italy, Netherlands, Spain, and U.K.) for the period 1994-2001 using the European Community Household Panel. Similar to Hyslop (1999) for the United States, we pay particular attention to the dynamics of employment decisions, distinguishing true state dependence from unobserved heterogeneity, as well as the dynamic effect of births and permanent/transitory non-labor income on employment decisions.[4] This is motivated by the observation that in countries where there is high persistence in employment decisions of females, we also find the lowest rates of employment. To validate how these dynamics explain cross-country differences in the levels, we perform a decomposition exercise of the cross-country differences in employment rates as revealed by the empirical model. The econometric framework used is directly derived from a discrete-time version of the labor supply model with market frictions and heterogeneous home production proposed by Garibaldi

[1] The Lisbon targets also call for an increase of overall employment rates at 70% and employment rates of workers aged 45-64 to 50%.

[2] In 1999, The female unemployment rate for the age group 25-54 was 12.7% in Italy while it was practically the same in France (12.6%) (OECD Labor Market Statistics, 1999). However, both countries differ substantially in their female employment rates (near 70% in France across the 1994-2001 period and closer to 45-50% for Italy).

[3] From the OECD Employment Outlook of 2004 (p.296), women aged 15-64 in Italy had an employment rate of 46.3% in 2000 while this figure was 68.9% in the U.K.

[4] Other country specific studies using such type of models exist: Arulampalam et.al. (2000) for U.K., Croda and Kyriazidou (2004) and Buddelmayer and Voicu (2003) for Germany. Unfortunately, the data sets used are rarely comparable and so is the specification used.

and Wasmer (2003). A Key insight of the model is that the effect of rigidities or contagion effects on the equilibrium rate of employment is indeterminate and largely depends on the composition of the labor force.

We confirm that what seems to matter most to explain cross-country differences is the composition of the labor force, both in terms of educational attainment (and its associated employment "gradient"), and unobserved heterogeneity, while the effect of fertility is not able to explain these differences. As it turns out, the difference in the dynamics, contagion or state-dependence effects, do not explain the differences across countries when we permute dynamic employment regimes across countries. Furthermore, countries with low employment protection legislation tend to have higher state-dependence effects which provides additional evidence that state-dependence is not per se the explication for the cross-differences in employment rates.

2 Employment Patterns

2.1 Data

The analysis is based on individual data from the European Community Household Panel (ECHP, 1994-2001). The ECHP is a survey based on a standardized questionnaire that involves annual interviewing of a representative panel of households and individuals in each country, covering a wide range of topics including demographics, employment characteristics, education etc. In the first wave, a sample of some 60,500 nationally represented households - approximately 130,000 adults aged 16 years and over - were interviewed in the then 12 Member States. There are three characteristics that make the ECHP relevant for this study. That is, the simultaneous coverage of employment status, the standardized methodology and procedures yielding comparable information across countries and the longitudinal design in which information on the same set of households and persons is gathered. The sample is constructed as a balanced panel of all married females aged between 18 and 60 years old with their husband employed continuously during all the available waves. We condition on the husband's employment status in order to avoid having to specify jointly employment decisions. This is similar to Hyslop (1999) and enables some comparison with intertemporal labor supply of married women in the U.S. However, we must acknowledge that the conclusions drawn from this analysis may not be applicable to other types of couples in Europe. Nevertheless, in most couples, the husband is usually employed continuously throughout the period covered.[5]

We first look at general trends in employment rates stratified by education and number of child in the household. We then document the persistence in employment decisions and relate it to employment rates.

2.2 Evolution of Employment Rates

Table 1 shows employment rates stratified by education level for the six countries over three selected years. We observe an increase in female employment rates over time which is highest for the Netherlands and Spain (about 8 percentage points). On average, employment rates are higher in France, Germany, and the UK (between 65%-70%), and lower in the Netherlands, Italy (around 50%), and Spain (40%). Table 1 suggests also that in all countries those with higher education are more likely to be employed. Quite clearly, the difference in employment rates across education levels (the education "gradient") is higher in the Netherlands, Italy, and Spain, countries where participation of women is low. The cross-country differences in employment appear to be stronger for lower educated women than for higher educated as can be seen for Italy and Spain compared to the U.K. or France for example. This suggests that part of the explanation for differences in employment rates could be due not only to differences in the average education level but also due to the segmentation of the labor market along education level, with those in low levels having a harder time to find jobs in southern countries. Another explanation might be differences in preference for work related to the male breadwinner norm combined with low compensation from the market compared to household production.

[5] The number of couples for which the husband is not continuously employed and thus dropped from the sample are 65 (8.11% of total sample) for France, 152 (15.93% of total sample) for Germany, 48 (4.53% of total sample) for Italy, 38 (6.32% of total sample) for the Netherlands, 107 (13.44% of total sample) for Spain, and 93 (14.72% of total sample) for the UK. In terms of characteristics, these females are younger and less educated in Italy and Spain, more educated in Germany, older with less kids in the Netherlands, while they are observationally similar in France and the UK. Even when restricting to families where the husband is continuously employed, we still get for Italy, Germany and the U.K., employment rates for 2001 which are very similar to OECD figures. As for the Netherlands, Spain and France, some differences arise. These are the countries that suffered the biggest loss of observations when deleting couples where the husband was not continuously employed.

Over recent decades, the negative correlation that appeared between fertility and employment at the country level seem to have switched sign from negative to positive (Ahn and Mira, 2002). Countries in which high fertility is associated with high employment (mostly in the north) have implemented public childcare systems (or subsidized private ones, maternity and paternity leaves) and others introduced child benefit provisions during the 80s and 90s. Differences in childcare institutions and child benefit policies could potentially be associated with the variation in employment rates. But although the effect of both type of measures on fertility is probably non-negative, the effect of both on employment may work in opposite directions. While the existence of childcare may help women combine work and childbearing, child benefits may give the opposite incentive through an income effect. For example, in Germany, a non-working mother could receive 300 euros per month for a period of 2 years if she was taking care of a child under 2 years old in 2004. Since 1994 in France, that amount is nearly 500 euros per month for a period 3 years for the second child and onwards (European Commission DG 05 MISSOC 2004).

Table 2 shows employment rates for women in childbearing age (25-40) by whether they have one kid under the age of 16 in the household.[6] As expected females without kids have higher employment rates relative to those with one except for France where the difference is statistically insignificant. Focusing on 1994, these differences in employment rates are higher in absolute term in the case of the Netherlands, Germany and the UK. The biggest gradient in absolute value is found in the U.K. with women with one kid having 26.6 p.p. lower employment rates.

Institutional differences such as parental/maternity leave and child benefits might be able to explain part of the difference in the effect of one kid and employment differences across countries. In particular, the provision for parental leave and no provision for child benefits for the first kid in France relative to the lack of optional leave in the UK may partly explain their employment differences for the first kid in Table 2. Although options for parental leave are available also in Germany, child benefits for every kid including the first might create negative income effects which lower employment probabilities. From Table 2, employment rates for females in Germany

[6] Using families with more than one kid would render the comparison more difficult because these families would arguably be different in other ways which are not controlled for in this table.

and UK with one kid are similar. This difference in the provision of benefits for the first kid might also play a role in the employment differences between France and Germany. In the South, similarly to the U.K. institutions are much less developed, however the difference in employment rate of females without kids relative to those with one is smaller in absolute terms. One explanation may be the role of the extended family in the South as a substitute to formal institutions.

The extent to which fertility affects overall female employment rates depends also on the education level of mothers. As Table 1 suggests, employment rates differ across educational levels with high educated being more likely to be employed. Table 3, presents birth rates stratified by education level showing that births are more frequent among the more educated except for Italy. For Italy, low educated who tend to work less have higher birth rates.

2.3 Persistence

Table 4 presents statistics on labour market transitions in the sample. The average employment flows in the first two columns of Table 4 are calculated by pooling all transitions over all wave. The last four columns present the frequency of the number of transitions for females in each country. In Italy, about 75% of women do not experience any transition in their labour market status. In Spain and France too, only about 28% of women experience a transition. Finally, in the Netherlands, Germany and the UK the share of women in the sample who do not change status is much lower (64% for NL, and about 58% for GER and UK). Particularly in the U.K. and in Germany, there are respectively 13.2% and 15.5% of women making 2 transitions in the 8 years period of the panel.

Generally, there is therefore a lot of persistence in the employment behavior of these women. More so in countries where the employment rate is the lowest (Spain and Italy). Therefore, similar to the argument of Layard et al. (1991) for the U.S. - European difference in unemployment rates, persistence can potentially explain differences within Europe with respect to employment rates of women. In Table 4, we report a measure of persistence proposed by Shorrocks (1978).[7] Countries with low employment rates tend to

[7] For a state-space with S states, an index of mobility is given by $[S - tr(P)]/(S-1)$ where P is the estimated transition matrix.

exhibit more persistence (Italy, Spain) while countries with less persistence (Germany, U.K.) have higher employment rates. Manning et.al. (2004) show that in Mediterranean countries (Italy, Spain, France) there is also a gender gap in unemployment rates, which is associated with differences in transition rates between unemployment and employment. In particular, females are more likely to exit from employment to unemployment and less likely to enter from unemployment to employment compared to males. Therefore, it seems that countries which exhibit high persistence in employment patterns are those with low employment. This finding serves as a motivation to think about a model that investigates the sources of this persistence in order to explain the cross-country variation in employment rates.

3 Labor Supply with Imperfections

To illustrate how persistence can be associated with the equilibrium level of employment, we draw from models proposed by Garibaldi and Wasmer (2003) which is also similar to that in Hyslop (1999). Assume a married women making decisions at t regarding consumption c_t and market work y_t. We assume that market and non-market work are perfect substitutes which leads to the result that she will specialize in one of the activities.[8] Therefore, y_t is assumed to be dichotomous (0,1). These activities are rewarded at a rate w_t and a_t for market and non-market work respectively, which she can consume in addition to non-labor income m_t (mostly the husband's income). She has a well-behaved utility function $u(c_t, x_t)$ where x_t is a vector of taste shifters. In making decisions today, she discounts future utility at a rate $\rho < 1$ and has an infinite life horizon. She has beliefs about future states of the world (income, value of home production and taste shifters) summarized by an expectation operator E_t. Present discounted utility is given by

$$U_t = \sum_{s=0}^{\infty} \rho^s E_t(u(c_{t+s}, x_{t+s})). \tag{1}$$

[8] Interestingly, what seems to differ across countries is the extent of home production and market work, not leisure activities (Freeman and Schettkat, 2002). The assumption that market and non-market work are substitute leads to the clear result that each women will want to specialize in one of the activities. This is in line with our desire to focus on employment rates and behavior at the extensive margin of work where most action is anyway observed (Heckman, 1993).

We assume for exposition that she does not have the possibility of saving and therefore faces the following constraint in each period,

$$c_t = m_t + (w_t - \gamma(1 - y_{t-1}))y_t + (1 - y_{t-1})a_t, \ \forall t = t + 1, ... \qquad (2)$$

If performing non-market work at $t-1$, she faces search costs if she is to accept a job. These cost have a value γ which may include information search costs, time spent to search, etc. Although the model does not consider the effect of the duration of stay in employment on wages, one can also interpret γ as skill depreciation in non-employment. This decreases the reward to work (effectively the wage rate) and this is the mechanism which leads workers to face different reservation wages than non-workers. Assuming that income processes are stationary along with the infinite horizon and static budget constraint implies that the value function at the beginning of period t given the participation state variable y_{t-1} is $V(y_{t-1}) = \max(V^1(y_{t-1}), V^0(y_{t-1}))$, where the superscripts 1 and 0 denote period t employment and non-employment respectively.[9] For each option the value function is

$$V^1(y_{t-1}) = u(m_t + w_t - \gamma(1 - y_{t-1}), x_t) + \rho E_t V(1) \qquad (3)$$

and

$$V^0(y_{t-1}) = u(m_t + a_t, x_t) + \rho E_t V(0). \qquad (4)$$

Because of search costs, it is easy to see that two reservation wages should arise, essentially depending on the state-space variables. She will be indifferent between working and not working at a reservation wage $w_t^*(y_{t-1})$ such that $V^1(y_{t-1}|w_t^*(y_{t-1}))) = V^0(y_{t-1})$. Since $V^0(0) = V^0(1)$, it follows trivially that the two reservation wages will be related as

$$w_t^*(1) = w_t^*(0) - \gamma. \qquad (5)$$

The interpretation is trivial, the reservation wage for home-workers must compensate for the income lost incurred while searching for jobs (or the loss in skills). Garibaldi and Wasmer (2003) show the following inequality

$$w_t^*(1) < w_t^* < w_t^*(0) \qquad (6)$$

where w_t^* is the reservation wage when $\gamma = 0$, the neoclassical reservation wage. Note that the level of $w_t^*(1)$ depends on γ only if uncertainty is present.

[9] We suppress the dependence on x_t which is part of the state space at t

This is produced by "participation hoarding", even under the possibility that home-production produces more consumption than market work, uncertainty about future outcomes yields a positive value of holding on the job. Therefore, $w_t^*(1)$ is lower than w_t^*.

This characterization of reservation wages gives rise to the following decision rules,

$$y_t = I(w_t^*(q) + \gamma y_{t-1} > 0) \tag{7}$$

where $w_t^*(q) = w_t - w_t^*(0)$ and variables without subscript denote current values as well as anticipation about future values $q = (x, m, a)$.[10]

The dynamic equation for the employment rate at t is given by (where $G(w_t^*(q))$ is the distribution function for $w_t^*(q)$)

$$e_t = (1 - G(0))(1 - e_{t-1}) + (1 - G(\gamma))e_{t-1}. \tag{8}$$

The proportion of non-working women with $w_t^*(q) > 0$ is $1 - G(0)$ and therefore equilibrium employment $\overline{e}$ is given by

$$\overline{e} = \frac{1 - G(0)}{1 - G(0) + G(\gamma)}. \tag{9}$$

Interestingly, decreasing γ has two effects. It first leads to an increase in the flow at the entry margin because search costs are now lower. But, there is a flow out at the quit margin because some women who were staying on the market in order to avoid search costs (or loss of skills) if future states reveal lower productivity at home, do not face these costs anymore. Clearly, the effect of changing γ depends on the composition of the labor force in terms of w. The distribution of w which may be observed or unobserved, combined with state-dependence (γ) influences the equilibrium employment rate as well as the dynamics out of equilibrium. Clearly as we have seen, the distribution of transitions is different across countries in Table 3 and countries who are not doing well in terms of employment (Italy and Spain) do have the highest

[10] Implicitly, from the demand side of the market, we assume that the wage offer distribution depends only on q (education, age, etc). such that a reduced form "utility" from employment is summarized by $w_t^*(q)$. Hyslop (1999) assumes a different search process where search costs are born even if the wife does not accept a job. In this case, the parameter γ in front of y_{t-1} involves the premium that a non-worker will require because of the likelihood that he does not find a job after searching (provided she is risk averse). Hence, we assume search is always fruitful in this paper.

proportions of women who never work throughout the 8 years we survey them. Given that the composition of the population in terms of observables and unobservables can explain these differences in persistence, one way to infer the sources of persistence is to estimate an explicit form of (7) on the longitudinal data in each country. This is likely to be informative about what type of policies might be most likely to yield increases in employment of married women in countries that have lower employment rates.

4 Econometric Model

We define an indicator y_{it} that takes value one when respondent i reports being employed and zero otherwise in year t. We observe data for a married women for $t = 0, ..., T-1$. We assume the first-order Markov model for these binary employment decisions considered in Heckman (1981) which is consistent with equation (7),

$$\begin{aligned} y_{it} &= I(\mathbf{x}_{it}'\beta + \gamma y_{it-1} + \alpha_i + \varepsilon_{it} > 0) \\ t &= 1, ..., T-1 \end{aligned} \tag{10}$$

where $\mathbf{x}_{it}$ are observed characteristics that may or may not vary over time. One generally assumes that the unobservables can be decomposed in two parts, one time-invariant and the other time-variant. Of course, the interpretation that one will give to α_i will depend on its assumed relationship with the whole past, present and future of $\mathbf{x}_i = \mathbf{x}_{i0}, ..., \mathbf{x}_{iT-1}$. The same applies for time-variant unobservables that one typically interprets as shocks. A first assumption we make is that, conditional on unobserved heterogeneity α_i, $E(\varepsilon_{it}|\mathbf{x}_{is}, \alpha_i) = 0$ for all t, s. This assumption is defined as strict exogeneity. It implies that ε_{it} is independent of the past, present and future values of $\mathbf{x}_{is}$.

In economic terms, the parameter γ has a direct correspondence to the model outlined previously and corresponds to the wedge in the reservation wage due to search cost (or loss of skills) , whether or not a job was held in the last period. A high γ could imply that policies targeted at reducing this wedge could be effective (Heckman and Willis, 1977). The interpretation of γ in statistical terms has a long history in econometrics. It is usually referred to as true state-dependence (as opposed to spurious state-dependence). It captures a dependence of the participation decision, irrespective of unobserved heterogeneity, on past employment decisions. In contrast, heterogeneity creates persistence because of self-selection of those with high propensity in employment and those with low propensity in non-employment. This creates spurious dependence of aggregate probabilities of transitions on previous states. This has different implications for policy and targeting can be an effective mean of increasing employment. Longitudinal data with at least three repeated observations is necessary to distinguish the two (Heckman, 1981). Unless the $\mathbf{x}$ process varies over time, the identification of γ would rely on functional form. Fortunately, several characteristics included in $\mathbf{x}$ will vary over time (see section 4.4).

As for the assumption regarding the relationship between α_i and the $\mathbf{x}_i$ process, it is generally more restrictive to assume that $\mathbf{x}_i$ is strictly exogeneous ($E(\alpha_i|\mathbf{x}_i) = 0$). If one is prepared to neglect the non-linear nature of (10), then moment conditions of the form $E(\Delta\varepsilon_{it}|y_i^{t-2}) = 0$ in addition to standard orthogonality conditions for $\mathbf{x}_{it}$ can be used to estimate by GMM parameters of the model (e.g. Hyslop, 1999). In this fixed effect formulation, the estimator uses lags of employment decisions to instrument Δy_{t-1} (Arrellano and Bond, 1991) which identifies γ if the process is stationary. However, this estimator will have all undesirable properties that linear probability models have, especially when making simulations and predictions.

In non-linear models, the use of fixed effects is notoriously problematic. Neyman and Scott (1954) show that for a general class of non-linear models, in particular the probit model with fixed effects, the MLE estimator is inconsistent because it suffers from the incidental parameter problem. Chamberlain (1980) proposed a conditional logit approach in the static case and in the dynamic version without any observable heterogeneity that varies over time (Chamberlain, 1985). Honoré and Kyriazidou (2000) proposed a conditional logit approach for the general dynamic model but one that imposes strong requirements on the distribution of observable heterogeneity over time and that has a slower rate of convergence than the classical root-N rate. As Wooldridge (2004) notes, one problem with non-linear fixed effect models that goes often unnoticed is that they do not allow to compute average partial effects. In the cross-country comparisons we wish to make, it will be crucial to compute average partial effects because parameters are identified up to scale and there is no reason to believe that this scale is the same across countries. Hence, we consider the use of quasi-fixed effects (Chamberlain, 1984).

The key insight of that approach is that instead of assuming strict exogeneity at one extreme (random effects), or making no assumptions at the other (fixed effects), one can specify parametrically the relationship between observable heterogeneity and unobservable heterogeneity. In that case, one has to be careful about how to interpret average partial effects. For time-invariant characteristics, one can always interpret α_i as a residual unobserved heterogeneity term, net of association of observable and unobservable heterogeneity and the causal effect of observable heterogeneity on outcomes. This makes clear that unless strict exogeneity is assumed, one cannot identify the causal effect of time-invariant observable heterogeneity on outcomes. But for time-variant characteristics such as non-labor income which varies over time and includes for example the husband's earnings, one could expect that α_i is not independent of average earnings of the husband but that perhaps it is still possible to separate causal effects from heterogeneity effects by using transitory income over and above permanent income.

This is potentially applicable to non-labor income since Blundell and Smith (1986) clearly reject the exogeneity of non-labor income in a cross-section from U.K. data. If this rejection comes entirely from the correlation between α_i and non-labor income, then using that strategy deals with the endogeneity of non-labor income. Similarly to Hyslop (1999) we can write

$$\alpha_i = z_i'\delta + \eta_i \tag{11}$$

where η_i is conditionally mean independent from $z_i' = z_{i0}, ..., z_{iT-1}$, a subset of $\mathbf{x}_i$ for which we assume that strict exogeneity with respect to α_i fails.

We use a similar strategy for fertility. The difference is that we consider fertility to be a predetermined variable such that we condition on the initial number of children before making the conditional mean independence assumption between new births and employment decisions. This allows for feedback from employment to fertility but does not allow for simultaneous determination of fertility and employment decisions.[11] It turns out that this is a direct application of a strategy proposed by Wooldridge (2000) for predetermined variables.[12]

[11] We could use the strategy proposed by Angrist and Evans (1998) to isolate the effect of fertility on employment identifying it from employment probability differences between families having a third birth, after having two kids of the same sex or not. Unfortunately, the ECHP does not provide the gender of the kids.

[12] An alternative which we do not pursue here would be to model fertility as a separate decision and allow interactions with employment (see references in Del Bocca et al. 2004 for such attempts). The problem is then to find valid instruments or a quasi-experiment to get exogeneous variation in one of the outcomes.

In a quasi fixed effect framework, one needs to postulate a distribution for unobserved heterogeneity η. It can be assumed that η_i is normally distributed (as in Hyslop, 1999) or one could rely on a non-parametric approach as suggested by Heckman and Singer (1984). We favor the second approach, as combined with the quasi-fixed effect assumption it constitutes a good compromise to the fully non-parametric fixed effect specification.[13]

We assume that the distribution of η_i has K points of support $\eta_k, k = 1, ..., K$ with associated mass probability $p_k = P(\eta_i = \eta_k)$. Since the probit model is identified up to scale we fix $var(\varepsilon_{it}) = 1$ and in addition fix $\eta_1 = 0$. In the analysis we specify

$$p_k = \frac{\exp(\pi_k)}{1 + \sum_{k' \neq 1} \exp(\pi_{k'})}, k = 2, ..., K$$

with $\pi_1 = 0$ and $p_1 = 1 - \sum_{k' \neq 1} p_{k'}$. Using a mass point heterogeneity distribution is an attractive alternative to parametric distributions for two reasons. First, it allows comparisons of distributions of unobserved heterogeneity without relying of a unifying distributional assumption across countries. Second, the parametric alternative involves numerical methods (quadrature or simulations from a parametric distribution) that are not always precise when persistence is high (Hyslop, 1999).

4.1 The Initial Condition Problem

Since the whole history of y is not observed, the initial observation y_{i0} is potentially correlated with η_i such that integrating over the marginal distribution of η_i will yield inconsistent estimates. This is known as the initial condition problem (Heckman, 1981). To see why, denote $P(y_{it}|\mathbf{x}_i, y_{it-1}, \eta_i)$ to be the conditional probability implied by imposing (10) and (11). The probability of observing y_i, the whole sequence from $t = 1, ..., T-1$ conditional on the initial observation y_{i0} (included in $P(y_{i1}|\mathbf{x}_i, y_{i0})$) and η_i is

$$P(\mathbf{y}_i|\mathbf{x}_i, \eta_i, y_{i0}) = \prod_{t=1}^{T-1} P(y_{it}|\mathbf{x}_i, y_{it-1}, \eta_i) \tag{12}$$

[13] See Hansen and Lofstrom (2005) for a similar application of mass point heterogeneity distribution to dynamic binary choice models applied to social assistance decisions of immigrants.

where $\mathbf{y}_i = (y_{i1}, ..., y_{iT-1})'$.The marginal likelihood is given by

$$P(\mathbf{y}_i|\mathbf{x}_i) = \int P(\mathbf{y}_i|\mathbf{x}_i, \eta_i, y_{i0})dF(\eta_i, y_{i0}|\mathbf{x}_i). \tag{13}$$

Since y_{i0} is observed and η_i not, Heckman (1981) suggested writing (13) as

$$P(\mathbf{y}_i, y_{i0}|\mathbf{x}_i) = \int P(\mathbf{y}_i|\mathbf{x}_i, \eta_i, y_{i0})P(y_{i0}|\eta_i, \mathbf{x}_i)dF(\eta_i) \tag{14}$$

and specifying $P(y_{i0}|\eta_i, \mathbf{x}_i)$ as a reduced-form approximation to the recursive solution from $t = 0$ to the start of the process

$$y_{i0} = I(\mathbf{x}_{i0}'\beta_0 + \mathbf{z}_i'\delta_0 + \lambda\eta_i + \varepsilon_{i0} > 0). \tag{15}$$

An alternative approach suggested by Wooldridge (2004) is to replace (11) with

$$\alpha_i = \mathbf{z}_i'\delta + \psi y_{i0} + \eta_i \tag{16}$$

and assume a distribution for η_i. This equivalent to specifying a distribution for $\alpha_i|\mathbf{z}_i, y_{i0}$. Wooldridge then shows that the conditional MLE estimate from maximizing the log sum of (12), where (16) is substituted in (10) for α_i and η is integrated out, yields consistent estimates of parameters of interest (if (12) and (16) are correctly specified).

Taking into account the mass-point specification for $F(\eta_i)$, the integration takes the simple form

$$P(\mathbf{y}_i|\mathbf{x}_i, y_{i0}) = \sum_{k=1}^{K} p_k \prod_{t=1}^{T} P(y_{it}|\mathbf{x}_i, y_{it-1}, y_{i0,}\eta_i^k). \tag{17}$$

We do not allow for serial correlation in the errors ε as in Hyslop (1999) for two reasons. First, the solution proposed by Wooldridge does not seem to provide consistent estimates for fixed T if serial correlation is present. This arises since y_{i0} is correlated with future ε. A (approximate) test for serial correlation can be implemented from estimating the model from moment restrictions with linear probability models (Arrellano and Bond, 1991). In results not reported here, we do not find evidence of serial correlation in the errors once state-dependence and unobserved heterogeneity is allowed for. Estimation of the parameters in the quasi-fixed effect models is done by maximum likelihood using the BFGS algorithm.

4.2 Marginal Effects

Parameters are not directly comparable across countries because they are normalized under a different scale in binary choice models. Therefore we rely on average partial effects (commonly referred as to as marginal effects) to compare effects across countries.

Wooldridge (2004) suggests computing the average partial effects of changes in x using

$$m_W(\mathbf{x}, y_{t-1}) = \frac{1}{N}\sum_{i=1}^{N} \nabla_{\mathbf{x}} P(y_{it} = 1|\mathbf{x}_i, y_{it-1}, y_0). \tag{18}$$

and similarly for y_{it-1} where $\nabla_{\mathbf{x}}$ denotes a finite difference for discrete outcomes while the continuous derivative is used for continuous xs. Note that information on initial condition y_0 is used to compute for each individual the marginal effect. In this sense, the initial conditions are used to weight the effects, since they are informative about α_i. Asymptotically, if the cross-section size is large, then $m_W(\mathbf{x}, y_{t-1})$ does not depend on the distribution of y_0 as noted by Wooldridge. Therefore, if we turn to estimates obtained using the Heckman initial condition solution, computing

$$m_H(\mathbf{x}, y_{t-1}) = \frac{1}{N}\sum_{i=1}^{N} \nabla_{\mathbf{x}} P(y_{it} = 1|\mathbf{x}_i, y_{it-1}) \tag{19}$$

where parameters are obtained from maximizing over (14) should be equal to $m_W(\mathbf{x}, y_{t-1})$ if both initial condition methods yield similar average partial effects. It is these quantities that we will use to compare both methods and also assess the importance of persistence to explain cross-country differences in employment rates of married women.

4.3 Decomposition of Cross-Country Differences

The model is (without the i subscript)

$$\begin{aligned} y_t &= I(x_t'\beta + \gamma y_{t-1} + \alpha + \varepsilon_t > 0), t = 1, ..., T \\ y_0 &= I(x_0'\beta_0 + \lambda\alpha + \varepsilon_0 > 0) \end{aligned}$$

with ε_0 i.i.d. (0,1) and α with discrete distribution $F = (\pi, p)$ where π are points of support and p their associated probabilities.

Given estimates of the parameters for each country, we can try to decompose the cross-country difference in employment rates. We can do that for differences that are explained by different composition in terms of observed and unobserved heterogeneity. But, these differences could also be due to differences in the way this heterogeneity is associated with (or causing) employment outcomes. The differences in state-dependence can also explain cross-country differences. Therefore, we can perform a decomposition exercise similar to what is usually referred to as Oaxaca decomposition (Oaxaca, 1973).

Denote the relevant elements of the regime for country a ,

$$R_a = (\beta_a, \gamma_a, \mathbf{x}_a, F_a).$$

and similarly for country b where $\mathbf{x}_m = \{\mathbf{x}_{im}\}$ is the sample of observed heterogeneity in country m.

Denote the simulated employment rates using a draw j from the distribution of ε and F , $e_j(R_a) = e_j(\beta_a, \gamma_a, \mathbf{x}_a, F_a)$ and $e_j(R_b) = e_j(\beta_b, \gamma_b, \mathbf{x}_b, F_b)$ using the estimates of β_a, γ_a and F_a along with the respective initial conditions in each country. The raw difference in employment rates is

$$\Delta_j = e_j(R_a) - e_j(R_b).$$

We can decompose this difference as

$$\Delta_j = \Delta_{\beta j} + \Delta_{\gamma j} + \Delta_{xj} + \Delta_{Fj}.$$

where differences in effects are given by

$$\begin{aligned}
\Delta_{\beta j} &= e_j(\beta_a, \gamma_a, \mathbf{x}_a, F_a) - e_j(\beta_b, \gamma_a, \mathbf{x}_a, F_a) \\
\Delta_{\gamma j} &= e_j(\beta_b, \gamma_a, \mathbf{x}_a, F_a) - e_j(\beta_b, \gamma_b, \mathbf{x}_a, F_a)
\end{aligned}$$

and differences in the composition of each sample are given by

$$\begin{aligned}
\Delta_{xj} &= e_j(\beta_b, \gamma_b, \mathbf{x}_a, F_a) - e_j(\beta_b, \gamma_b, \mathbf{x}_b, F_a) \\
\Delta_{Fj} &= e_j(\beta_b, \gamma_b, \mathbf{x}_b, F_a) - e_j(\beta_b, \gamma_b, \mathbf{x}_b, F_b).
\end{aligned}$$

Now we compute each employment rates for J draws ($j = 1, .., J$) from the error distributions and average them. Then we compute the differences and report each component. This is done for all years recognizing the dynamic structure of the model (simulating both initial conditions and the main equation using lagged simulated outcomes). We take the outcome in 2001 to report decompositions since the dynamics in the different regimes take some time to reach an equilibrium given their new characteristics. We compute the share of the raw difference explained by each component in order to get an idea of what explains the gap across countries. Unfortunately, because the model is non-linear, this does not enable to look at the separate contribution of each component of observed heterogeneity to explain the employment gap. We can however look directly at the average partial effects to get an idea of their respective contribution.

4.4 Specification

We include broad age indicators with the reference group being respondents aged 40 to 50. Along with these indicators, we include a cohort or "vintage" variable (expressed as birthyear - 1930 divided by 10).

We account for many of the circumstances that could explain variation over time in behavior. We include one health indicator for whether the respondent reports being in very good health and fertility indicators. These consist of indicators for the net increase in the number children under 16 years old.[14]

We also allow for the transitory component of the husband's income to have a separate effect from the association and causal effect of permanent income which cannot be separated apart. As for time-invariant heterogeneity, it consists of education level (low, medium and high) where low education is expressed as the reference group.

We control for the initial condition correlation with taste for work by including the number of kids in the first wave. The initial condition specification when using the Heckman initial condition solution is the same where the lagged employment is not included (these last results are available upon request).

[14]Unfortunately, the ECHP does not provide the age of the children which is argued in Hyslop (1999) to be a good indicator of the parental care needs.

5 Discussion of Results

5.1 Marginal Effects

Table 5 presents marginal effects from the dynamic probit estimation using the Heckman initial conditions. We focus our presentation of results on the relationships with human capital, fertility and income. Typical age patterns of employment rates are found for all countries and cohort effects in Spain and the Netherlands explain a big part of the positive momentum in these countries' employment rates. It implies nearly a 10 p.p. increase in employment rates for those born in 1965 compared to those born in 1955.

Having completed secondary or higher education (compared to the baseline elementary education) is associated with higher employment probabilities in all countries. In particular, higher education is associated with increased employment, 34 percentage points in Italy, 36 p.p. in Spain, 17 p.p. in the Netherlands and Germany, 9 p.p in France, and 5 pp. in the UK. This mirrors the descriptive evidence in Table 1 and perhaps shows that a significant share of the difference in employment rates across countries comes from differences in employment rates by education levels within each countries.

We now turn to the potential effect of fertility and childbearing on employment outcomes. In our specification we control for the number of kids at the first year in the sample. This is consistent with the view that fertility is a predetermined variable and therefore we rule out simultaneous contemporaneous effects going from employment to fertility in order to refer to an effect of births on employment. We do however consider the correlation between unobserved tastes for work and fertility and this correlation appears to show up for all countries except for France and the UK.

The direct effect of fertility is captured by controlling for the first and second or more births during the period that the sample is observed.[15] Both effects are negative and significant for the Netherlands, Spain and the UK, while for Italy and Germany there is no direct effect of fertility. For France, the second or more birth has a negative and significant effect on employment. This lines up well with the presence of the *Allocation Parentale d'Éducation* (APE) at the time the panel was in place, which gives a child benefit of nearly 500 euros per month for 3 years, for births of second rank or more. We find effects of the order of 11 p.p. for second or more births, while the first birth

[15] For those with no kids at first year, the dummy for first birth will coincide with the birth of the first kid, while the dummy for the second or more births with additional kids.

does not attract a statistically significant parameter estimate. Since this is a pattern unobserved in the other five countries, it gives us some confidence that it may be picking up some of that incentive structure. Laroque and Salanié (2003) find that the elimination of the APE would have lead to the creation of 45,000 jobs in France in 1997 while Choné et al. (2003) estimate the effect to be 4 p.p. on employment probabilities. Additional evidence comes from Piketty (1998) who estimates that 17% of the women eligible to the APE in 1994, when it was extended to second births, did quit the labor market after its introduction.

In Germany, although child benefits are generous there is no effect of births on employment. However, German women use in far greater proportion maternity, parental and sick leaves than their French counterparts (Evans, 2001). The figures from the Second European Survey on Working Conditions and the Employment Options of the Future Survey reported in Evans (2001) show that 92,3% of employed women with a child under 15 in Germany report having extra statutory arrangements maternal, parental or sick leaves, while that proportion is 57.9% in France. Less dramatic but still indicative of the degree of family-friendly policies in Germany, flexible time arrangements are also available for 33.2% of them while this fraction is 26% in France. Therefore, these childcare program may serve to counterbalance the negative incentive effect on employment that generous child benefit may have. Merz (2004) provides evidence that the generous parental leave legislation in Germany, expanded in 1986 and in following years, may explain why, through the period 1970-2002, employment rates of married women with children rose steadily compared to employment of other groups.

In Italy, the absence of an effect of fertility might be due to the fact that births are more predominant within the low educated group, as Table 3 shows, who are less likely to be employed. Moreover, the support from the extended family might play its own role.

Compared to a standard random effect probit our specification allows for correlation between fertility and husband's income with unobserved heterogeneity as describe in equation (11). The results from Table 5 show that exogeneity of permanent husband's income fails in all countries. Of course, in order to make that claim, we must assume that permanent income itself does not have a causal effect on employment. Then the rejection of the null gives support to the endogeneity hypothesis. In general there appears to be a strong association between the husband's income and the wife's unobserved propensity to work. The disincentive effect of the tax system is probably better captured by the effect of the transitory component of income. It turns out that the transitory component of the husband's income does not have a significant effect except for Germany and the UK, where a 10 thousand euros increase in husband's income reduces the employment rate by 28 p.p and 21 p.p., respectively. This represents an elasticity of -0.68 at the median husband income in the sample.

This is consistent with the institutional setting in Germany in which the tax system benefits one-earner families. Marginal tax rates for the second earner are relatively higher than in any other European country (OECD, 2003). This rate was 57% in 2000 compared to 30-40% in other countries. In particular, due to joint taxation an increase of husband's income, which increases household income, reduces the splitting advantage that the joint taxation system provides (Gustafsson, 1992). For the UK, the working families tax credit (WFTC) which is payable to families with net income lower than 90 pounds a week (as of 1999, and 75 pounds before) might create incentives to withdraw from employment for females as eligibility can be based only on husband's income. (Blundell et.al., 2000).

The specification using Wooldridge's initial conditions is shown in Table 6. Estimation of main regressors is less precise compared to the Heckman specification. However, it is quite striking how the average partial effects of many characteristics are similar given that the Wooldridge solution is far more easier to implement than the Heckman method. However there are some differences. The exogeneity test of permanent husband's income differs between the two specifications. With the Wooldridge solution we do not reject exogeneity of husband's income for France, Italy, the Netherlands, and Spain as we do with Heckman initial conditions, but we do reject for Germany and the UK. Moreover, the countries for which exogeneity of fertility is rejected, as depicted by the number of kids at first year, differs between the two methods. Of course, this is probably largely due to the fact that including initial employment in the quasi-fixed effect picks up a lot of the correlation between taste for work and any other outcome related to the husband. In what follows, we will refer mainly to the results using the solution proposed by Heckman since effects are more precisely estimated.

5.2 True vs Spurious State-Dependence

The key insight from the theoretical model we used is that the amount of state-dependence and the distribution of heterogeneity in the population will affect the equilibrium employment rate but also the dynamics of the aggregate employment rates following changes in the composition of the labor force or aggregate shocks to the labor market. Table 5 presents the marginal effect for state dependence (working at t-1) and Table 7 the distribution of unobserved heterogeneity using the Heckman solution for the initial conditions. Note that some of that heterogeneity is already picked up by time-invariant differences in education level, husband's permanent income and initial number of kids at baseline controlling for cohort differences in employment rates. These were already found to have important associations with employment rates of respondents.

Layard et.al. (1991) present an index of labor market rigidities in Europe for which the countries we look at split in two groups. First, there are the northern countries such as the U.K., the Netherlands and to some extent Germany are seen as more liberal labor market. The second group is then composed of France, Italy and Spain. In the model we presented, one source of rigidity is the presence of search cost or skill depreciation/job-specific human capital. When we look at the evidence from Table 5, we see a reverse ordering of countries based on the level of the state-dependence effect if it were completely attributed to search costs. The Netherlands has the highest structural difference in employment probabilities for women who worked in the last year compared to those that did not. The effect is 48.9% while France has a state-dependence effect of 40.0%. For the rest of the countries state-dependence effect varies between 27.0% and 31.0%.[16] Therefore, it has to be that state-dependence is capturing other "state" effects such as skill depreciation. We are not aware of international comparisons that would give support to such hypothesis based on the ranking of countries in terms of state-dependence effects.

As we mentioned earlier, the equilibrium level of employment under different state-dependence effects will depend on the composition of the labor force in terms of observable heterogeneity and also unobserved heterogeneity at the two margins. If participation hoarding dominates the entry effect, then a country where state-dependence is higher may have a higher aggregate level of employment such as in the U.S. and the U.K.

In terms of unobserved heterogeneity, Table 7 shows the estimates of the three mass points and their probabilities for each country. The points are

[16] Compared to the results from the Wooldridge solution in Table 6, we observe that the marginal effect of previous participation is slightly higher under the Heckman solution.

not directly interpretable in absolute terms but one can say something about the dispersion and the asymmetry in the distribution. For example, in the Netherlands, the distribution appears to be skewed towards low propensity women, with nearly 40% of the sample having a very low unobserved propensity to work compared to only 20% who have a high propensity. The same is true in Spain and to some extent in Italy. All of these are countries with low employment. In the case of Netherlands, coupled with high state-dependence this means that potentially many more women pile up at the entry margin than at the exit margin. In contrast, France has a high share of women with very high propensity to work and high state-dependence, which could mean that there is more participation hoarding than women piled up at the entry margin. This is even more dramatic for Germany and the U.K. who have disproportionately larger shares of women having high propensity to work than low propensity to work. These differences remain however unexplained. One can only speculate, in the same way as for the education gradient, that these unobserved differences are potentially due to cultural and social norms, unobserved dimensions of human capital, or segmentation of the labor market in ways not captured by observable heterogeneity we used, that lead specific groups in southern countries to stay out of the labor market.

5.3 Cross-Country Differences in Employment

To shed more light to the importance of state dependence, unobserved and observed heterogeneity and composition effects in explaining the differences in employment rates across countries we perform the decomposition described in section 4.3. We perform this decomposition for groups of countries as shown in Table 8. The term Δ_β measures for country A the differences of returns of characteristics due to living in country A, while Δ_γ measures the difference of state dependence. The term Δ_x represents the "explained" component of differences in employment rates between groups of countries. Finally, Δ_F measures differences due to unobserved characteristics.

The comparison between the U.K. and Spain illustrates well one of our main point. Comparing United Kingdom with Spain the employment rate difference predicted through our model is 33 percentage points. Table 8 shows that most of this difference is due to differences in the return to characteristics (more than 75%) and to a lesser extent due to differences in the distribution of these characteristics between the two countries. In terms of characteristics, the biggest difference observed is in the share of high educated (43% for the UK compared to 22% for Spain). But the damage to employment is reinforced by the large difference in the effect of high education on employment probabilities (36.5 p.p higher employment for high educated in Spain vs. 5.5 p.p. in the UK - Table 5). State-dependence cannot be the explanation for differences between the two countries since it is higher in the U.K., a country with higher employment rates. So the difference in persistence between the two countries is mostly due to observable heterogeneity (education, income and fertility) and their returns in the labor market.

Similarly for the comparison between United Kingdom and Italy, Table 8 shows that we predict an employment rate difference of 21 p.p which is also mostly due to differences in the return of characteristics, driven again plausibly by the effect and composition of education. Differences in time-invariant unobserved characteristics seem also to play a more important role when comparing UK with Italy. In Italy, the distribution is skewed towards low employment probabilities which means more women plausibly at the entry margin. When the U.K. contagion effect is used in Italy, this widens the gap. So the correlation between persistence and employment levels is again entirely due to observable and unobservable heterogeneity and their returns.

Therefore, although state dependence is significant in each country, this decomposition shows that differences across countries in state dependence is not able to explain the observed differences in employment rates. Differences in the return to observed characteristics and composition effects especially related to human capital and differences in unobserved characteristics are mostly behind the observed cross-country differences in employment. This can be most convincingly seen for the differences between the U.K, France. and the Netherlands.

6 Conclusion

Our aim was to study employment dynamics of married women in Europe using comparable panel data for six countries that differ considerably in their labor market institutions and family care provisions. We showed how a theoretical model such as the one proposed by Garibaldi and Wasmer (2003) could be used to derived dynamic employment equations in the same way as in Hyslop (1999). We then compared two estimation methods to deal with the initial condition problem and used non-parametric techniques to deal with unobserved heterogeneity. Finally, we proposed a decomposition method to assess the importance of several factors in explaining differences in employment rates.

The main question we asked was whether persistence could explain low employment rates in Southern Europe and, if it did, what was the source of that persistence. Our findings illustrate that it is not the differences in contagion or state-dependence that explain differences in employment rates across countries. In fact, the contagion effects found cannot be linked to rigidities per se: countries with high state-dependence have in general more flexible labor markets. Rather, it is the distribution of observed and unobserved heterogeneity, along with their respective returns that appear to capture most of those differences. The segmentation of the labor market across education levels seems to be the most important factor with low educated in the South facing lower employment rates, while the effect of fertility is not able to explain these differences.

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Table 1: Employment rates stratified by education level

% employed	1994	1997	2001
Netherlands			
total	44.9	43.6	52.9
low ed	34.0	30.9	41.2
high ed	63.5	58.3	68.8
France			
total	69.1	67.7	70.7
low ed	63.8	63.3	65.8
high ed	75.9	77.3	82.1
Italy			
total	47.4	48.3	49.1
low ed	30.2	30.0	30.6
high ed	81.6	84.4	86.7
Spain			
total	31.9	32.5	39.6
low ed	18.6	18.3	24.8
high ed	62.3	64.9	74.7
Germany			
total	62.4	60.9	65.2
low ed	52.6	52.1	56.8
high ed	78.4	77.8	81.1
United K.			
total	67.1	69.7	69.5
low ed	61.2	67.4	64.7
high ed	70.1	75.0	74.1

NOTES: Married women aged 18-60 with husbands continuously employed 1994-2001 Middle level of education ommitted.

Table 2: Employment rates stratified by number of kids

% employed	1994	1997	2001
Netherlands			
no kids	86.1	78.1	85.7
one kid	64.1	58.9	48.0
France			
no kid	77.7	70.0	88.0
one kid	80.7	72.0	76.0
Italy			
no kid	66.7	71.7	75.0
one kid	51.8	51.6	58.7
Spain			
no kid	57.5	63.4	63.3
one kid	40.0	38.7	55.5
Germany			
no kid	87.5	85.1	88.2
one kid	68.3	68.5	74.2
United K.			
no kid	94.3	92.7	90.0
one kid	67.7	64.3	67.3

NOTES: Married Women aged 25-40 with husband continuously employed 1994-2001. presence of a kid under 16.

Table 3: Birth rates stratified by education

	Low Education	Medium Education	High Education
Netherlands	.0427	.0559	.0744
France	.0426	.0596	.0680
Italy	.0550	.0567	.0524
Spain	.0478	.0476	.0565
Germany	.0557	.0523	.0623
United K.	.0547	.0582	.0714

NOTES: Married Women aged 25-40 with husband continuously employed 1994-2001. birth rate from wave 2 onwards

Table 4: Persistence and Mobility

	average emp. flow		mobility	number of transitions			
	in (%)	out	Index	0	1	2	3+
Netherlands	8.7	7.7	0.164	64.6	22.9	6.4	6.0
France	11.6	4.9	0.165	71.5	13.9	10.6	3.9
Italy	6.3	6.1	0.123	76.9	9.7	8.5	5.0
Spain	6.2	8.9	0.151	72.7	13.4	7.6	6.4
Germany	16.0	9.1	0.258	58.5	16.3	15.5	9.7
United K.	19.0	7.9	0.269	57.7	19.9	13.2	9.3

NOTES: Married women aged 18-60 with husbands continuously employed 1994-2001. Average employment flow calculated pooling all transitions over all waves. The mobility index lies between 0 and 1, values closer to zero meaning higher persistence.

Table 5: Dynamic Probit Model - Heckman Initial Conditions

Marginal effects	Country		
	Netherlands	France	Italy
age <30	-.145*	-.049	-.023
age 30-40	-.093*	-.034*	-.036*
age 50+	-.005	-.026	-.046*
med education	.053*	.053*	.205*
high education	.172*	.088*	.346*
good health	.046*	.0008	-.001
husband income (10,000 euros)	-.018	.0006	-.081
1st birth	-.058*	-.024	-.038
2nd birth	-.078*	-.112*	-.007
Permanent husband income	-.454*	-.037*	-.200*
# kids at first year	-.027*	-.012	-.028*
Cohort Year/10	.075*	.007	.005
Employed at $t-1$	.489*	.401*	.288*
Number of married women	563	736	1011
	Spain	Germany	U.K.
age <30	-.132*	-.039	-.088*
age 30-40	-.098*	-.024	-.059*
age 50+	.022	-.037	.022
med education	.078*	.054*	.002
high education	.365*	.178*	.055*
good health	-.006	.009	.046*
husband income (10,000 euros)	-.038	-.284*	-.211*
1st birth	-.087*	.052	-.076*
2nd birth	-.059*	.008	-.100*
Permanent husband income	-.277*	-.791*	-.448*
# kids at first year	-.017*	-.045*	.004
Cohort Year/10	.087*	-.054*	.030
Employed at $t-1$	.276*	.294*	.311*
Number of married women	689	802	539

NOTES: Marginal effects calculated using equation (19). * denotes a p-value less than or equal to 5 percent.

Table 6: Dynamic Probit Model - Wooldridge Initial Conditions

Marginal effects	Country		
	Netherlands	France	Italy
age <30	-.146*	-.032	-.011
age 30-40	-.093*	-.037*	-.023*
age 50+	-.013	-.017	-.033*
med education	.032	.024	.052*
high education	.071*	.032*	.143*
good health	.046*	.0002	.004
husband income (10,000 euros)	-.020	.002	-.067
1st birth	-.061*	-.024	-.034
2nd birth	-.078*	-.107*	-.012
Permanent husband income	-.108	-.020	-.059
# kids at first year	.004	.007	-.013*
Cohort Year/10	.048*	.016	.0003
Employed at $t-1$	.454*	.438*	.221*
Number of married women	563	736	1011
	Spain	Germany	U.K.
age <30	-.116*	-.036	-.088*
age 30-40	-.081*	-.027	-.055*
age 50+	.035	-.035	.028
med education	.044*	.050*	-.011
high education	.132*	.117*	.020
good health	-.014	.008	.044*
husband income (10,000 euros)	-.018	-.301*	-.209*
1st birth	-.079*	.046	-.068*
2nd birth	-.058*	.006	-.103*
Permanent husband income	-.120	-.552*	-.282*
# kids at first year	-.003	-.007	.048*
Cohort Year/10	.089*	-.060*	.027
Employed at $t-1$	.257*	.279*	.301*
Number of married women	689	802	539

NOTES: Marginal effects calculated using equation (18). * denotes a p-value less than or equal to 5 percent.

Table 7: Unobserved Heterogeneity Distribution

	Country		
	Netherlands	France	Italy
Point 1	-2.56*	-.055	-.71
Point 2	2.67*	-1.35*	-1.80*
Point 3	1.14*	1.24*	1.83*
Prob 1	0.46*	0.32*	0.32*
Prob 2	0.17*	0.21*	0.40*
Prob 3	0.37*	0.47*	0.28*
	Spain	Germany	UK
Point 1	-4.82*	.048	-2.02*
Point 2	2.25*	3.20*	2.94*
Point 3	4.15*	1.44*	1.42*
Prob 1	0.38*	0.33*	0.13*
Prob 2	0.45*	0.19*	0.44*
Prob 3	0.17*	0.48*	0.43*

NOTES: For estimation, the first point was normalized to zero while we report points adding the estimate of the constant term. These estimates correspond to the model estimated in Table 4. * denotes a p-value less than or equal to 5 percent.

Table 8: Decomposition of Employment Rate Differences for year 2001

	Country Comparisons			
	UK/Spain	UK/Italy	UK/NL	France/NL
Δ : Total Difference	.332	.212	.200	.191
Δ_γ: Difference SD	-.028	-.034	-.073	-.016
Δ_β: Difference Param.	.225	-.180	-.035	.369
Δ_x: Difference X	.091	.001	.003	-.058
Δ_F: Difference F	.044	.425	.305	-.102

NOTES: Sample used is married women aged 18-60 with husbands continuously employed 1994-2001. Decomposition performed as described in section 4.3. 500 draws are used to average employment rates. Estimates from Table 4 and 6 used.