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FEMALE LABOUR FORCE PARTICIPATION IN THE 1980s: THE CASE OF SPAIN*

Olympia BOVER

Banco de España

Manuel ARELLANO

CEMFI

This paper is motivated by the large increase in the labour force participation of prime age women that, like in many other countries, took place in Spain during the last decade. We develop an empirical time series participation equation for women aged 25-44, in which education and fertility are treated as endogenous variables. We regard participation, education and fertility as variables that jointly respond to changes in the demand for skilled non-physical labour and the wage structure. Although we find a significant business cycle effect on participation, we conclude that the dominant forces determining the increase in participation during the sample period were structural factors that shifted female earnings potential.

1. Introduction

The Spanish female labour force participation rate, which was 43.6 in 1991, is among the lowest of all OECD countries along with Ireland and Italy (see Table A1 in the Appendix). Participation has been steadily growing, but just a modest 9-point rise between 1973 and 1992, compared to the 17-point increase in the US and 23 points for Canada (countries which, with a participation rate of 52.7 and 44.9, respectively, in 1973, were at a substantially higher level), or the increase of 24 points in the Netherlands, having started in 1973 at a slightly lower participation rate than Spain. However, since these trends compound generational, life cycle and business cycle effects, it is very difficult to interpret the pattern of change in female participation rates without a breakdown by age group. This is specially so in the case of Spain due to the acute changes across generations that have taken place during the last two decades. For example, for the cohort of

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women born in the period 1942-1946 the percentage of those with a university degree was 1.5 percent while for those born between 1957 and 1961 the figure is 8 percent. In addition, fertility rates almost halved from one cohort to the other. In effect, looking at the change in participation rates for women aged 25-44 provides a very different picture. At the beginning of the 1970s the participation rate of this age group was below 30 percent and below the rate for the total female labour force; during the 1970s it increased around 4 points, but in the period 1981-1991 the increase was slightly over 22 points!

This upward trend in prime age female participation started at a period of rapid destruction of employment in the Spanish economy, with rising unemployment that reached 21 percent of the labour force in 1986. This suggests that the dominant factors affecting participation are not cyclical but long term. Indeed, from the evidence presented in this paper, the changes in the 1980s appear as a distinctive structural phenomenon. Measuring to what extent this is so, relative to the contribution of cyclical factors, is important for understanding the trends in the Spanish labour market and for policy.

The increases in female participation rates in all developed countries over the 1980s are probably the response to technical changes in production practices, and to the resulting increase in the demand for skilled non-physical labour and, in particular, the demand for female workers. In countries where participation rates were already high, life cycle factors would be expected to play an important role in determining participation decisions. However, we feel that generational and time series variation are still the dominant features of the recent changes in Spanish female participation [1]. Thus, we are not so much interested in following the sequences of participation decisions of particular adult individuals (as one would do with micro panel data), for which education and children would be mostly predetermined, but rather in following the labour supply decisions of different cohorts and their interactions with education and fertility decisions. The kind of question we are concerned with is as follows: why is a prime age woman in 1990 much more likely to participate in the labour market than a woman of the same age group in 1980, once cyclical factors have been accounted for? For this reason, time series data for a narrowly defined age group incorporate the variation we are most interested in analyzing. In this paper we focus on the evolution of the participation of women aged 25 to 44, using quarterly averages from 1976-III to 1991-IV and annual regional data over 1980-1990, from a sequence of Spanish labour force surveys. No doubt, some of the cross-sectional variation in the individual data would also be very valuable, but at the same time it would introduce much heterogeneity, measurement error and computational problems, given the very large cross-sectional sample sizes available, which would call for a rather more ambitious project than the present one.

[1] It is likely that future cohorts of Spanish women will have to take more time out of the labour market for child care than the current cohort, which has often benefited from high-quality, free child care from their non-working mothers or other relatives.

TABLE 1
Female labour force participation rates in Spain

	Aged 25-44	Aged 25-54	Aged 25-44 with university education	Total (Age ≥ 16)
1976	30.20	29.63	71.96	28.48
1977	30.24	29.46	74.28	27.85
1978	30.75	29.57	75.50	27.68
1979	31.75	30.24	78.06	27.56
1980	32.53	30.62	83.18	27.19
1981	32.60	30.42	83.92	26.81
1982	34.71	31.76	83.66	27.23
1983	36.96	33.29	82.67	27.82
1984	38.14	34.03	83.48	27.72
1985	39.76	34.97	83.12	27.80
1986	41.20	36.38	84.72	28.47
1987	45.71	40.15	84.92	31.07
1988	48.77	42.93	86.34	32.51
1989	51.10	44.85	86.35	32.76
1990	53.40	46.92	85.89	33.36
1991	55.18	48.64	87.38	33.60

In Table 1 we present participation rates for various age groups. We can see how the inclusion of older women in the overall rate introduces a significant amount of inertia that masks the evolution of the participation behaviour of prime age women.

We regard time series changes in female participation, education and fertility as variables that jointly respond to changes in the demand for skilled workers and to wage differentials. Therefore, we measure the effects of education and fertility on participation treating those variables as endogenous and using as instruments exogenous shifters of the relative demand for skilled labour and the wage structure. In addition, contrary to standard practice in previous work, we do not distinguish between married and non-married women, which we consider to be an essentially endogenous grouping. The reason is that we would expect participation, education and fertility decisions to have an effect on marriage decisions. If rising education and participation have a negative effect on the proportion of married women, a time-series analysis of participation of married women would suffer from endogenous self-selection.

Concerning previous work on Spanish female labour supply trends, Hernández and Riboud (1985) studied participation from 1900 to 1980 and therefore their sample stops before the large increase in participation during the 1980s we are most concerned with. However, their theoretical framework and interpretation of results are close to ours. Novales (1989) presented a detailed description of the time series changes in female participation and female employment by occupation. Albarracín and Artola (1989) and Novales and Mateos (1990) emphasized the potential influence of labour market conditions on women's participation. Novales and Mateos (1990) also considered the increase in education as a potential determinant of participation. Both studies use time series transfer function models, which focus on the modelling of the dependence over time in participation changes.

The paper is organized as follows. Section 2 describes the model and the data. Econometric methods and results are presented in Section 3. Finally, Section 4 states the conclusions.

2. The model and the data

Since the work of Mincer (1962), economists tend to attribute the secular changes in participation to changes in real wages. The basic model depicts women as choosing between market work, work at home, and leisure. If the market wage rises relative to productivity at home, market participation increases provided the positive substitution effect is greater than the negative income effect due to higher family income, which would increase the demand for leisure.

Despite the central role that the theory assigns to wages, it is difficult to choose an appropriate wage variable for a time series empirical participation equation. Even if wage series were available by gender (in Spain a measure of female average wages did not start until 1989), using average wages of working women is problematic. As Mincer (1985) pointed out, the change in average female wages may be a result of changing self-selection into the labour force as the latter grows, rather than a cause of labour force growth. On the other hand, at a period of rapid generational change, average wages of workers may underestimate the average potential earnings for all women in our chosen age band given the appearance of possibly more educated new cohorts. As a consequence, in this paper we use education variables, instead of the average wage of working women, as indicators of all women's potential earning power. In most of the specifications we also include aggregate wages, but we would expect them to capture the exogenous general effect related to the prevailing economic conditions, whereas education would partly reflect the fact that women's wages are growing faster than men's because the increase in women's education is stronger than for men. Although we take the view that education only affects participation through potential earnings, whether this is so or, on the contrary, there is an additional effect of education on participation given potential earnings, is something we cannot test. The existence of such additional effects would alter the interpretation of the estimated coefficients.

Obviously, education is also a simultaneous decision together with participation; partly you become educated because you would like to participate. In our analysis we do consider education as an endogenous variable and we face the nontrivial problem of finding adequate instrumental variables. In the empirical specifications we distinguish between secondary education and university education (treating both as endogenous), since their effect on participation could be potentially different.

Plausibly important determinants of the observed increase in women's education are the increase in the demand for skilled labour and the increase in the returns to skill. Rising wage differentials by skill during the 1980s have been documented for several countries (eg. Katz and Revenga (1989), Katz, Loveman and Blanchflower (1993), Schmitt (1993) or Gosling, Machin and Meghir (1994)) and particularly for the US (see Mincer (1991), Krueger (1993), Bound and Johnson (1992) and Murphy and Welch (1992)). For Spain, Revenga (1991) reports that, after a narrowing of wage differentials between 1963 and 1977, from 1977 wage differentials according to education increased substantially. The potential explanations of these changes that have received attention in the literature include, from the demand side, the changing pattern of international trade (with increased trade deficits in the 1980s in the developed countries) associated with a decrease in manufacturing employment, skill biased technological change, and the decline in union power. Specifically, the instruments motivated by demand side considerations that we use are real expenditure in R & D, trade balance excluding food, industry real added value and services real added value (see the Appendix for a precise definition of these variables). We also use as an instrument reflecting changes in the composition of demand the percentage of employment in Public Administration. Indeed, during the 80s there was in Spain a considerable increase in the proportion of people employed by the public sector, partly due to the development of the regional authorities. Supply side factors influencing the evolution of the returns to skill have also been considered but most of them would be endogenous for our problem. However, a potential exogenous determinant of the increase in education is the increase in the supply of (highly subsidised) higher education, which has been substantial in Spain during the 1980s. In this respect, we also use as an instrument a measure of the change in the supply of university courses in Spain.

Our equation also includes a fertility variable as a determinant of participation. The increase in participation has been associated with a decrease in fertility rates. To some extent, children and labour market status are two joint decisions, and as such fertility rates in the participation equation will be treated as an endogenous variable. We would expect fertility to have a negative effect on female participation rates once its endogeneity has been taken into account, given that changes in the demand for women jobs and earnings potential will also affect participation through fertility given education, by triggering changes in life-cycle patterns across cohorts. Aside from the previous more obvious effects, the fertility variable may also be capturing potential wage effects. Indeed, Becker (1985) argues that a decrease in fertility would increase hourly earnings through an increase in work effort. A reduced fertility rate may also increase wages as a consequence of an increase in work experience.

Lastly, female participation may depend on the cyclical conditions of the economy. To allow for this effect the aggregate unemployment rate will be included in the equation as an exogenous determinant. Obviously, there will be some feedback from participation to unemployment, but given that we concentrate on the participation of a specific segment of the population (women of a certain age), its effect on the aggregate unemployment rate will be small and, hopefully, negligible. Nevertheless, in some of the estimated equations we treat the total unemployment rate as an endogenous variable, using the industry unemployment rate or the unemployment rate in the building sector as instruments. Moreover, as explained above, in most of the specifications we include an aggregate wage variable as a further indicator of the prevailing economic conditions, given the inclusion of education and fertility, that would capture the structural trends in the evolution of female potential earnings.

Turning to the description of the data and the definition of variables, our participation variable refers to a relatively narrow band of prime age women. This choice is motivated by the desire to explain changes in participation decisions across certain cohorts that we regard as being particularly important, while abstracting from other aspects that influence the overall participation rates but are of less interest to us. We adopted an age band starting at 25, to avoid picking the increase in women delaying participation due to college education, and ending at 44, to leave out older women whose participation decisions are mostly predetermined by past conditions.

We work with two different data sets. The first one consists of quarterly aggregate observations from 1976-III to 1991-IV, while the second contains annual panel data for the 17 Spanish regions, from 1980 to 1990. These are

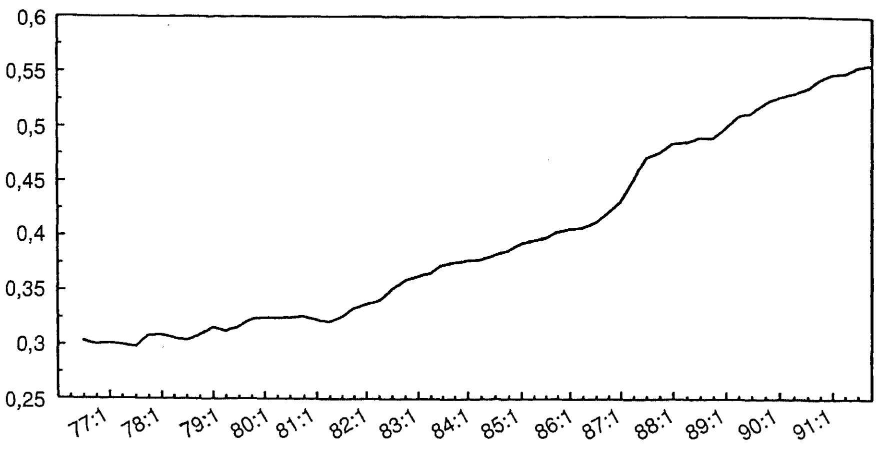

Figure 1
Labour force participation
Women aged 25-44

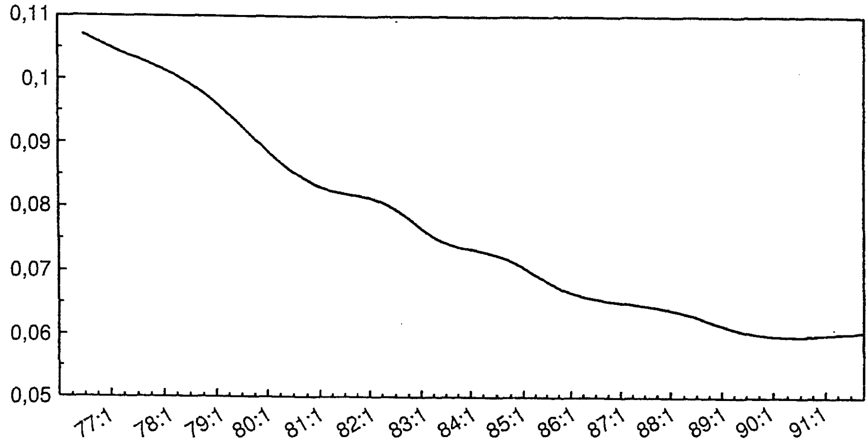

Figure 2
Fertility
Women aged 25-44

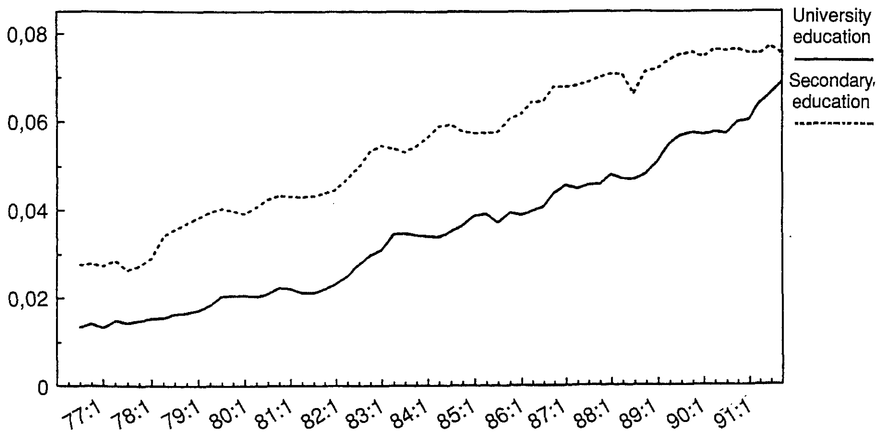

Figure 3
University and Secondary Education
Women aged 25-44

the longest sample periods available to us. The lack of certain variables at the regional level (in particular the variables used as instruments), will prevent us from estimating all the specifications with the regional data, but the regional estimates will serve as a check of the findings obtained with the aggregate quarterly series.

The main source of data are the «Encuestas de Población Activa» (EPA), the Spanish quarterly Labour Force Surveys, which provide information at different levels of disaggregation (in particular by gender and five year age bands), both at the national and at the regional levels. From these sources we have constructed, by age bands aggregation, our variables for education, labour force and population referring to women aged between 25 and 44. The remaining variable that refers to women aged 25 to 44, the fertility variable, has also been constructed from 5 year bands data. This variable measures the percentage of births in a given period and is, therefore, a flow measure of fertility. Figures 1 to 3 plot the quarterly participation and fertility rates, and the two education variables that are used in the empirical analysis. For more details on the construction of these and the other variables, see the Appendix.

Unfortunately, in the second quarter of 1987 the definition of labour force participation used in the EPA changed, with the result that the number of individuals counted as being participants increased (the definition of employed also changed, leaving the number of unemployed essentially unaltered). Despite the fact that the INE provided homogeneous series for the whole period, there is a noticeable change just at that date. In order to take into account this change, we introduced two dummy variables. The first dummy allows for a change in the level of the series from 1987-II onwards, while the second allows for an abnormal observation in that particular quarter. This is a particularly delicate period over which to introduce dummy variables because factors other than definitional changes may be at work, but the difference in the magnitude of the observations at that time was too obvious.

3. Econometric methods and results

3.1. Econometric models

Since a participation rate p is a variable bounded between zero and one, we wish to specify a functional form for the participation equation that also has this property. If we were interested in the regression of p given a set of conditioning variables x, a natural specification would be

$$E(p \mid x) = F(x' \gamma)$$

where F is some probability distribution function. The standard choice of F, due to its simplicity, is a logistic function of the form

$$F(r) = e^r / (1 + e^r),$$

which is the one we employ in this paper. The resulting model can be estimated by nonlinear least squares, and indeed some of the estimates that we have obtained use this method. This is the logit model used in binary choice analysis. However, as discussed in the previous section, we are mostly concerned with estimating response functions in which the explanatory variables may be correlated with the error term due to simultaneity. From this point of view, it is more natural to treat observables and unobservables in a symmetric way and to specify an equation of the form.

$$p = F(x'\beta + u),$$

in which u is an error term assumed to be independent of a vector of instruments z, but not of x. This is a commonly used simple model, that leads to a linear instrumental variables equation in the transformed variable $y = F^{-1}(p) = \log [p/(1-p)]$:

$$y = x'\beta + u$$

which can be estimated by the generalized method of moments. In all the equations the regressors are in logs, since this produced a slightly better fit than the specifications in levels, although we also tried squared and cubic powers of some of the variables.

Turning to the specification of the timing of the effects, given the frequency of the data we specified Almon distributed lags for most of the variables, allowing for a maximum lag length of four and an approximating polynomial of order two. This reduces for each variable the number of parameters to be estimated from five to three. We parameterized the Almon lags in such a way that the coefficients represent short run or contemporaneous effects, long run effects and the effect of the fourth lag. Specifically we have:

$$y_t = \sum_{j=0}^{4} \beta_j x_{t-j} + u_t$$

with

$$\beta_j = \alpha_0 + \alpha_1 j + \alpha_2 j^2 \qquad (j = 0, \ldots, 4)$$

We can parameterize the Almon restrictions in terms of β_0, β_4 and the long run effect $\beta^* = \sum_{j=0}^{4} \beta_j$ by noting that

$$\beta_1 = \frac{3}{10}\beta^* - \frac{1}{2}\beta_4$$

$$\beta_2 = \frac{4}{10}\beta^* - \frac{1}{2}(\beta_0 + \beta_4)$$

$$\beta_3 = \frac{3}{10}\beta^* - \frac{1}{2}\beta_0$$

which gives the transformed equation

$$y_t = \beta^* z_{1t} + \beta_0 z_{st} + \beta_4 z_{ct} + u_t$$

where

$$z_{lt} = (3x_{t-1} + 4x_{t-2} + 3x_{t-3})/10$$

$$z_{st} = x_t - (x_{t-2} + x_{t-3})/2$$

$$z_{ct} = x_{t-4} - (x_{t-1} + x_{t-2})/2$$

An alternative specification could be based on Koyck-type distributed lags using lagged values of the dependent variable. However, since we expect the errors to contain slow moving cohort effects and therefore to be serially correlated, lagged dependent variables would be correlated with the errors. Clearly, this would create problems of identification since it is usually difficult to find appropriate instruments to take this correlation into account. Nevertheless, we estimated some equations including the first lag of the dependent variable and using the second lag as instrument. The result was that the lagged dependent variable was not significant and the remaining coefficients did not change.

Our strategy will be to estimate the equations by the generalized method of moments, and to calculate standard errors and test statistics taking into account that the errors may be serially correlated in a unspecified form. The estimates themselves are not ordinary instrumental variable statistics since we use a weighting matrix of the orthogonality conditions that is optimal under error autocorrelation. We regard all the variables as being either stationary or functions of stationary processes and (local) deterministic trends. Any trend would necessarily be a local one, since the original variables are bounded. For such processes, the standard asymptotic approximations to the distributions of GMM estimators and test statistics have been shown to be valid under very general conditions by Andrew and McDermott (1995).

3.2. *Empirical analysis*

Table 2 contains our main instrumental variables estimates, using the generalized methods of moments. Table 3 contains linear and nonlinear least squares estimates of specifications comparable to some of the equations reported in Table 2. The results from the nonlinear specification are essentially the same as those from the linear one, so the nonlinear case will not be discussed further. In Table 3 we also check for the potential importance of the lagged dependent variable. Table 4 presents additional instrumental variables estimates to evaluate the sensitivity of the main results to the addition or the exclusion of some of the deterministic variables (trend, and step and seasonal dummies) from the equation, and to the endogeneity of the unemployment rate.

TABLE 2
Female participation equations
Instrumental Variables Estimates
Dependent variable: log $(p/(1-p))$ p = participation rate age 25-44
Sample period: 1976.III - 1991.IV

	Outside instruments from current period		Instruments current and lagged one period	Outside instruments lagged one period
Constant	−3.026 (4.24)	−2.248 (2.31)	−2.42 (2.15)	−2.393 (3.44)
Secondary education				
long run*	–	–	0.231 (1.13)	–
short run*	–	–	0.262 (1.00)	–
fourth lag*	–	–	0.211 (0.98)	–
University education				
long run*	0.391 (6.69)	0.414 (5.00)	0.308 (2.77)	0.441 (8.16)
short run*	0.015 (0.16)	0.020 (0.16)	−0.144 (0.66)	0.145 (1.82)
fourth lag*	−0.052 (0.32)	−0.202 (0.92)	−0.124 (0.47)	0.148 (1.29)
Fertility				
long run*	−1.093 (6.03)	−1.041 (4.05)	−0.998 (3.87)	−0.921 (5.43)
short run*	−2.095 (3.83)	−1.405 (1.91)	−1.713 (2.22)	−1.664 (3.48)
fourth lag*	−3.306 (5.58)	−2.813 (3.43)	−3.212 (3.44)	−2.579 (5.20)
Real wage	0.429 (3.84)	–	0.421 (1.63)	0.434 (3.79)
Unemployment	−0.323 (16.78)	−0.303 (11.47)	−0.357 (7.62)	−0.307 (15.79)
DF87	0.100 (7.65)	0.120 (6.95)	0.084 (2.76)	0.112 (8.64)
D872	−0.053 (2.97)	−0.082 (3.61)	−0.041 (1.11)	−0.044 (2.81)
DS1	0.009 (2.09)	0.006 (0.95)	0.003 (0.34)	0.011 (2.62)
DS2	−0.011 (2.56)	−0.011 (1.74)	−0.020 (1.72)	−0.010 (2.20)
DS3	−0.012 (2.67)	−0.016 (2.44)	−0.010 (0.88)	−0.009 (2.16)
Test of overidentifying restrictions	4.36 (3)	9.53 (4)	0.74 (2)	1.38 (3)
Standard error of the regression	0.013	0.017	0.020	0.013
DW	1.56	1.30	1.83	1.27

Notes:
[1] Estimates and t-ratios (in brackets) robust to first order serial correlation.
[2] Starred variables are treated as endogenous. The instruments not included in the equation are: real expenditure on R & D, total population rate aged 20 to 24, trade balance (excluding food), industry real added value, services real added value, real disposable income, employment in Public Administration, number of university faculties and trend. In column 4 these outside instruments are lagged one period.
[3] All explanatory variables are in logs.
[4] Degrees of freedom for the tests of overidentifying restrictions are in brackets.

TABLE 3
Female participation equations
Least Squares Estimates
Sample period: 1976.III - 1991.IV

	OLS		NLS	Linear model with lagged dependent variable	
	With secondary education	Without secondary education		OLS	Instrumenting only $\log[p/(1-p)]_{-1}$ using $\log[p/(1-p)]_{-2}$
Constant	−2.491	−1.886	−2.493	−1.671	−2.00
	(6.12)	(3.69)	(6.19)	(4.71)	(5.51)
Secondary education					
long run	−0.055	–	−0.060	–	–
	(1.10)		(1.22)		
short run	0.021	–	0.021	–	–
	(0.67)		(0.64)		
fourth lag	0.045	–	0.043	–	–
	(1.36)		(1.30)		
University education					
long run	0.462	0.468	0.464	0.337	0.419
	(15.82)	(12.00)	(15.98)	(5.88)	(5.29)
short run	0.152	0.159	0.158	0.122	0.132
	(3.79)	(3.55)	(3.95)	(3.89)	(3.24)
fourth lag	0.039	0.058	0.040	0.028	0.043
	(1.31)	(1.42)	(1.41)	(1.08)	(1.28)
Fertility					
long run	−0.925	−0.832	−0.927	−0.698	−0.826
	(11.55)	(7.53)	(11.7)	(7.15)	(7.73)
short run	−1.661	−1.385	−1.615	−1.132	−1.429
	(6.24)	(4.30)	(6.07)	(3.69)	(3.99)
fourth lag	−2.535	−2.294	−2.484	−1.870	−2.277
	(9.50)	(7.94)	(9.54)	(5.08)	(5.25)
Real wage	0.488	0.303	0.440	0.232	0.322
	(4.84)	(2.20)	(4.88)	(2.00)	(2.65)
Unemployment	−0.293	−0.290	−0.292	−0.224	−0.277
	(22.66)	(18.83)	(23.07)	(6.79)	(6.39)
DF87	0.119	0.124	0.120	0.095	0.111
	(16.27)	(13.31)	(17.16)	(7.64)	(7.87)
D872	−0.052	−0.066	−0.053	−0.030	−0.047
	(9.18)	(9.61)	(9.76)	(2.32)	(3.29)
DS1	0.010	–	0.010	0.006	0.009
	(2.41)		(2.55)	(1.39)	(1.91)
DS2	−0.011	–	−0.010	−0.012	−0.011
	(2.90)		(2.63)	(2.99)	(2.61)
DS3	−0.009	–	−0.008	−0.009	−0.010
	(2.98)		(2.63)	(2.59)	(2.81)
Dependent variable lagged one period	–	–	–	0.234	0.072
				(2.33)	(0.54)
Standard error of the regression	0.011	0.014	0.002	0.011	0.011
DW	1.67	1.74	1.69	1.86	1.62

Note: [1] t-ratios (in brackets) robust to heteroskedasticity.

Our results show that university education has a very significant long run effect on participation, with an estimated elasticity of 0.23.[2] The short run effect becomes insignificant as soon as endogeneity is considered (compare the first columns of Tables 2 and 3).[3] In contrast, secondary education does not have any significant effect on participation given college education. This is so for both instrumental variables and least squares estimates (column 3, Table 2, and columns 1 and 3, Table 3, respectively). In this regard, we interpret the university education variable as a successful indicator of earnings potential. The coefficient of the secondary education variable might be affected by an anomalous observation in 1988-III for which we do not have an explanation. Nevertheless, interpolating this observation did not change the results. It should be noted that, since we need instruments for the entire distributed lag of endogenous explanatory variables, in the equation that includes secondary education we used outside instruments lagged one period as an additional instrument to secure formal identification.

Fertility has a substantial and highly significant effect on participation, both in the long run (elasticity -0.65) and in the short run (elasticity -1.2). Moreover, these effects persist after instrumenting. The strong short run effect remaining after endogeneity is taken into account points to fertility as a predetermining constraint on participation. On the other hand, the large negative estimated coefficient for the fourth lag implies that the restricted coefficients at the intermediate lags are positive (the implied estimates of β_1, β_2 and β_3 are 1.2, 2.1 and 0.7, respectively). This could be the result of bias due to lag truncation. However, our time series are not sufficiently long to experiment with longer distributed lags.

To some people the recent developments in female participation, education and fertility would be a consequence of a general change of «attitudes» that simultaneously determines them. Leaving aside the fact that regarding changing attitudes as exogenous events that occur by chance is not very appealing in general, this approach becomes even less compelling when trying to explain the changes in participation that occurred between the 70s and the 80s. The exogenous forces behind secular changes in female participation should probably be traced to such factors as the reduction in the mortality of children, or changes in the occupational and residential structure of the population. However, one would not expect these factors to have a substantially stronger impact in the 80s than in the 70s. For this reason, it is more likely that the changes in the composition of the demand for labour, with an increase in the demand for women jobs, and the rise in

[2] Given that the regressors are in logs, the sample elasticities, ε_t, will be obtained as

$$\hat{\varepsilon}_t = \hat{\beta}(1 - \hat{p}_t)$$

where $\hat{p}$ represents the participation rate predicted by the chosen model and $\hat{\beta}$ is the corresponding coefficient estimate. We report the mean of the estimated sample elasticities (see Table 5).

[3] In the fourth column of Table 2, the basic equation is re-estimated with the outside instruments lagged one period. As can be seen, this has a negligible effect on the results.

the returns to skill since the 80s, have played a major role. Certainly, the significance and the magnitude of our estimated long run effects of both fertility and education are consistent with this view. This is so because endogeneity is taken into account using as instruments economic determinants of the demand for skill and of the wage structure.

Aside from being sensitive to structural developments like changes in the rate of fertility and university education, we also find evidence that female participation responds to the business cycle and in particular to unemployment. We find a very significant negative effect of unemployment with an elasticity of −0.19. High unemployment rates discourage participation in so far as individuals perceive a lower probability of finding a job. The effect of aggregate wages, that we also interpret as reflecting the prevailing economic conditions, is significant but less so. Furthermore, we noticed that the wage effect was slightly unstable across specifications. The second column of Table 2 reports an estimated IV equation excluding wage effects which shows that the coefficients on the other variables change very little. Given that the education variable captures the positive substitution effect associated with the change in female wages, the estimated effect of aggregate wages would reflect the incentive to participate associated with a favorable economic situation through its impact on the wage level. However, it is not surprising that such effect is small and unstable given the lack of a clear cyclical behaviour of wages during the period under consideration.

The last column of Table 4 presents estimates treating the total unemployment rate as an endogenous variable. We use the industry unemployment rate as an instrument and the results are qualitatively the same. This is also true if the unemployment rate in the building sector is used as an alternative instrument, although in this case some of the coefficients are less precisely estimated.

As shown by the Durbin-Watson statistics, the residuals of our equations are serially correlated. This is to be expected, because the errors are likely to contain unobserved cohort effects that change slowly over time. We tried alternative distributed lags, additional variables and nonlinear transformations of the existing variables, to make sure that serial correlation was not the result of misspecification. Notice that including lags of the dependent variable as explanatory variables is not attractive in our case, since they may capture the effect of slow moving components of the errors and bias the estimates. Nevertheless, in the last two columns of Table 3 we report two equations that include the first-order lagged dependent variable. The first one is estimated by OLS and provides an estimated autoregressive coefficient of 0.23 together with downward biases in all the other coefficients relative to previous specifications. The last column presents IV estimates of the same equation that use the dependent variable lagged two periods as an instrument for the dependent variable lagged one period. The result is that the coefficient on the latter is now insignificant and the remaining coefficients

TABLE 4
Female participation equations
Additional Instrumental Variables Estimates to Check the Robustness of the Results
Sample period: 1976.III - 1991.IV

	1	2	3	4	5
Constant	−9.162	−2.326	−6.453	−6.019	−3.221
	(0.97)	(3.01)	(5.42)	(5.26)	(3.99)
University education	0.618	0.438	0.204	0.243	0.378
long run*	(1.65)	(7.02)	(1.86)	(2.32)	(5.96)
short run*	−0.015	0.109	−0.308	−0.229	−0.097
	(0.06)	(1.17)	(2.07)	(1.65)	(0.61)
fourth lag*	−0.370	0.049	−0.316	−0.206	−0.117
	(0.58)	(0.31)	(1.29)	(0.91)	(0.75)
Fertility					
long run*	−3.411	−0.931	−1.905	−1.802	−1.144
	(0.96)	(4.82)	(6.37)	(6.32)	(5.76)
short run*	−3.877	−1.667	−4.755	−4.634	−2.209
	(1.26)	(2.71)	(4.66)	(4.65)	(3.32)
fourth lag*	−7.206	−2.737	−6.377	−6.116	−3.532
	(1.18)	(4.21)	(6.57)	(6.50)	(5.20)
Real wage	1.528	0.357	0.821	0.798	0.441
	(0.91)	(2.93)	(3.58)	(3.65)	(3.26)
Unemployment	−0.602	−0.302	−0.448	−0.442	−0.329
	(1.41)	(14.26)	(18.64)	(18.93)	(13.86)
DF87	0.055	0.114	–	–	0.094
	(0.72)	(7.81)			(5.52)
D872	−0.077	−0.062	–	–	−0.054
	(1.31)	(3.28)			(2.70)
DS1	0.015	–	0.009	–	0.009
	(1.07)		(0.94)		(1.60)
DS2	−0.016	–	−0.017	–	−0.012
	(1.20)		(1.67)		(2.25)
DS3	−0.017	–	−0.016	–	−0.014
	(1.26)		(1.65)		(2.45)
Trend	−0.028	–	–	–	–
	(0.66)				
Test of overidentifying restrictions	0.21 (2)	27.99 (6)	11.19 (5)	21.90 (8)	2.06 (3)
Standard error of the regression	0.034	0.014	0.028	0.027	0.016
DW	1.06	1.72	1.19	1.24	1.55

Notes:
[1] Estimates and t-ratios (in brackets) robust to first order serial correlation.
[2] The instruments not included in the equation are the same as in Table 2 columns 1 and 2.
[3] Column 5 is similar to column 1 from Table 2, except that the unemployment rate is treated as an endogenous variable, and the industry unemployment rate is included in the instrument set.

are similar to those obtained for the previous models. We take this as evidence that the OLS estimate of the autoregressive coefficient was upward biased due to endogeneity. Serial correlation would bias the ordinary estimates of the standard errors of both least squares and IV estimates. Another effect of serial correlation is to make these estimates to be asymptotically inefficient within their class. To address these problems the main equations reported in Table 2 were estimated by two-step GMM robust to first-order serial correlation of arbitrary form using the method proposed by Newey and West (1987). We also obtained GMM estimates and standard errors robust to second order serial correlation but the results were unchanged. In fact, this was also the case for uncorrected estimates or for estimates corrected for heteroskedasticity. The results did not vary significantly with the different corrections that we tried.

TABLE 5
Means of the estimated sample female participation elasticities

	Wage	Unemployment	University education	Fertility		
			Long run	Long run	Short run	Fourth lag
Column 1 Table 2	0.256	−0.192	0.233	−0.651	−1.247	−1.968
Column 4 Table 2	0.258	−0.182	0.262	−0.548	−0.990	−1.534

So far we have commented on the factors that significantly affect the participation of women aged 25 to 44 in the labour force, with their elasticities being reported in Table 5. However, to pin down the factors that were ultimately responsible for the over 72 percent increase in participation from 1981:I to 1991:IV we have to consider as well the changes in these factors over that period. The most spectacular development was the 211.76 percent increase in the number of women aged 25-44 that have university education. The birth rate for women in this age band fell by over 27 percent, while aggregate unemployment grew by 28.65 percent and real wages by 10.52 percent. Therefore, according to the equation, keeping other factors constant[4], the increase in college education would have increased participation over 30.7 percent from 1981:I to 1991:IV; the downward

[4] To calculate the contribution of factor X_1 to the increase in participation from 1981-I to 1991-IV, keeping other factors constant, we have used the following expression:

$$\frac{[F(\beta_1 \log X_{1\,(1991.4)} + \beta_2 \log X_{2\,(1981.1)} + \dots) - F(\beta_1 \log X_{1\,(1981.1)} + \beta_2 \log X_{2\,(1981.1)})]}{F(\beta_1 \log X_{1\,(1981.1)} + \beta_2 \log X_{2\,(1981.1)} + \dots)}$$

where $F(.)$ is exp (.)/ 1 + exp (.). Another approximate measure (that gives very similar results in our case) is

$$\varepsilon_{1981.1}\, \Delta_{81.1}^{91.4} \log X_1$$

where $\varepsilon_{1981.1}$ is the 1981.I elasticity of participation to X_1 and $\Delta_{81.1}^{91.4}$ represents the increase between 1981.I and 1991-IV.

TABLE 6
Female participation equations from regional panel data
Sample of 17 regions, 1980-1990
Ordinary Least Squares Estimates

	1980	1981	1982	1983	1984	1985	1986	1987	1988	1989	1990	1980-1990
Constant	–0.833	–2.363	–0.534	–2.161	–1.591	–1.003	–2.454	–1.386	–0.208	0.358	–0.056	–1.410
	(1.65)	(3.16)	(0.44)	(1.93)	(1.08)	(0.64)	(2.22)	(1.58)	(0.34)	(0.46)	(0.08)	(3.5)
Secondary education	–	–	–	–	–	–	–	–	–	–	–	–0.008
												(0.10)
University education	0.199	0.178	0.155	0.203	0.199	0.274	0.220	0.210	0.127	0.160	0.276	0.217
	(2.71)	(2.60)	(1.60)	(2.03)	(1.75)	(2.15)	(2.58)	(2.57)	(1.45)	(1.44)	(2.61)	(6.60)
Fertility	–0.786	–1.176	–0.545	–1.080	–0.905	–0.897	–1.102	–0.835	–0.522	–0.402	–0.478	–0.833
	(4.08)	(4.57)	(2.01)	(3.69)	(2.39)	(2.43)	(4.86)	(4.02)	(3.14)	(2.61)	(2.78)	(9.7)
Real wage	–	–	–	–	–	–	–	–	–	–	–	–0.007
												(0.17)
Unemployment	–0.450	–0.244	–0.349	–0.182	–0.214	–0.314	–0.081	–0.155	–0.312	–0.350	–0.141	–0.202
	(5.17)	(1.96)	(1.30)	(0.96)	(1.06)	(1.75)	(0.42)	(0.91)	(2.89)	(2.31)	(1.08)	(4.68)
D87	–	–	–	–	–	–	–	–	–	–	–	0.184
												(3.77)
D88	–	–	–	–	–	–	–	–	–	–	–	0.309
												(8.80)
R^2	0.75	0.70	0.45	0.65	0.60	0.67	0.70	0.59	0.59	0.57	0.60	0.81

Note:
[1] *t*-ratios (in brackets) are robust to heteroskedasticity.

evolution of fertility would have produced a 28 percent increase, and the increase in unemployment would have reduced participation by 5.4 percent. The increase in real wages, aside from being the least stable of the determinants, would have produced only a 3 percent increase in participation. These results are important because they show that, despite the fact that participation responds to business cycle fluctuations, the large increase in female participation that took place in Spain during the 80s was mainly due to structural factors.

As a further check on the robustness of the results, we constructed a data set with annual observations from the 17 Spanish regions («Comunidades Autónomas») for the period 1980 to 1990.

We estimate a model similar to the quarterly specification, except that we do not allow for dynamic effects given the data frequency and the short sample period. The dummies that capture the change in the definition of labour market participation in 1987.II are slightly altered according to the annual frequency of the regional data (see the Data Appendix). Our outside instruments are not available by regions and therefore we only report least squares estimates. The results are shown in the last column of Table 6. As we can see, the same kind of conclusions emerge. In particular, female participation reacts to unemployment, fertility and university education but not to secondary education. As indicated in the previous discussion, the wage effect is unstable as confirmed here where it is completely insignificant.Of particular interest are the year-by-year regressions which show a remarkable stability over time and also that the cross-sectional regional variation captures a pattern of effects that are qualitatively similar to those obtained from the time series equations.[5] The year specific results shown on Table 6 exclude secondary education. This variable was found to be insignificant in all years except two while the estimates of the other coefficients were very similar to the ones reported.

4. Conclusions

In this paper we have analysed the determinants of the formidable increase in the participation of prime age women in the labour market that took place in Spain during the 80s. We conclude that this increase in participation was mainly due to structural factors that shifted female earnings potential, in particular to the increase in university education and the decrease in birth rates, after controlling for their endogeneity. This is important because it implies that the levels of prime age female participation that have been reached and that were never experienced before are here to

[5] Notice that pooled estimates with regional fixed effects are not of particular interest in our context, since we are not concerned with measuring the effect of the explanatory variables on female participation rates relative to unexplained permanent regional differences, but rather in comparing our previous results with those obtained from similar regressions based on cross-sectional regional variation.

stay. Even if these levels remained constant in the future, we would expect the total female participation rate to increase due to the replacement of the older cohorts.

Appendix

Participation rate: $\dfrac{\text{Women in the labour force, aged 25 to 44}}{\text{Population of women aged 25 to 44}}$

Source: «Encuesta de Población Activa» (Tempus). INE, and own.

Secondary education: $\dfrac{\text{Women with secondary education, aged 25 to 44}}{\text{Population of women aged 25 to 44}}$

Source: «Encuesta de Población Activa» (Tempus). INE, and own.

University education: $\dfrac{\text{Women with university education, aged 25 to 44}}{\text{Population of women aged 25 to 44}}$

Source: «Encuesta de Población Activa» (Tempus). INE, and own.

Fertility rate: $\dfrac{\text{Number of births of mothers aged 25 to 44}}{\text{Population of women aged 25 to 44}}$

Source: «Movimiento natural de la población española», INE. The births series is annual and, at the national level has been quarterly interpolated.

Unemployment Rate. Source: from 1987-II, «Encuesta de Población Activa», INE; before 1987-II, (1) national data in García Perea (1991); (2) regional data, «Series Revisadas EPA (1977-1987)», INE.

Industry unemployment rate. Source: «Encuesta de Población Activa», INE, and García Perea and Gómez (1994).

Unemployment rate in the building sector. Source: «Encuesta de Población Activa», INE, and García Perea and Gómez (1994).

Employment in Public Administration (percentage). Source: «Encuesta de Población Activa», INE.

Nominal wage: Source (1) national data: National Accounts Wages for the whole economy, INE, quarterly interpolation, Banco de España (2) regional data,

$\dfrac{\text{«Remuneración Asalariados»}}{\text{Empleo Asalariados}} \left(\dfrac{\text{Workers' Salaries}}{\text{Employed Workers}}\right)$ in «Contabilidad Regional de España», INE.

Real wages: Nominal wages divided by the Consumer Price Index (IPC), INE.

Real income: Source: National Accounts, INE, quarterly interpolated, Banco de España.

Expenditure on R & D: «Gastos intramuros totales en actividades de I + D», Source: «Estadística sobre las actividades en Investigación Científica y Desarrollo Tecnológico», INE, quarterly interpolation, Banco de España. Deflated by real GDP, Source: National Accounts.

Added Value from Industry and Added Value from Services: Source: National Accounts, INE, 1986 constant prices.

Net Real Exports: Real Exports in consumption goods (excluding food) –Real Imports in consumption goods (excluding food). Source: «Economía Española: series históricas», Dirección General de Previsión y Coyuntura, Ministerio de Economía y Hacienda.

Population aged 20 to 24: Total population between 20 and 24 years of age. Source: Encuesta de Población Activa», INE.

Number of university faculties: quarterly interpolation of a series elaborated by Indalecio Corugedo on the basis of «Estadística de la Enseñanza en España» (INE).

Dummy variables:

For quarterly aggregate data:

- DF87 = 1 from 1987-II to 1991-IV
  = 0 otherwise

- D872 = 1 in 1987-II
  = 0 otherwise

For annual regional data
- D87 = 1 in 1987
  = 0 otherwise

- D88 = 1 from 1988 to 1990
  = 0 otherwise

TABLE A.1
Female participation rates: Total (aged 15-64)
Source: OECD, Labour force statistics 1971-1991 (ed. 1993)

	Spain	West Germany	France	Italy	Nether-lands	Portugal	United Kingdom	Ireland	Canada	United States	Japan	Australia	Finland	Norway	Sweden
1973	34.8	50.3	51.2	28.9	31.1		57.2		44.9	52.7	54.1	47.8	66.0	51.8	63.9
1974	34.9	50.6	51.9	29.4	31.5	52.1	58.0		46.1	53.9	52.4	48.7	67.5	51.2	66.3
1975	33.8	50.8	53.0	29.8	32.0	50.6	58.6	34.8	50.5	54.9	51.7	49.7	67.7	53.6	68.9
1976	34.8	51.0	53.9	30.5	32.2	51.4	60.2		51.4	56.1	51.9	49.8	67.5	57.7	70.1
1977	34.4	51.2	55.0	36.8	32.8	52.3	61.2	34.0	52.5	57.5	53.1	51.2	68.2	58.9	71.4
1978	34.3	51.6	54.8	36.8	33.3	53.1	61.2		54.7	59.3	54.2	50.8	67.9	60.7	72.8
1979	34.3	52.2	56.0	37.9	34.5	55.4	61.2	35.2	56.1	60.5	54.7	50.2	68.9	62.2	74.4
1980	33.8	52.8	56.0	38.9	36.3	54.9	61.7		57.9	61.3	54.9	52.5	70.2	64.2	75.8
1981	33.3	53.1	55.9	39.3	38.4	55.6	61.3	35.8	59.5	62.2	55.2	52.1	71.1	65.8	77.0
1982	33.9	52.9	55.9	40.5	39.4	55.1	60.8		59.6	63.1	55.9	51.7	72.2	65.2	77.6
1983	34.7	52.5	55.6	39.6	40.5	61.1	60.2	37.8	60.8	63.4	57.2	51.7	72.9	67.9	78.3
1984	34.7	52.3	55.9	40.1	40.7	59.4	62.8	36.9	62.1	64.3	57.2	52.6	73.1	68.2	78.8
1985	34.8	52.9	56.1	40.5	41.0	58.8	63.3	36.7	63.5	65.5	57.2	54.1	73.9	70.1	79.7
1986	35.4	53.8	56.9	41.7	41.4	58.5	64.0	37.2	64.6	66.5	57.4	56.4	73.4	73.7	80.5
1987	38.9	54.5	57.0	42.8	48.9	57.7	65.1	38.4	66.1	67.5	57.8	57.4	72.9	74.8	81.7
1988	41.0	55.4	56.9	43.4	50.6	60.0	66.0	37.5	67.5	68.3	58.3	58.8	73.0	74.4	82.5
1989	41.3	55.9	57.4	43.8	51.1	58.7	68.1	37.5	68.2	69.4	59.3	60.6	73.5	72.9	82.5
1990	42.2	57.0	57.6	44.4	53.1	59.7	68.5	38.9	69.0	69.6	60.4	61.9	72.9	72.6	82.7
1991	42.6		57.9	44.3	54.5	62.3	68.2	43.4	68.9	69.6	61.5	61.7	72.1	72.4	81.5
1992	43.6		58.5	44.6	55.5	60.3	68.3		68.4	70.2	62.0	61.9	70.7	71.9	80.0

TABLE A.1 *(continuación)*
Female participation rates: aged 25-54
Source: OECD, Labour force statistics 1971-1991 (ed. 1993)

	Spain	West Germany	France	Italy	Nether-lands	Portugal	United Kingdom	Ireland	Canada	United States	Japan	Australia	Finland	Norway	Sweden
1973	27.9	51.1	54.1	29.5	25.7		58.6		44.0	52.0	54.3	47.5	75.6	53.4	68.9
1974	29.4	51.9	55.8	30.7	27.2	47.9	60.4		45.5	53.8	53.0	48.7	78.2	52.8	71.4
1975	27.9	52.8	57.5	31.3	28.5	46.5	61.0	26.1	50.5	55.0	52.3	49.6	78.5	55.3	74.3
1976	29.3	53.5	59.0	32.2	29.8	48.2	62.1		52.1	56.7	52.9	50.3	78.1	60.0	75.6
1977	29.4	54.1	60.7	37.7	30.9	51.2	63.3	26.0	53.6	58.4	54.3	51.4	79.7	61.5	77.5
1978	29.6	54.6	61.2	37.9	32.0	51.7	63.3		56.6	60.5	55.5	51.8	79.8	63.5	79.3
1979	30.2	55.4	63.0	38.9	34.0	54.8	63.3	27.6	57.9	62.2	56.2	51.4	81.2	66.0	81.1
1980	30.6	56.6	63.8	39.9	36.7	54.9	63.4		60.1	63.8	56.7	53.3	82.7	68.9	82.9
1981	30.4	57.8	64.8	40.7	38.6	56.2	63.4	29.1	62.7	65.2	57.1	53.5	83.4	71.3	84.8
1982	31.8	58.3	66.0	42.4	40.6	56.1	63.6		63.3	66.4	58.1	54.0	84.7	70.7	85.9
1983	33.3	58.3	67.0	42.4	42.4	64.3	63.7	32.8	65.1	67.0	59.5	53.2	85.8	73.2	87.0
1984	34.0	58.5	68.2	43.6	43.4	63.9	66.9	32.1	67.0	68.1	60.0	55.0	86.1	74.2	88.1
1985	35.0	59.5	68.9	44.1	44.4	64.1	67.7	32.0	68.5	69.5	60.3	57.1	86.9	75.8	88.9
1986	36.4	60.3	70.3	45.6	45.3	64.1	68.5	34.1	70.1	70.6	60.8	60.3	87.3	79.3	89.8
1987	40.2	60.8	70.5	46.9	53.5	64.3	69.5	36.7	71.7	71.7	61.4	61.7	87.3	80.4	90.4
1988	42.9	61.8	71.2	47.8	55.7	67.2	70.8	36.8	73.4	72.5	62.2	63.0	87.0	79.7	91.1
1989	44.9	62.6	72.1	48.6	56.3	66.3	72.1	37.2	74.5	73.5	63.2	65.3	86.9	79.0	90.2
1990	46.9	64.1	72.9	49.5	58.5	68.0	73.0	39.6	75.6	73.9	64.2	66.6	86.0	79.2	90.4
1991	48.6		73.8	49.9	60.8	71.2	72.9	47.0	75.8	74.0	65.0	67.1	85.6	79.1	89.1
1992	50.4		74.9	50.8	62.0	70.4	73.4		75.3	74.6	65.4	67.2	84.5	79.1	88.7
Breaks		1987	1982	1977	1981										1987

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Resumen

Este trabajo está motivado por el gran aumento en la participación laboral de las mujeres de edad intermedia que, al igual que en muchos otros países, tuvo lugar en España durante la última década. Desarrollamos un modelo empírico de participación para mujeres de 25 a 44 años, utilizando datos de series temporales, en el cual las variables explicativas de educación y fertilidad se tratan como endógenas. Consideramos participación, educación y fertilidad como variables que responden conjuntamente a cambios en la demanda de trabajo cualificado y a la estructura salarial. A pesar de encontrar efectos cíclicos significativos sobre la participación, concluimos que las fuerzas dominantes que determinaron el incremento de la participación durante el período muestral fueron factores estructurales que aumentaron los ingresos potenciales de las mujeres.

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