Women's Employment, Children and Transition: An Empirical Analysis on Poland[§]

Elena Bardasi [#] Chiara Monfardini[*]

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Abstract

The effect of transition from centrally planned to market economies on female employment is unclear a-priori. Many studies have pointed out that the emergence of labour markets created obstacles to but also new opportunities for women's employment. A frequently mentioned potential explanation of the lower female participation during the transition period is represented by the reduction of childcare facilities, which created a major constraint on the participation of women with dependent children. However, we must not forget the effect of forces of opposite sign, first of all the household necessity of having two earners during the turbulent transition period. The aim of this paper is to give an empirical assessment on how the transition to a market economy affected the relationship between motherhood and labour force outcomes in Poland. We estimate random effects probit models on two PACO panel datasets covering a four year period before the reform (1987-1990) and a three year period afterwards (1994-1996). Our findings indicate that during transition small children were much less of a deterrent to the employment probability of their mother than it was before transition.

JEL classification: J13, J22, P23, C23

Keywords: female employment, fertility, transitional economies, Poland, panel data, PACO database.

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[#] World Bank, Washington DC, USA; e-mail: ebardasi@worldbank.org

[*] Department of Economics, Università di Bologna, Italy, e-mail: chiaram@economia.unibo.it

1. Introduction

In this paper we investigate the role of motherhood in shaping employment outcomes of Polish women before and during transition. The almost universal female labour market participation in the pre-transition period and the rapid deterioration of child care facilities during transition, which made more difficult for women to combine paid work and family responsibilities, have been greatly emphasised. However, we will show that employment was more responsive to motherhood before than during transition. In fact, in the pre-transition period a set of constraints existed that represented obstacles to mothers' employment; similarly, during transition, together with forces of negative sign, there were also pressures pushing mothers into the labour market.

There has been very little research on mothers' labour supply choices in transition economies. Fong and Lokshin (2000) examined how women's labour force activity depends on household child care choices in Romania. They showed that after transition the labour market activity of women was strongly affected by the presence of children and was highly responsive to the cost of childcare. Saget (1999) estimated a model of labour supply decisions of married women in Hungary to measure the impact of the number and age of children on women' labour supply choices (in addition to her wage and the non-labour income of her household). Using post-transition data, she found that the presence of young children decreases the probability of a woman participating in the labour market, while the presence of teen-agers has a positive influence on the decision to participate.

These studies suggest that, during transition, women in Central and Eastern European countries tended to display a labour market behaviour more and more similar to the one of women in Western economies. Bonin and Euwals (2001), in their study of participation behaviour of East German women after German unification, investigated precisely this question. They wondered how the high unemployment rate and the change in public policies that followed transition impacted the high female labour market participation of the pre-transition period and whether East German women will eventually adapt their behaviour to West German patterns. They found that participation is negatively affected by the number of pre-school children, and that this disincentive effect increased over time after unification.

However, they do not use data from the period *before* the unification to estimate how different women's labour market behaviour was before the transition.

To our knowledge no specific study exists for Poland and in particular no comparison of pre- and post-transition female labour market behaviour has ever been carried out. While the position of women in the Polish labour market was weak before the transition, it possibly weakened even further during the transition. Before 1989, despite the very high female labour market participation, women were concentrated in low skilled and low paid occupations. In a very traditional society such as the Polish one, women were considered to be the main childcare providers; the rhetoric about female's emancipation through participation in the labour market was mainly functional to the need of the production system to resort to women's labour in a period of rapid industrialization (Nowakowska and Swedrowska, 2000). At the same time, women's wage was needed to support the family. Almost all employed women were employed full-time; the availability of childcare facilities and family-support measures provided by the state allowed women to combine paid work and family responsibilities. However, women being regarded as the main responsible of child-care and as secondary earners inside the family, they had to sustain the double burden of carers and workers (Ruminska-Zimny, 2002).

During transition, activity rates dropped considerably, for both women and men. Polish women experienced a dramatic drop in employment – about 900,000 jobs were lost by women from 1989 to 1995. Unemployment, non-existent in the communist period, was 13 per cent in 1997 for women of all ages, and 10 per cent for men (UNICEF, 1999). Women with children were disproportionately more likely to be out of work. According to a study by UNICEF based on Polish Labour Force Surveys "Poland marriage status – a reliable proxy for the presence of children – was not a factor for unemployed men in finding work, but was a serious handicap for unemployed women" (UNICEF, 1999, p. 29). However, the same analysis found that, at the same time, the presence of children was not increasing the chances of employed women of losing their job with respect to men, indicating that many women needed and were actually able to combine both paid job and family commitments.

Some legislative advancements during the 1990s were aimed at increasing the protection of employees and women's position in particular. The amendments of the Labour Code in 1996 introduced for the first time the principle of gender equality in the labour market. Although the anti-discrimination, equal treatment measures have been indicated as insufficient and ineffective in many cases (Nowakowska and Swedrowska, 2000), they clearly represent an attempt at protecting the deteriorating position of women in the labour market.

To sum up, while several factors represented obstacles and disadvantages for mothers' employment during transition, other factors created pressures for women to stay in paid work. The drop in real wages had a substantial role in pushing women to participate in the labour market. At the same time, the emergence of more flexible forms of employment – such as part-time employment, fixed-term contracts, and self-employment (in Poland particularly relevant in agriculture, because of the importance of family farms) – offered more opportunities to women, even if at the price of an increased competition for jobs. By contrast, the drop in the number of jobs, the increased job insecurity, the lower availability and higher cost of childcare facilities were all going in the direction of decreasing female labour market participation. Women with children had also to face additional difficulties, including increasing discrimination in the emerging private labour market and the persisting social expectations that women should have taken care of the children and assume the burden of family responsibilities. Despite the concentration of women in the sectors that were going to remain public (education, health, social services, public administration) could have helped mothers to overcome the difficulties of combining paid work with maternity responsibilities, women with children may have been the big losers of the transition period. There is actually evidence that, despite the government effort to promote parental leave (in large part to compensate for the cut in workplace nurseries provision), women's position remained very weak. While men's use of parental leave is still negligible, working women tend not to use the whole of their parental leave to avoid adverse consequences on earnings and career. More generally, employers are likely to regard the potential absence of women following childbirth as a disincentive to promote them or even hire them. There is growing evidence of dismissals from work associated to parental leave.

In this paper we ask the following questions: is motherhood in Poland a significant constraint to women's employment after transition? Was it the case before the transition? In our analysis we use panel data for both periods in order to disentangle the direct effect of children on female employment from women's unobserved taste for children and employment. We will argue, based on our results, that in the pre-transition period characterized by virtually no constrains to join the labour market (no unemployment, availability of childcare facilities), ‘preferences’ – including also cultural attitudes and social expectations about the role of mothers in a traditional society such as the Polish one – played a fundamental role in shaping mothers’ employment patterns. On the other hand, during transition, the need of a second earner in the household acted as a pressure stronger than the cultural disincentives to mothers’ employment and the increasing difficulties represented by the deterioration of childcare facilities.

The paper is structured as follows. Section 2 describe the dataset we use for our investigation. In section 3 we provide some descriptive statistics on the participation and fertility behaviour of women in Poland across transition. Section 4 presents the econometric approach we adopted to estimate the female employment probability. The empirical evidence from the estimated probit models, and the interpretation of the main findings are reported in Section 5, while Section 6 concludes.

2. Data

We address the above issues using the panel data for Poland available in PACO – the original source is the Polish Household Panel survey. PACO is an international comparative database provided by CEPS/INSTEAD (based in Luxembourg) integrating micro-data from various national household panels over a large number of years. For Poland, the available data consists of two separate panels: the first includes four waves from 1987 to 1990, the second three waves from 1994 to 1996. Although the individuals observed are not the same in the pre- and post transition period, we can still exploit all the advantages of using panel data to estimate a binary choice model of mothers’ employment and compare the results in the two periods. The main advantages of using the Polish datasets through PACO are the consistency of variable definition over time (the original variables have been standardised and codified according to standardised international classifications), and (potentially) the comparability with other countries included in the database (Hungary, among transitional economies, and some West-European economies like Germany, UK). Moreover, in the case of Poland PACO offers two panels, one covering the pre-reform period and the other the transition period – overall a relatively long time span that allows for comparisons before and after the reform.

The data contains extensive information on labour force and work history (employment status, professional status, working time, working sector), and includes a set of household- and individual-level income variables, education and family background variables, household composition and demographic characteristics of each individual, and territorial division.

3. Descriptive evidence on female labour supply and fertility in Poland

In this section we use the Polish Household Panel in PACO to provide some stylised facts about the employment and fertility behaviour of Polish women. Figure 1 shows the proportion of women aged 16 to 44 with a partner and household responsibilities (i.e. who are either head of the household or the spouse of the head) who were in employment in the period before the transition (1987-1990) and after the transition (1994-96), by age of the youngest child. We chose this age interval in order to account for the existence of a non zero percentage of very young mothers (lower bound); the upper threshold of 44, on the other hand, is typically chosen as the oldest fertile age in studies on fertility behaviour.[1]

For women in this age interval, the employment rate dropped very little, from 69 per cent in the first period to 66 per cent in the second period on average. However, the employment rate of mothers varied substantially depending on the age of the youngest child; moreover, the patterns were quite different in the two periods. Before the transition, only 42 per cent of women with a child aged 0 to 2 were in employment, compared with 65 per cent of mothers of a youngest child aged 3 to 5 and 79 per cent of mothers of a youngest child 6 to 11. During transition, the employment rate of mothers was also increasing with the age of the youngest child, but the differences were much less pronounced. So, a higher percentage of mothers of a child 0 to 2 were in employment than before the transition – 50 per cent of them. However, only 59 per cent of mothers of a youngest child 3 to 5 and 71 of mother of a youngest child 6 to 11 were employed during transition. The largest difference in the employment rate before and during transition existed for the group of women without

[1] Moreover, we focused on women who were either the head or the spouse because only for this group of women it was possible to unambiguously identify their own child from other children living in the household. See footnote 4.

dependent children, 86 vs. 73 per cent. These patterns suggest that, first, there were substantial differences in the probability of women to be employed between the two periods that were unrelated to motherhood (see the differences in employment rates for women without dependent children between the 1987-1990 and the 1994-1996 period). Second, while before transition the presence and the age of children appeared to influence quite heavily the employment behaviour of women, during transition the female employment rate were much less responsive to the age of the youngest child.

Figure 1 – Percentage of women aged 16-44 with partner and household responsibilities (either head or spouse) who are employed, by age of youngest child

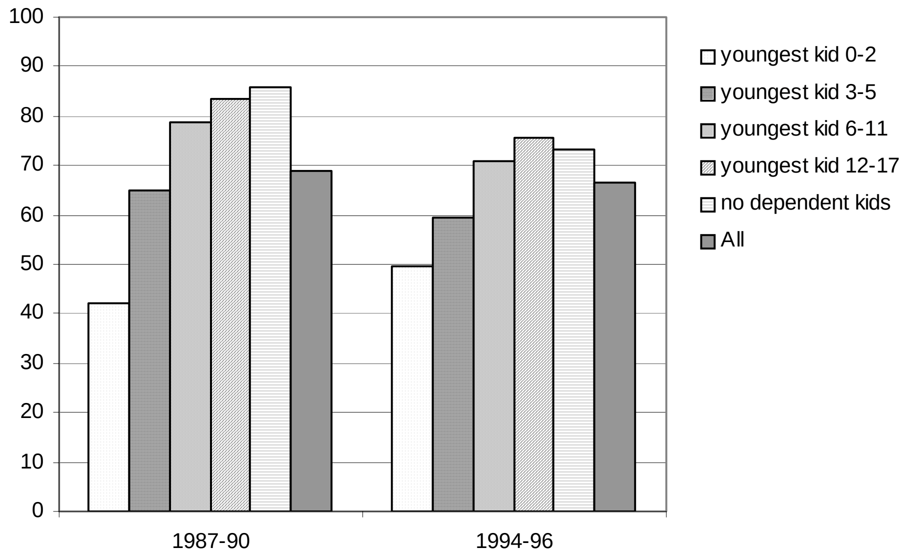

Figure 1 shows the employment rates of women by age of the youngest child but does not take into account the total fertility and the changes in fertility patterns between the two periods. It could be that the large differences in employment rates of mothers of children of different age before the transition be driven by a total higher fertility – so that the age of the youngest child influenced quite heavily the employment rate of the mother because of the presence of more siblings in the first than in the second period. Table 1 shows that, if there is some validity in this explanation, it is however very limited.

Table 1 – Average number of dependent children of women with partner and household responsibilities, aged 16-44, by mother's employment status

	Working	Not working	All	Working	Not working	All
Average number of kids 0-2	0.14	0.43	0.23	0.13	0.27	0.18
Average number of kids 3-5	0.32	0.59	0.41	0.24	0.37	0.28
Average number of kids 6-11	0.82	0.87	0.83	0.71	0.77	0.73
Average number of kids 12-17	0.72	0.47	0.64	0.75	0.64	0.71
Average number of kids 0-17	2.00	2.36	2.11	1.83	2.05	1.90

It is true that the average number of children generally decreased between the two periods, but not dramatically so. For example, the average number of children 0 to 2 for each woman aged 16 to 44 was quite similar in the two periods, but Figure 1 showed that the employment rates of mothers of these very young children was 8 percentage points higher in the second period. Moreover, Table 1 confirms that there was less difference in the average number of children between working and not working women in the second period with respect to the 1987-1990 period, especially in the average number of 0-2 and 3-5 children.

The evidence presented so far seems to suggest that the presence and the age of children was a crucial element in explaining women's employment before transition rather than during transition. This is in contrast with the general notion of universal female labour market participation before transition and the often mentioned increased difficulties of mothers during transition due to the deterioration of childcare facilities.

Another element that should not be overlooked in this analysis is the role of educational qualification. Because women with different education levels are likely to display different attitudes toward employment and were also differently affected by labour market changes during transition, in Table 2 we show the proportions in employment disaggregated by educational qualifications as well. As it was shown in Figure 1, both before and after transition mothers had lower employment rates than non-mothers or mothers of children 18 or older. Moreover, the younger the youngest child was, the lower the employment rate of the mother. This pattern was common across education levels, although some important differences ought to be stressed.

Table 2 - Proportion of women with household responsibilities aged 16-44 who are employed, by age of the youngest child and educational qualifications

Age of the youngest child	Second level, 1st stage and lower	Second level, 2nd stage high school	Second level, 2nd stage professional	Third level
1987-1990				
Children 0-2	0.473	0.401	0.391	0.564
Children 3-5	0.657	0.690	0.590	0.828
Children 6-11	0.724	0.887	0.741	0.893
No dependent children	0.855	0.986	0.946	0.984
1994-1996				
Children 0-2	0.479	0.522	0.415	0.712
Children 3-5	0.533	0.632	0.521	0.828
Children 6-11	0.620	0.745	0.664	0.895
No dependent children	0.744	0.850	0.672	0.912

Women with household responsibilities are defined as those who are either the head of household or the spouse of the head. Dependent children are children aged 17 and younger. Each child category indicates that the woman is mother of children of the corresponding age category and has no younger children.

First, highly educated mothers (third level education) had substantially higher employment rates than mothers of children of the same age but with lower educational qualifications. Moreover, highly educated women tended to re-enter employment when the child was still very young – the employment rates of highly educated mothers with a child aged 3-5 was over 80 per cent both before and after transition. Second, the transition does not seem to have affected the employment rate of mothers as much as it affected the employment rate of non-mothers. While women with non-dependent or no children had sensibly lower employment rates in the period 1994-1996 than they had in the period 1987-1990, the employment rate of mothers was lower in the post transition period but not that dramatically so and in the case of mothers of children aged 0 to 2 there was actually no difference. An age effect is likely to drive this result – transition negatively affected older women more than it did younger women. Third, women were differently affected by the drop in employment rates during transition depending on their educational qualifications. Highly educated women did not experience any drop in employment rates, especially those who were mothers (a drop in the employment rate of non-mothers with third level education occurred, but at 7 percentage points this is not as dramatic as for other categories of women). Women with lower educational qualifications, by contrast, experienced large drops in employment rates. It seems therefore that the education level was a key factor in explaining employment outcomes following transition, more than the fact of having or not having children.

Finally, Table 3 shows the employment rates of married women depending on the participation in employment of their partner. Before transition – when the employment rate of men and women was very high – having a partner out of employment was associated with a very high employment participation of the wife, while women whose partner was working were much less likely to work. This could be due to a genuine effect of the husband's employment status on his wife's employment status (compatible with the added worker effect, as well as with the effect produced by an increase in the household non-earned income on the woman's labour supply choice). Interestingly, after transition, the difference in the employment rates of the two groups of women is reversed, and having a partner out of employment makes the women less likely to work. During this period, the employment rate of men also drops, so it is much more likely to observe couples where both partners do not work. These patterns suggest that in both periods a lot of heterogeneity is likely to exist among women (women whose partner is in employment are very different from women whose partner is out of work). In order to (partially) control for heterogeneity we estimate our model separately for the sample of women whose partner is in employment.

Table 3 - Proportion of married women 16-44 who are employed, by employment status of their partner

	Partner not employed	Partner employed	% of women with partner in employment
1987-1990	0.730	0.686	94.1
1994-1996	0.620	0.673	84.3

4. Econometric modelling

The study of female labour supply and its correlation with the presence of young children has a long tradition in the applied econometric literature. Two exhaustive surveys are given by Browining (1992) and Nakamura and Nakamura (1992). These papers review the main issues concerning the estimation of the effects of the presence of children on women's labour supply and examine some selected studies on this topic. The vast majority of the empirical studies finds a negative correlation between fertility and labour supply of women. However, the interpretation of this result must be confronted with a number of issues. The existence of heterogeneity among women, differing in their fixed unobserved preferences for work and children, introduces selection effects that can bias 'genuine' effects. For example, it could be that women with a higher 'taste' for children are those with lower unobservable abilities and skills on the labour market. In that case, it would be this selection effect to drive the observed association between a higher number of children and a lower labour market participation, rather than a 'true' causal effect due to discrimination against mothers or to a change in tastes following maternity. Availability of panel data in the more recent empirical investigations of female labour supply allows for a proper consideration of unobserved heterogeneity. However, when this woman specific time invariant component is correlated with the desired number of children, the problem of endogeneity of the fertility variables arises. A possible approach to this problem is to follow Chamberlain (1984) and parametrically specify the individual unobserved heterogeneity allowing for correlation with the observed fertility variables. We follow this suggestion in our empirical analysis, given the major problems characterizing alternative solutions, like the Instrumental Variable estimation.[2]

A further important feature to consider to understand women's work behaviour has to do with the inter-temporal nature of labour supply. There is evidence of a high degree of serial persistence in individual participation decisions, possibly due to state dependence (with current participation depending on past participation) and persistent heterogeneity (when the unobserved individual component determines current participation irrespective of past participation). As a consequence, econometric models of female labour supply should explicitly include dynamic factors. Recent studies by Hyslop (1999) and Buddelmeyer et al (2003) adopt a dynamic framework to evaluate the interaction between fertility and labour supply decisions. Unfortunately, this approach is not viable in our case, given the short

[2] Angrist and Evans (1998) suggested the use of the sibling-sex composition to construct IV estimation, given the observed correlation between having two children of the same sex and further childbearing. However, this strategy is difficult to implement because in Central and Eastern Europe, as in most developed countries, the number of women with at least two children is typically very small.

period characterizing the two panel available to us.[3] In the following sub-section we present the empirical strategy we adopt for our analysis, which is based on static models, but fully exploits the panel dimension of our dataset.

4.1 Empirical specification

We included in our sample women aged between 16 and 44, who were either the spouse of the head of the household or the head of the household themselves. In fact, we were able to ascertain the parental relationship with children living in the household only for the head and his/her spouse.[4] Finally, we excluded from the analysis single women, whose higher heterogeneity with respect to partnered women is well documented in the literature on female labour supply (see Browining, 1992). Following a similar argument, we limit our analysis to women with employed partner in each of the sample year. These selection criteria left us with a sample of 3452 person-waves in the first panel (period 1987-1990), and 3933 person-waves in the second one (period 1994-1996).

We defined participation as the status of being employed, similarly to what is done in the most recent empirical studies of female labour force participation (see Hyslop, 1999), and modelled the binary indicator (employed/not employed) by means of probit models. As a starting point and benchmark for our analysis, we considered two classical models: the pooled cross-section probit and the random effects panel probit model. The cross section probit model ignores the repeated observations on the same individuals, and can be sketched in terms of a continuous latent variable as:

$$y_{it}^{*} = x_{it}^{'}\beta + u_{it} \tag{1}$$

i=1,…N; t=1,…T, with the binary dependent variable indicating participation given by the indicator function: $y_{it} = 1(y_{it}^{*})$. The vector x_{it} contains observed human capital, demographic, family structure and other variables which may affect the participation decision, while u_{it}

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[3] Voicu and Buddelmeyer (2003) adopt an alternative approach and model participation with panel data without specifying a random effect individual component, but allowing for a general correlation pattern of error terms over time.

[4] The PACO dataset only records the relationship of household members with respect to the head of the household. We assume that the children of the household head are automatically children of the spouse of the household head or have the same effect on her labour market behaviour as her own children.

captures the effects of unobserved factors and it is assumed to be *I.I.D.* standard normal. Accordingly, the likelihood function of the cross section probit model is built on the assumption that:

$$\Pr(y_{it} = 1 \mid x_{it}) = \Phi(x_{it}^{'}\beta)$$

with $\Phi(.)$ denoting the standard normal cumulative density function.

The random effect probit model splits the error term into a time invariant individual random effect and an idiosyncratic random error:

$$y_{it}^{*} = x_{it}^{'}\beta + \alpha_i + \varepsilon_{it} \qquad (2)$$

with $\varepsilon_{it} \sim I.I.D.N(0,1)$, $\alpha_i \mid X_i \sim N(0,\sigma_{\alpha}^{2})$. The unobserved individual effect α_i is assumed to be independent of X_i. The presence of such a time invariant component implies a non-diagonal covariance matrix of the composite error term $u_{it} = \alpha_i + \varepsilon_{it}$, with $Var(u_{it}) = \sigma_{\alpha}^{2} + 1$ and $\rho = Corr(u_{it}u_{is}) = \sigma_{\alpha}^{2}/(\sigma_{\alpha}^{2} + 1)$, interpretable as the proportion of the overall error variance which is explained by the individual effect. The likelihood function of the random effect probit model relies on the probabilities:

$$\Pr(y_{it} = 1 \mid X_i, \alpha_i) = \Pr(y_{it} = 1 \mid x_{it}, \alpha_i) = \Phi(x_{it}^{'}\beta + \alpha_i)$$

Integrating out the unobservable component introduces an integral of the above probabilities in each individual contribution to the sample likelihood function.

Estimation by Maximum Likelihood of both models can be automatically implemented by STATA (version 8), although the maximisation of the second model likelihood function requires the evaluation of an integral by Gaussian Quadrature (Butler and Moffit, 1982). The first model ignores individual heterogeneity, but provides nevertheless consistent estimates of the parameters under the classical assumption of incorrelation between regressors and error terms. Indeed, the inference requires the use of a robust covariance matrix to take into account the error serial correlation induced by the unobserved individual component not varying with time. The second model allows us to incorporate the individual unobserved effects, which is typically found to represent a very substantial component in estimated female labour participation equations (Hyslop finds that the individual-specific random effect accounts for 78 percent of the latent error variance) but relies on two main assumptions: 1) incorrelation between the individual heterogeneity and the regressors, 2) strict exogeneity of the explanatory variables conditionally on the individual heterogeneity component. Whenever either of these hypotheses fails, the estimates we obtain for the effects of the children on labour supply are not consistent (notice that in this instance the pooled probit would suffer from the same problem). We cope with the first limitation adopting as a final specification a correlated random effect probit models (CRE), which allows the unobserved preferences for working to be correlated with observed fertility. We follow Chamberlain's suggestion (1984) and postulate:

$$\alpha_i = \bar{z}_i^{'}\xi + \eta_i \qquad (3)$$

where: $\bar{z}_i$ denotes the time average, $t=1,\ldots T$, of the vector $z_{it} = (z_{i1t},\ldots,z_{iMt})'$ containing the M elements of x_{it} describing the observed fertility; $\eta_i \mid X_i \sim N(0,\sigma_\eta^2)$, *I.I.D.* and independent of X_i.[5] This model can also be estimated resorting the predisposed routines for random effect panel probit estimation in STATA, as it only needs to include the time averaged fertility variables as extra-regressors.

Among the regressors, we include the classical labour supply variables. The educational qualification of the woman enters the model through a set of dummies indicating the highest level of education attained. The woman's age is inserted in quadratic form. To proxy for her husband's earnings we use his age and education.[6] Completed fertility is represented by a set of variables indicating the number of kids in different age classes (0-2, 3-5, 6-11, 12-17). Finally, we included among the explanatory variables a set of year dummies to capture trend effects and a set of territorial dummies – corresponding to the 49 Polish voidvoiships – to capture the effect of the unobserved factors describing the different economic context at a regional level. Some recent applied studies (Newell and Pastore, 2002, Sibley and Walsh, 2002 and Boni, 2002) illustrated the crucial importance of accounting for the strong geographical differentiation in development levels and potential of the Polish regions.

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[5] We also tried the more general alternative specification of the random effect, proposed by Hyslop (1999): $\alpha_i = \sum_{s=1}^{T}(\delta_{1s} z_{i1s} + \delta_{2s} z_{i2s} + \ldots + \delta_{Ms} z_{iMs}) + \nu_i$, but got results very similar to those obtained with Chamberlain's model. We therefore selected the latter model, as its parametrization is more parsimonious.

5. Results

In Tables A.1 and A.2 in the Appendix we report the estimation results relative to the two periods: 1987-1990 (pre-transition) and 1994-1996 (transition period). For each period, we present three sets of results corresponding to the pooled cross section (CS) probit model (equation 1), the random effect probit model (RE) (equation 2), and the correlated random effect probit model (CRE) (equations 2 and 3). Before discussing the details of the evidence we obtained on the determinants of women's employment in the two periods, we notice that each of the estimated models represents an improvement over the benchmark pooled cross section specification.

First, the estimated value $\hat{\rho}$ within the CRE specification is about 0.90, and statistically significant, in both periods. This result shows that the relative importance of the unobserved effect is very high, and that allowing for unobserved heterogeneity represents an important improvement of the model (the huge increase in the log-likelihood with respect to the pooled cross section specification supports this conclusion). Second, when we look at the CRE estimation output, we find that some of the coefficients (the elements of the parameter vector ξ) of the variables entering the expected value of the unobserved effect are statistically significant in both periods. We are also able to reject in both periods the null hypothesis that all the elements of ξ are simultaneously equal to zero. This evidence supports the existence of correlation between the individual effect and the inserted fertility variables, selecting the CRE model against the RE one. A further consideration supporting the adoption of the CRE model as the more adequate to represent women's employment outcomes is based on the evidence we get on the exogeneity status of the fertility variables. In all models the hypothesis of strict exogeneity of the included regressors is maintained, implying that there is not feedback from employment history to future fertility decisions. We follow the indication of Wooldridge (2002) and test for this hypothesis. This is achieved by inserting the 1 year forwarded fertility

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[6] Although the variable 'labour earnings' is included in the dataset, the majority of the observations had missing values; so we preferred to replace it with age and education, labour earnings being essentially a control variable in our case.

variables on the right hand-side and testing whether their coefficients are jointly equal to zero. The results of the tests are reported in Table A.3 in the Appendix. Interestingly, the hypothesis could not be rejected for both panel data models within the CRE framework. This indicates that the specified correlation pattern between the fertility variables and the individual unobserved heterogeneity is able to capture part of the mechanism generating a feedback effect from past participation to future fertility behaviour. For the abovementioned reasons we provide discussion and interpretation of the estimation results obtained with the CRE model, referring to the CS and RE formulations only for some comparative comments.

The coefficients of the fertility variables – the key variables of interest in our analysis – display the usual negative effect on the probability of working in both pre-reform and post-reform periods. Not surprisingly, the size of this effect decreases as the age of the child increases, indicating that a lower amount of mother's time is requested for child-care of older children.[7] The most striking pattern emerging from our results is the diminished effect of the fertility variables observed in the second period. While in the first period all the fertility variables display significant coefficients with the only exception of that conveying information about the number of children aged 11+, during the transition period the only significant effect is associated with the presence of small children (aged 0-2). This finding was unexpected to some extent. A stronger negative impact of the number of children of very young age would have signalled the prevailing effect of the well known drop in child-care facilities in the post-communist period (Chase, 1995). According to our results, the increasing difficulty of combining family and work after transition due to less and less financed child-care does not result in a lower employment probability of mothers with respect to the pre-transition period (or it is totally compensated by forces of opposite sign).

A potential explanation of this pattern could be an increase in part-time employment in the transition period – availability of part-time employment typically mitigates the negative

[7] We incidentally note that within the RE model the estimated coefficients of the fertility variables display a higher absolute value than those obtained within the pooled CS model. This pattern suggests that the unobservable characteristics are positively correlated with both the probability of employment and the realized fertility. Within the CRE specification, the magnitude of the children is attenuated with respect to the RE model: the coefficients that are mostly affected by the change in the estimation technique are those of the dummies indicating the presence of very young children (number of kids aged 0-2 and 3-5). This suggests that it is the

effect of children on women's labour supply. However, there is evidence of no increase in part-time job opportunities in the second period; data provided by enterprises indicate instead that a slight drop in the share of part-time employment occurred, from about 10 per cent before transition to 8.8 per cent in 1998 (Kwiatkowski, Socha and Sztanderska, 2001). Therefore, this result calls for alternative interpretations. A plausible explanation is the change in the general economic environment represented by the transition to a market economy, characterized by tremendous adjustments and restructuring, high layoff rates and low re-employment rates. The resulting uncertainty in the labour market appears to have represented a powerful incentive for women to work despite the presence of young children. On the other hand, in the first period, given the very high female participation rate, the decision of working is strongly dependent on the presence of young children.

To analyse the numerical effect of the presence of children on employment outcomes we present in Table 4 a set of estimated probabilities of working for a 'typical' woman – whose characteristics are described at the bottom of the table – over the analysed years. The reported probabilities are evaluated using the CRE estimated model, averaging over the distribution of the individual heterogeneity component in order to compute average partial effects (see Wooldridge, 2002). More precisely, we look at the percentage change in the (average population) probability of working, observed when the typical woman moves from her condition of childless woman to be a mother of a small child (aged 0-2). We limit the analysis to this age class for the child as it is the only one displaying a significant effect in both periods. The descriptive analysis of section 3 suggests to look at this impact by level of education of the woman.

A first inspection to the estimated probabilities over the years, reveals that – regardless of the presence of a small child and at any level of education – the probabilities are not dramatically different in the two periods. In order to give a proper interpretation of this pattern, we recall that these results refer to a selected sample of women, married and with a partner continuously in employment during the analysed years, characterized as the reference woman. For each profile, a trend effect is evident, making the employment probabilities decreasing in the first period until reaching a minimum in 1990. In the transition period the

presence of very young children that is most likely correlated with the mother's taste for work – a very plausible result.

pattern is reversed and the trend effect becomes positive. This reflects the documented increase in women's aggregate employment over the transition period (in 1998 it was 9.6 per cent higher than in 1992), which some analysts explain with the faster rate of job creation in services with respect to industry – services traditionally employing more women than industry (Kwiatkowski, Socha and Sztanderska, 2001).

Table 4. Impact of having one child 0-2 on the probability of working by educational level. Average population effects from model CRE

	1987	1988	1989	1990	1994	1995	1996
Low education							
P(no children)	0.67	0.66	0.66	0.63	0.68	0.70	0.71
P(1 child 0-2)	0.52	0.51	0.50	0.48	0.56	0.59	0.59
% change	**-22.84**	**-23.36**	**-23.57**	**-24.63**	**-17.53**	**-16.59**	**-16.25**
High school							
P(no children)	0.77	0.77	0.76	0.74	0.71	0.73	0.74
P(1 child 0-2)	0.64	0.63	0.62	0.60	0.60	0.62	0.63
% change	**-17.50**	**-17.98**	**-18.17**	**-19.15**	**-16.19**	**-15.28**	**-14.94**
Professional education							
P(no children)	0.67	0.66	0.65	0.63	0.62	0.65	0.66
P(1 child 0-2)	0.52	0.50	0.50	0.48	0.50	0.53	0.54
% change	**-22.95**	**-23.47**	**-23.68**	**-24.74**	**-19.77**	**-18.80**	**-18.44**
Third level							
P(no children)	0.83	0.83	0.82	0.81	0.85	0.87	0.87
P(1 child 0-2)	0.72	0.71	0.70	0.68	0.77	0.79	0.79
% change	**-14.07**	**-14.51**	**-14.68**	**-15.59**	**-9.97**	**-9.25**	**-8.99**

'Typical' woman: aged 30, partner aged 35, with low education (primary or education or first level secondary education), with no children, residing in the Warsaw region.

Turning to the estimated impact of being a mother of a small child, the comparison between the two sample periods confirms the generally diminished effect in the second period, occurring – with different magnitude – at all levels of education of the woman. Before the transition to a market economy, women with low or professional education were the most penalized by the presence of a small child (the estimated decrease in the probability of working was about 23 per cent). Being highly educated was associated with a lower but still sizable (around 15 per cent) percentage drop in the probability of working. During the transition period, the biggest effect – in absolute terms – was estimated for women with professional education (with a percentage decrease of about 19 per cent), while the likelihood of employment of women with high education was very little affected by the presence of a small child (with a percentage decrease about 9 per cent). These results point out that educational qualifications were a key dimension –eventually more important than family arrangements- in determining employment after transition. There is abundant evidence that during transition the increase in unemployment hit the lowest educated people in particular,[8] while the demand for employees with high level of human capital increased steadily.[9] On the other hand, the demand for employees with primary or incomplete primary education and for graduates from post-secondary schools declined significantly (Kwiatkowski, Socha and Sztanderska, 2001).

6. Concluding remarks

In this paper we have analysed the relationship between motherhood and employment probabilities of Polish women before and during transition using panel data for the 1987-1990 and 1994-1996 period. Before transition, despite the very high female participation in employment, Polish mothers – especially mothers of young children – were substantially less likely to work than childless women. During transition, to our surprise, the negative effect of motherhood was substantially lower than in the pre-transition period. A significant negative effect was estimated only for mothers of children age 0 to 2 and even in this case the coefficient were much smaller than in the pre-transition period. This result goes in the opposite direction of what one would have expected, given the very high female labour market participation in the communist Poland, sustained by the existence of an extensive network of childcare facilities that were dismantled during transition, making women's participation in employment theoretically more difficult.

Our findings are not driven by a change in the association of the 'taste for children' and the 'taste for employment' in the two periods, because our estimates –obtained with correlated random effect panel probit models- control for unobserved heterogeneity in the two

[8] Among the unemployed, the majority had only primary (33.4 per cent) or basic vocational (37 per cent) educational qualifications, while only 2.5 per cent of the unemployed were graduates of universities (Kowalska, 2002).

[9] In the period 1994-1998, the employment of Polish people with higher education rose fastest, by 29.1 per cent, while for those with secondary vocational education rose by almost 17 per cent.

samples. Therefore the most plausible interpretation is that, before transition, in absence of constraints in accessing the labour market, women could realize their preferences for lower employment in cohincidence of motherhood or were ‘forced’ out of work by the expectations of a traditional society that required mothers to care for their children rather than working. On the other hand, during transition, the uncertainty in the labour market, the process of restructuring, and the need of relying on two earners in the household turned out to be more powerful pushing factors than the deterioration of childcare facilities for women to work. This is not to say that the increasing difficulties in accomodating mothers’ necessity of combining paid work and family responsibilities were without negative consequences for women’s employment – it rather stresses how powerful the demand side constraints were during transition in shaping women’s employment opportunities.

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APPENDIX

Table A.1 – Estimation results, pre-transition period (1987-1990)

	Pooled Cross Section Probit model (1)		Random Effect Probit model (2)		Correlated Random Effect Probit model (3)	
	Coefficient	*S.e.*	Coefficient	*S.e.*	Coefficient	*S.e.*
Age	0.203*	*(0.121)*	0.507***	*(0.178)*	-0.186	*(0.184)*
Age squared	-0.003	*(0.002)*	-0.007***	*(0.003)*	0.002	*(0.003)*
High school	0.469***	*(0.134)*	1.150***	*(0.207)*	1.056***	*(0.180)*
Professional education	0.231*	*(0.120)*	0.079	*(0.196)*	-0.020	*(0.174)*
University	1.004***	*(0.234)*	2.160***	*(0.328)*	1.803***	*(0.314)*
Number of kids 0-2	-0.715***	*(0.083)*	-1.517***	*(0.127)*	-1.341***	*(0.167)*
Number of kids 3-5	-0.408***	*(0.059)*	-0.923***	*(0.104)*	-0.693***	*(0.177)*
Number of kids 6-11	-0.182***	*(0.053)*	-0.592***	*(0.085)*	-0.418**	*(0.180)*
Number of kids 12-17	0.034	*(0.064)*	-0.424***	*(0.101)*	-0.281	*(0.196)*
Average number of kids 0-2					-1.359***	*(0.270)*
Average number of kids 3-5					-0.868***	*(0.234)*
Average number of kids 6-11					-0.010	*(0.200)*
Average number of kids 12-17					-0.251	*(0.233)*
Year 1988	-0.066*	*(0.040)*	-0.175	*(0.116)*	-0.096	*(0.120)*
Year 1989	-0.110**	*(0.050)*	-0.238**	*(0.121)*	-0.132	*(0.131)*
Year 1990	-0.200***	*(0.059)*	-0.462***	*(0.126)*	-0.327**	*(0.147)*
Partner high school education	-0.095	*(0.141)*	-0.332	*(0.205)*	0.151	*(0.202)*
Partner professional education	-0.119	*(0.116)*	-0.216	*(0.165)*	-0.070	*(0.158)*
Partner university education	-0.475**	*(0.215)*	-1.036***	*(0.304)*	-0.708**	*(0.305)*
Age of the partner	-0.009	*(0.075)*	0.122	*(0.118)*	0.350***	*(0.123)*
Age of the partner squared	0.000	*(0.001)*	0.000	*(0.002)*	-0.002	*(0.002)*
Constant	-3.046*	*(1.755)*	-11.378***	*(2.753)*	-2.897	*(2.845)*
$\hat{\rho}$			0.887***	*(0.010)*	0.907***	*(0.008)*
Observations	3452		3452		3452	
Number of groups			863		863	
Pseudo R squared	0.1788					
Log-likelihood	-1806.35		-1209.1		-1197.26	
Likelihood ratio test of ρ=0			1194.51		1207.43	
($Pr>\chi^2$)			(0.00)		(0.00)	
Joint significance of average number of kids variables in CRE					42.52	
($Pr>\chi^2$)					(0.00)	

Notes: *** significant at the 1% level; ** significant at the 5% level; * significant at the 10% level. The base categories of the dummy variables are: primary education or first level secondary education; no children or children over 18 years old; year 1987; partner with primary education or first level secondary education. We also control for geographical differences by including 49 dummies corresponding to the regions (voivodships). Standard errors are adjusted for clustering on individuals.

Table A.2 – Estimation results, post-transition period (1994-1996)

	Pooled Cross Section Probit model (1)		**Random Effect Probit model (2)**		**Correlated Random Effect Probit model (3)**	
	Coefficient	*S.e.*	Coefficient	*S.e.*	Coefficient	*S.e.*
Age	0.132	*(0.093)*	0.402**	*(0.163)*	0.450***	*(0.167)*
Age squared	-0.002	*(0.001)*	-0.004*	*(0.002)*	-0.006**	*(0.002)*
High school	0.187	*(0.115)*	0.364*	*(0.212)*	0.313	*(0.221)*
Professional education	-0.109	*(0.109)*	-0.531**	*(0.236)*	-0.506**	*(0.220)*
University	0.722***	*(0.170)*	1.963***	*(0.366)*	1.927***	*(0.349)*
Number of kids 0-2	-0.385***	*(0.073)*	-1.174***	*(0.148)*	-1.056***	*(0.207)*
Number of kids 3-5	-0.235***	*(0.059)*	-0.343***	*(0.114)*	-0.052	*(0.202)*
Number of kids 6-11	-0.086*	*(0.045)*	-0.131	*(0.107)*	0.012	*(0.186)*
Number of kids 12-17	0.023	*(0.044)*	0.024	*(0.096)*	-0.031	*(0.159)*
Average number of kids 0-2					-0.094	*(0.265)*
Average number of kids 3-5					-0.765***	*(0.258)*
Average number of kids 6-11					-0.348	*(0.221)*
Average number of kids 12-17					-0.064	*(0.186)*
Year 1995	0.060**	*(0.029)*	0.179*	*(0.093)*	0.217**	*(0.095)*
Year 1996	0.089**	*(0.037)*	0.214**	*(0.096)*	0.298***	*(0.101)*
Partner high school education	0.039	*(0.128)*	0.424	*(0.298)*	0.689***	*(0.259)*
Partner professional education	0.026	*(0.111)*	0.337	*(0.229)*	0.267	*(0.200)*
Partner university education	-0.002	*(0.168)*	0.390	*(0.339)*	0.310	*(0.353)*
Age of the partner	-0.016	*(0.071)*	-0.038	*(0.129)*	0.149	*(0.127)*
Age of the partner squared	0.000	*(0.001)*	0.000	*(0.002)*	-0.002	*(0.002)*
Constant	-1.604	*(1.138)*	-4.859**	*(2.479)*	-8.553***	*(2.561)*
$\hat{\rho}$			0.901***	*(0.010)*	0.910***	*(0.009)*
Observations	3933		3933		3933	
Number of groups			1311		1311	
Pseudo R squared	0.0944					
Log-likelihood	-2225.47		-1608.13		-1601.41	
Likelihood ratio test of ρ=0, (Pr>χ^2)			1234.68 (0.00)		1239.64 (0.00)	
Joint significance of average number of kids variables in CRE (Pr>χ^2)					10.76 (0.029)	

Notes: *** significant at the 1% level; ** significant at the 5% level; * significant at the 10% level. The base categories of the dummy variables are: primary education or first level secondary education; no children or children over 18 years old; year 1996; partner with primary education or first level secondary education. We also control for geographical differences by including 49 dummies corresponding to the regions (voivodships). Standard errors are adjusted for clustering on individuals.

Table A.3 Tests of strict exogeneity of fertility variables

	χ^2	$Pr>\chi^2$
Pre-transition period (1987-1990)		
Model (1) – pooled cross section model	6.86	0.1436
Model (2) – random effect model	4.40	0.3548
Model (3) – correlated random effect model	46.02	0.0000
Transition period (1994-1996)		
Model (1) – pooled cross section model	3.56	0.4685
Model (2) – random effect model	20.26	0.0004
Model (3) – correlated random effect model	22.23	0.0002

Notes: this is a test of joint significance of the 1 period forwarded variables: number of kids 0-2, number of kids 3-5, number of kids 6-11, number of kids 12-18. The required estimation involves the loss of the last year in both samples, corresponding to 863 observations in the first period, 1311 in the second.