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Employment and the Risk of Domestic Violence: Does the Breadwinner's Gender Matter?*

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Abstract

This paper studies the effect on the risk of female victimization of the employment statuses of both partners, conditional on income and a set of sociodemographic characteristics. Using cross-sectional data from the Violence Against Women (VAW) surveys for Spain in 1999, 2002, and 2006, we address the potential endogeneity of employment and income variables using a multivariate probit model. We exploit geographical-level information on employment and unemployment rates by gender and age, and on household income, to identify the parameters of the model. Our estimation results, for which proper account of the endogeneity problem proves critical, show that male partner employment plays a major role in the risk of physical violence, while female employment only lowers the risk of violence when her partner is employed too. The lowest risk of physical abuse appears for more egalitarian couples in which both partners are employed.

JEL classification: J12, D19, J16, C25, C26.

Keywords: intimate-partner violence, employment, discrete choice, multivariate probit, endogeneity.

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1 Introduction

Violence against women perpetrated by their intimate partners is an issue of social and political concern. It is rooted in social inequalities between men and women and a form of gender-based discrimination (Garcia-Moreno et al, 2005). Despite the increased attention paid to it, the problem is still present in many countries. According to the Special Eurobarometer on Domestic Violence against Women from the European Commission (2010), one European woman in four experiences domestic violence at some point in her life, and between 6-10% of women suffer domestic violence in a given year. It is also a widespread problem in the U.S. where, according to OECD (2013), 5.9% of women reported having experienced domestic violence during 2010 and 35.6% in their lifetime.

This problem imposes a significant economic burden on society as a whole, in the form of health care costs, policing and legal costs, or declines in productivity, among others. It can also have harmful consequences for children who witness it in terms of their emotional, cognitive, and behavioral development. Moreover, it is well-documented that domestic violence passes through the generations, in that children who are exposed to violence in their families of origin are more likely to become involved in violent relationships as adults. Therefore, we should consider not only the short term consequences of the domestic violence, but also its long-lasting effects.

One potential way advocated to prevent abuse is to improve the outside opportunities of the woman so as to increase her bargaining power within the household. The pioneering study of Gelles (1976) finds a negative relationship between female resources and intimate-partner violence. However, the existing empirical evidence on this relationship is mixed. The empirical studies differ in the level of disaggregation of the data, the geographical scope of the analysis, the subpopulation under study and, very specially, in the methodological approach, particularly whether endogeneity of both partners' resources is acknowledged or not.

In this paper, we assess the separate and interacted effects of both partners' employment statuses on the risk of intimate-partner violence (IPV), conditional on household income and a set of sociodemographic characteristics. Using cross-sectional data from the Violence Against Women (VAW) surveys for Spain in the years 1999, 2002, and

2006, we consider separately binary indicators for two types of intimate-partner abuse, physical and non-physical. The qualitatively different forms of IPV has been acknowledged by previous studies, which may indicate distinctive correlates and distinctive behaviors for the different types of violence (see MacMillan and Gartner, 1999).

The challenge of identifying the causal effect of employment and income indicators on the risk of IPV originates from the potential endogeneity of these variables. Their discrete nature makes two-stage or instrumental variable methods inappropriate. Thus, we consider a multivariate probit model in which we specify three additional equations for the employment indicators of the woman and her partner, as well as for household income. We rely on the jointly normal distributional assumption and, additionally, exploit exclusion restrictions to identify the parameters of the model. In particular, we exploit exogenous geographical information on the employment and unemployment rates by gender and by age, and on household income. To ascertain the role of the endogeneity in the estimation results, we also estimate univariate probit models for the probability of IPV, which assume exogeneity of the employment and income variables.

If employment were just an indicator of availability to economic resources within the household, the main mechanism linking lack of employment and violence should be the stress derived from the lack of resources within the household. Thus, one would expect that employment statuses of the partners should have little or at most a limited effect on domestic violence, once other measures of economic resources, such as income, are accounted for. Moreover, if in addition to the absolute amount of resources, the relative contribution of each partner affects their corresponding bargaining power, then we should expect woman employment and partner employment to have differential roles in the incidence of intimate-partner violence. However, male employment might also have a "symbolic" role and entail more than just economic resources. As Macmillan and Gartner (1999) point out, "notions of masculinity remain strongly tied to beliefs about being a good provider and breadwinner". From this perspective, the effect of women employment would be different depending on her partner's employment status. Equally, a strong effect of men's employment status as opposed to women's could also reveal the symbolic consideration of men's employment.

Our estimation results show that reckoning endogeneity has critical consequences on the estimation of the causal effects of employment and income on both types of IPV. The main estimation results, which account for endogeneity, can be summarized as follows. There is an asymmetry in the role of employment statuses of the partners in the risk of physical IPV, with male employment playing the major role, and female employment only reducing physical IPV provided that her partner is employed too. The risk of physical violence is significantly lower when the male partner is employed, ranging from 2.9 to 1.9 percentage points less depending on whether the woman is employed or not, respectively. Female employment reduces the probability of physical violence by 2.5 percentage points, provided that her partner is employed too. Regarding non-physical IPV, the risk is significantly reduced by 4.9 percentage points for employed woman whose partners are employed too, while the reduction for employed partners when women are employed too amounts to 2.7 percentage points. Considering that the respective sample rates of physical and non-physical IPV are 3.9% and 8.2%, the marginal effects of employment statuses are quantitatively relevant.

Our results agree with the interpretation that equal couples, in which both partners are employed, tend to have the lowest risk of physical violence. On the contrary, the highest risk arose for couples in which the male partner is not employed, in accordance with the instrumental use of violence by abusing partners predicted by backlash theories: when the male partner feels challenged by a relative improvement in the woman's position within the household, he inflicts violence in order to assert his dominant position. Finally, the household income indicator is not significant, so that the importance of the labor market statuses of the partners matters beyond economic considerations in the risk of partner abuse.

The rest of the paper is organized as follows. In section 2, we provide an overview of previous theoretical and empirical contributions about intimate-partner violence. In section 3 we describe the sources of the data used in the analysis and provide a descriptive analysis of the main data. In section 4 we discuss the methodological appproach, and present the estimation results in section 5. Section 6 summarizes the main findings and concludes.

2 Previous research on intimate partner violence

Family violence has originally being a topic of study in the fields of criminology and sociology, for which intimate-partner violence serves two purposes. The first is expressive, to the extent that some men derive direct utility from the violence. The second purpose of violence is instrumental, thus increasing partner's utility indirectly through the control of female behavior (Gelles, 1974). The major issue of concern has been the relationship between the socio-economic situation and the risk of spousal violence. According to the sociological "absolute resource theory" (Gelles, 1974), diminished household resources lead to conflict that can culminate in female victimization. Thus, violence might be particularly prevalent in households in which the indicators of economic resources point to a difficult situation, as for instance low income households, low educated households, or households in which both partners are unemployed.

The economic research in family violence is relatively recent and scant. The main theoretical contributions consists of game-theoretical approaches which challenge the cooperative, altruistic models set up by Becker (1965, 1973). Tauchen, Witte and Long (1991) and Farmer and Tiefenthaler (1997) consider non-cooperative game models with partners' utilities being functions of their consumption levels and the violence exerted by the male partner, which plays both an expressive and an instrumental role. Improving economic opportunities for the woman outside the relationship might increase the woman's reservation utility. Among such economic opportunities, we should mention improved education, income and labor market participation, but also the odds of a generous divorce settlement and the quality of help services for battered women (Farmer and Tiefenthaler, 1996). It is generally assumed that the female reservation utility is binding and the woman receives net income transfers from her partner, which would mostly correspond to situations in which relative income is dominated by men. Under such case, according with Tauchen et al (1991), changes in male and female income have the opposite effects on violence, which only plays an expressive purpose. Instrumental use of violence would be discarded to the extent that women have the possibility of ending the relationship (Aizer, 2010). An increase in male income allows him to exert more violence, while an increase in female income forces her partner to reduce violence in order to assure her reservation utility. Such situation provides support to fostering women economic empowerment as a way to reduce the risk of abuse.

Notwithstanding, Tauchen et al (1991) remark that when the female reservation utility is not binding, violence exerted by the male partner can also have an instrumental purpose. In this case, violence is not necessarily decreasing with income if the male marginal utility of violence increases when the woman is better off, so that an income increase might lead to an increase in violence. In particular, when the female reservation utility is not binding and there are not income transfers between partners (so that each partner income increase is devoted to the own partner's consumption), whereas an increase in male income would just increase male consumption and decrease violence, an increase in female income might increase female consumption and increase violence.[1] The instrumental use of violence by male partners is on the basis of the sociological theory of "male backlash" (Hornung, McCullough and Sugimoto, 1981; Macmillan and Gartner, 1999), which predicts that an improvement in the economic situation of women might increase the risk of abuse. Exercise of violence emerges as a result of imbalance in access to resources within the household. Male partners can perceive women improvement in outside opportunities as a threat, to which they might respond violently to assert their dominance in the relationship. On the other hand, according to the "extractive theory" the male might exercise violence on a woman with more economic resources to extract a monetary transfer, thus weakening her bargaining position.

The theoretical predictions establishing a decreasing effect of woman economic opportunities on the incidence of violence are confirmed by many empirical studies (Aizer, 2010; Farmer and Tiefenthaler, 1997, Tauchen et al, 1991). However, several studies, particularly for patriarchal cultures where divorce is hardly an option for women, a relative increase in female income may increase the risk of victimization, as it challenges the socially prescribed male dominance, triggering male backlash (Luke and Munshi, 2011).

Female participation in the labor market is largely advocated as an effective way

[1] If there are transfers beteween partners and the male marginal utility of violence is increasing with female consumption, changes in income have the same sign irrespective of whether it is male's or female's income, but the effect of an income increase on violence can be positive, if the income effect were strong enough to allow a greater female consumption.

to improve her bargaining power within the household. There are numerous empirical studies that have analyzed the effect of female employment status on the incidence of violence, with mixed results. One of the first studies is Gelles (1976), who find a negative relationship. Other studies, such as DeMaris et al. (2003), find greater violence towards women who are employed. Interestingly, Kaukinen (2004) finds that the association between a woman's employment and her risk of abuse depends on the employment status of her partner, abuse being more likely if she is employed while her partner is not.

But the main shortcoming of these papers, and others that followed, is that they do not account for the potential endogeneity of labor market status.[2] To overcome this problem, some authors have estimated structural models (Bowlus and Seitz., 2006) or used panel data techniques (Tauchen et al., 1991) to control reverse causality and correlated time invariant unobserved heterogeneity. Other authors have used instrumental variables techniques, such as Villarreal (2007), who uses the level of control exercised by her partner; Chin (2012) who uses the exogenous variation in rural women's working status driven by rainfall shocks and the rice-wheat dichotomy; and Bhattacharyya, Bedi and Chhachhi (2011), who assume that children and family type affect female participation but not violence.

Another strand of the literature focuses on the signalling role of local unemployment rates by gender and its effect in the incidence of partner's abuse. Anderberg, Rainer, Wadsworth and Wilson (2015) and Tur-Prats (2015) find that male unemployment rate has a negative effect on abuse, while the effect of the female unemployment rate is negative. However, they do not consider the effect of the own employment status of the partners on the risk of abuse. Aizer (2010) considers potential, and not actual wages, to estimate the causal effect of the gender wage gap on the risk of abuse by exploiting exogenous changes in the demand for labor in female dominated industries relative to male dominated ones. Specifically, she uses a measure of the gender wage gap at the municipality level, which reflects gender-specific labor demand but not underlying worker characteristics potentially correlated with domestic violence.

Another line of theoretical research has considered the intergenerational trans-

[2]See, for instance, Strube and Barbour (1983), or Yllo and Strauss (1981).

mission of domestic violence within families. Pollak (2004) develops a model with expressive violence in which both men and women are heterogeneous in their corresponding probabilities of committing and suffering abuse, where such probabilities depend on whether each of them were raised in violent homes. The model recognizes the importance of the marriage market and divorce in the level of domestic violence, showing that factors that reduce the incidence of abuse in the short run might have strong effects in the long run because of its impact in intergenerational transmission.

3 Data and descriptive evidence

Our primary source of data consist of the 1999, 2002 and 2006 cross-sectional surveys on Violence Against Women (VAW) for Spain.[3] These surveys were fostered by the First National Response Programme against Domestic Violence[4], established in 1998, which led to subsequent legislative proposals that gave rise in 2004 to the first constitutional law against gender-based violence. This bill not only established harsher penalties for offenders, but also finance public help services and shelters for battered women, promoted training programmes for health professionals and judges, and campaigns in public education institutions and the media to raise awareness about violence against women. After its approval, both the number of legal actions, the number of emergency calls, and the number of women affiliated to the special phone service for victims of abuse has increased (see Table 1). However, this increasing trend has flattened in the aftermath of the recession, following a funding reduction in help resources for victimized women.

The VAW surveys form broad and representative samples, both at the national and at the regional level, of women living in Spain for each corresponding year. The surveys were conducted by phone to adult women (at least 18 years old). The original dataset contains $69,627$ observations. In order to conduct our analysis, we have imposed the following selection criteria. We have restricted the sample to women younger than 65 years old, cohabiting with her partner and not enrolled in school. Moreover, we

[3] Data for 2011 are also available. However, we have excluded them, given that there are important methodological differences, which entail the survey design and the way in which the survey was conducted.

[4] Primer Plan de Acción contra la Violencia Doméstica.

have removed those women reporting more than two relationships. Finally, we have discarded all respondents with missing covariate information. The sample size is thus reduced to $32,410$ observations. The VAW surveys include information on domestic violence, as well as sociodemographic characteristics for the woman and her partner.

The survey undertakes the widely accepted methods to measure exposures to violence by an intimate partner. Gold standard methods to estimate the prevalence of any form of violence are obtained by asking respondents direct questions about their experience of specific acts of violence over a defined period of time, rather than using more generic questions about whether the respondent has been "abused" or has experienced "domestic violence" or "rape" or "sexual abuse", which tends to yield less disclosure (World Health Organization, 2013). The information about domestic violence experienced by women is gathered by means of a multiple choice question module addressed to all women except for those currently either living alone or without a partner. This module presents a list of 26 behaviors that constitute domestic abuse. Women are asked to indicate which, among such behaviors, they have experienced lately, and if so, how often (usually, sometimes, rarely), and by whom (partner, different relatives living at home, other people living at home). We consider, following Anderberg et al. (2015), those behaviors that entail serious abuse by the partner, which, in the Spanish VAW survey, are circumscribed to 13 out of the 26 listed behaviors (see Tur-Prats, 2014). We then construct two binary indicators of IPV (Intimate Partner Violence), Physical and Non-Physical, which takes on value one if the respondent attributed her partner to behave "usually" or "sometimes" to the categories corresponding to physical and non-physical abuse, respectively.[5] Table 2 presents the list of behaviors from which we have constructed our IPV indicators. For descriptive purposes and to allow the comparison with previous studies, particularly for Spain, we have also reported the IPV indicator that considers either physical or non-physical serious abuse.

In Table 3 we report the IPV incidence for the three years we have considered.

[5] The measures under this methodology, however, cannot tell anything about the intensity of the abuse. In particular, the fact that among two women for which the IPV indicator takes on value one but the first woman declares to have experienced a higher number of listed behaviors than the second one does not imply that the first woman faces a more serious situation of abuse. The characteristics of the available IPV information precludes the use of principal components as an alternative to build a synthetic measure of IPV, as the values of such synthetic variable could be hardly interpreted in terms of more or less serious abuse.

About 11.7% of women reported to have experienced some situation of serious abuse. When breaking down into physical and non-physical abuse, we find that about 5% of women reported physical abuse, while 9.3% reported serious non-physical abuse. The incidence has significantly decreased over time, the rates of physical and non-physical abuse ranging from 6.4% and 10.3% in 1999 to 3.9% and 8.2% in 2006, about one year after the constitutional law against gender-based violence was passed by the Spanish Parliament.

In Table 4 we break down the IPV incidence by region. The incidence of any type of IPV ranges from 8.8% to 14.4%, and the differences across regions are statistically significant. It is interesting to mention that most Southern regions, whose average per capita income is clearly below the national average (Andalucia, Castilla-La Mancha, Extremadura, and the African autonomous cities of Ceuta and Melilla) exhibit the highest incidence of IPV, quite above the average value at the national level.

The main summary statistics by IPV status are reported in Table 5. These statistics point out that women, partner and household characteristics differ depending on the presence of abuse, where the stronger differences appear for physical abuse. Women experiencing abuse are significantly older and exhibit substantially lower levels of education than women who are not abused. Equally, the partners of women experiencing abuse tend to be significantly older and with significantly lower levels of education than those of women who are not abused. Particularly, many abused women have not even completed primary studies, while much fewer women with a college degree have experienced abuse. This same pattern in education arises for abusing partners. This differential effect of both partners' age in the incidence of abuse points out an age cohort effect in this phenomenon. The negative correlations betwen both partners' education and abuse also hold when conditioning on the ages of the woman and her partner.

Interestingly, there are strong differences in the income distribution between households with abusing partners and households where abuse is absent. Households with victimized women are substantially poorer, what appears consistent with Gelles' (1974) absolute resources theories by which diminished resources increase the risk of victimization. There are not differences in marital status between abused and non-abused women and in the size of the municipality of residence, and households with abusing partners are just slightly larger than households where abuse is absent.

In Table 6, we report the incidence of IPV in accordance with employment statuses of the woman and her partner, finding significant differences. Specifically, abused women seem to be less likely to work than non-abused women and, to a lesser extent, abusive partners are less likely work than non-abusive partners. We could then think that male partners are less likely to abuse if the woman is working; however, when we consider the interaction between the statuses of the woman and her partner, this does not seem to be the case. In fact, male partners are less likely to abuse when the woman is working, only if the male partner is working too: when the woman is working but her partner is not, the probability of abuse does increase.

In summary, the descriptive evidence indicates that sociodemographic characteristics of women and their partners differ depending on the presence or the absence of abuse. In particular, both partners tend to be older and much less educated, and live in much poorer households. Interestingly, the domestic abuse is correlated with the woman's employment, her partner's employment, and the interaction between employment statuses. Of course, these differences just express marginal correlations, behind which there can be differences in observed and unobserved individual and demographic characteristics, so we cannot necessarily infer the existence of causal effects.

We have also consider complementary data to control for geographical effects and to exploit exogenous variation in order to undertake endogeneity problems in employment and income variables. In order to control for local effects, we have also gathered province-level information on per capita GDP (Source: Spanish National Accounts, National Statistics) and population density (Source: Spanish Census, National Statistics). Regarding employment variables, we have constructed the employment and unemployment rates by year, province, age group and gender (Source: Spanish Labor Force Survey, National Statistics); for our purposes, we have calculated as the fraction of the total population in that age-gender group that is employed or unemployed, respectively. Regarding geographical information on income, we have computed the fraction of poor households, defined as the ones whose income is one standard deviation below the national average, by year, region and size of municipality (Source:

Spanish Consumer Expenditure Surveys, National Statistics).

4 Empirical model

To study the effect of female and male employment statuses on the probability of female abuse, we use nonlinear discrete models. Let IPV_i^* be the latent process that guides intimate partner violence, characterized by the following underlying behavioral model:

$$IPV_i^* = \alpha_0 + \alpha_1 f_i + \alpha_2 p_i + \alpha_3 (f_i \times p_i) + \mathbf{X}_i' \delta + v_i \equiv \mathbf{W}_i' \beta + v_i \tag{1}$$

where X_i is a set of exogenous variables and f_i and p_i are dummy variables for women and her partner employment, and $(f_i \times p_i)$ is the interaction between them. We observe a binary variable, IPV_i, which indicates whether women i experiences IPV,

$$IPV_i = \mathbf{1}(IPV_i^* > 0) = \mathbf{1}\left(\begin{array}{c} \alpha_0 + \alpha_1 f_i + \alpha_2 p_i + \alpha_3 f_i \times p_i \\ +\mathbf{X}_i' \delta + v_i \geq 0 \end{array} \right), \tag{2}$$

where $\mathbf{1}$ denotes the indicator function. This model is known as "dummy endogenous variable model" (Amemiya, 1985).

If $v_i \mid \mathbf{X}_i, f_i, p_i \sim N(0,1)$, model (2) becomes a standard probit model. If f_i, p_i were continuous endogenous variables, provided we had a set of instruments Z for them, such that the distribution of the f_i, p_i conditional on $\mathbf{X}_i, \mathbf{Z}_i$ were Normal, the reduced form for IPV_i would also be a probit model and, therefore, the parameters in (2) could be easily estimated by means of a two-stage method. Nevertheless, the presence of dummy endogenous regressors in a binary choice model makes the analysis differ substantially from that in continuous variable models. More precisely, the binary nature of the endogenous variables precludes the 2SLS approach, since their distribution is not normal.

Given this, we account for the endogeneity problem by considering a multivariate probit model. We must first specify reduced form equations for female and male employment:

$$f_i = \mathbf{1}(f_i^* > 0) = \mathbf{1}(\mathbf{Z}_{1i}' \lambda_1 + \varepsilon_{i1} > 0), \tag{3}$$

$$p_i = \mathbf{1}(p_i^* > 0) = \mathbf{1}(\mathbf{Z}_{2i}' \lambda_2 + \varepsilon_{i2} > 0), \tag{4}$$

where $(v_i, \varepsilon_{i1}, \varepsilon_{i2})$ are assumed to be jointly normally distributed with zero mean vector and covariance matrix

$$\mathbf{\Omega} = \begin{pmatrix} 1 & \rho_{v\varepsilon_1} & \rho_{v\varepsilon_2} \\ & 1 & \rho_{\varepsilon_1\varepsilon_2} \\ & & 1 \end{pmatrix}. \tag{5}$$

Notice that if $\rho_{v\varepsilon_1} = \rho_{v\varepsilon_2} = \rho_{\varepsilon_1\varepsilon_2} = 0$ consistent estimators of the parameters could be obtained by estimating each equation separately. $\mathbf{Z}_{1i}$ and $\mathbf{Z}_{2i}$ are sets of exogenous variables, that include $\mathbf{X}_i$. Although the model is identified by functional form assumptions, exclusion restrictions by which there exist regressors in the employment equations that do not directly affect the domestic violence outcome improve identification of the parameters of the model.

Model parameters are jointly estimated by Maximum Likelihood. The log-likelihood function (LF) of the model (omitting subscript i for the sake of simplicity) depends on the trivariate standard normal distribution function, $\Phi_3(.)$. It contains eight joint probabilities corresponding to the eight possible combinations of the endogenous dummies. The log LF is maximized with respect to $\beta, \lambda_1, \lambda_2, \rho_{v\varepsilon_1}, \rho_{v\varepsilon_2},$, and $\rho_{\varepsilon_1\varepsilon_2}$ and takes the following form:

$$L = \sum_{i=1}^{N} \log \Phi_3(\mu_i, \mathbf{\Sigma}_i),$$

where

$$\mu_i = (K_{i1}\mathbf{W}_i'\beta, K_{i2}\mathbf{Z}_{1i}'\lambda_1, K_{i2}\mathbf{Z}_{2i}'\lambda_2),$$
$$\mathbf{\Sigma}_i = (K_{i1}K_{i2}\rho_{v\varepsilon_1}, K_{i1}K_{i3}\rho_{v\varepsilon_2}, K_{i3}K_{i2}\rho_{\varepsilon_1\varepsilon_2}),$$

with $K_{i1} = 2IPV_i - 1$, $K_{i2} = 2f_i - 1$, and $K_{i3} = 2p_i - 1$.

Evaluating multivariate normal distribution functions by numerical approximation is computationally cumbersome. Instead, we use simulation to approximate them, implementing the most popular simulation method by Geweke, Hajivassiliou and Keane (GHK). This method is based on the expression of the multivariate normal distribution as the product of sequentially conditioned univariate normal distributions, which can be easily evaluated (see Borsch-Supan and Hajivassiliou, 1993).[6]

[6] The simulated maximum likelihood estimator is consistent as the number of draws and the number of observations tend to infinity. Increasing the draws increases accuracy, but at the cost of lengthening run time. According to Hajivassiliou and Ruud (1994) ensuring that the number of random draws is similar to the square root of the sample size ensures that the simulation bias is reduced to negligible.

In our empirical model we have an additional potentially endogenous binary variable, which is an indicator of being a poor household. Thus, we add to our specification a reduced form model with additional exclusion restrictions for the probability of being a low income household, and estimate a tetravariate probit model that takes into account the correlation between the error terms of the 4 different outcomes, domestic violence, female employment, male employment, and low household income.

Finally, in non-linear discrete models the parameters of policy interest are typically the marginal effects of the variables evaluated at certain values of the explanatory variables, for instance the mean. Given that our specification includes an interaction between the employment indicators for both spouses, the marginal effect of female and male employment depends on the value of the partner's employment status. For instance, for the sample mean values of the covariates $\overline{\mathbf{X}}$, the estimated effect of female employment on the probability of IPV when the male partner is employed is obtained by considering the difference

$$\begin{aligned}&\widehat{\Pr}(IPV = 1 \mid f = 1, p = 1, \overline{\mathbf{X}}) - \widehat{\Pr}(IPV = 1 \mid f = 0, p = 1, \overline{\mathbf{X}})\\&= \Phi(\widehat{\alpha}_0 + \widehat{\alpha}_1 + \widehat{\alpha}_2 + \widehat{\alpha}_3 + \overline{\mathbf{X}}'\widehat{\delta}) - \Phi(\widehat{\alpha}_0 + \widehat{\alpha}_2 + \overline{\mathbf{X}}'\widehat{\delta}),\end{aligned}$$

where $\widehat{\alpha}_0$, $\widehat{\alpha}_1$, $\widehat{\alpha}_2$, $\widehat{\alpha}_3$ and $\widehat{\delta}$ are the multivariate model estimates and $\Phi(\cdot)$ denotes the cumulative distribution function of the univariate standard normal distribution. Equally, the effect of female employment when the male partner is not employed is given by

$$\Phi(\widehat{\alpha}_0 + \widehat{\alpha}_1 + \overline{\mathbf{X}}'\widehat{\delta}) - \Phi(\widehat{\alpha}_0 + \overline{\mathbf{X}}'\widehat{\delta}).$$

5 Results

In this section, we discuss our estimation results to assess the effect on the probabilities of physical and non-physical IPV of the employment statuses of both partners, household income, and other measures of socioeconomic status such as partners' education. Our variables for employment status consist of each binary variable on whether the corresponding partner is employed or not, and the interaction between them. Our income variable consists on an indicator of low household income (when the respondent asked that her household income was significantly below the average). We have included sets of dummy variables on each partner's completed education, as well as a binary variable indicating whether the woman is more educated that her partner. We control for additional socio-demographic variables: woman's age, number of household members, and a binary variable for large municipality of residence (above 200,000 inhabitants). Moreover, we have included province-level variables on the logarithm of per capita income and the population density.

The estimates for the probability of being victim of physical and non-physical abuse are reported in Tables 7 and 8, respectively. In the first column we report the univariate probit estimates, which ignore the potential endogeneity of employment indicators and household income. The multivariate probit estimates for the equations of the corresponding type of IPV are reported in the second column, and the multivariate probit estimates for the three auxiliary equations of each partner's employment status and low income are reported in the subsequent columns.[7] The marginal effects of the employment status of each partner on each type of IPV based on the corresponding multivariate probit estimations are reported in Table 9.

The univariate probit estimates for physical IPV show a significantly positive coefficient for woman employment and a negative and significant coefficient for the interaction between woman and partner employment, while the coefficient of partner employment is small and non-significant. These results indicate that, conditional on income and other household and characteristics of both partners, in couples in which the woman is employed but her partner is not, the risk of physical IPV is higher than in couples in which neither partner is employed. Nonetheless, the sign of the effect of woman employment is reversed when her partner is also employed, that is, physical abuse risk were reduced only when both partners are employed. Therefore, the mere fact that the woman is employed would not make her less vulnerable to intimate-partner violence, as such effect is contingent to the employment status of his partner. These results resemble previous findings in MacMillan and Gartner (1999) and Terrazas-Carrillo and McWhirter (2015), among others, by which women's employment substantially increases risk of abuse when their male partners are not employed.

Our multivariate probit models, which undertake the endogeneity problem of em-

[7] To obtain our estimates, we use the Stata third-party application mvprobit, written by Capellari and Jenkins (2003).

ployment statuses and low household income, consider, in addition to the IPV equation, three additional reduced form equation for these three variables.[8] The specification of these equations is the usual in the literature.[9] In addition to the joint normal distributional assumption, we consider exclusion restrictions, so that we will include additional variables in the auxiliary equations but not in the IPV equation. Specifically, we use annual province-level female and male employment and unemployment rates by age –that are expected to be good predictors of the employment statuses of each partner– and the percentage of low income households by year, region and size of municipality –that is expected to be a good predictor of the indicator of low income household–. As an exclusion restriction in the employment equations, we assume that employment opportunities, measured by the province-level employment and unemployment rates by gender an age, influence the employment outcome, but have no direct effect on the incidence of domestic violence once the own employment status is accounted for. Likewise, we assume that the fraction of low income households at the local level affect the household poverty risk, but do not have a direct effect on IPV.

Unlike the univariate probit, multivariate estimates for physical IPV show that the estimated coefficient of partner status is significantly negative, whereas female employment status does not increase the risk of physical violence. In fact, the effect of female employment depends critically on the employment status of her partner. Whereas female employment does not reduce the risk of IPV when her partner is not employed, it does when her partner is employed too. Thus, according to these results, partner employment status plays a crucial role: women cohabiting with an employed man face significantly lower risk of experiencing physical abuse than women whose partner is not employed, which is even lower when the woman is employed too.

To compute the marginal effects of employment on physical IPV (first column of Table 9) we have taken as reference the mean values of the variables for 2006. The risk of violence is significantly lower in households where the male partner is employed,

[8] Notice that in this case the potential endogeneity of income arises, not only because of the obvious simultaneity bias, but because of potential measurement error too.

[9] The specification used for the corresponding binary employment equations for each partner resemble the baseline specifications for labor market participation that have been widely used in the empirical labor literature (see for instance Mroz, 1987). One of the major differences between labor force participation specifications for women and men is that in the case of women, a measure of household income (excluding the woman's labor income) is usually included among the covariates.

amounting to 2.9 and 1.9 percentage points less depending on whether the woman is employed or not. Considering female employment, we have already seen that it only has a significantly lowering effect on the risk of physical violence provided that her partner is employed. In this case, the risk of violence is 2.5 percentage points lower. To weigh the importance of these figures, it must be recalled that the sample rate of physical IPV in 2006 amounted to 3.9%.

To better understand the underlying causes of the differences between the univariate and the multivariate estimates, it is useful to look at the estimated correlation coefficients among the error terms in the four equations, reported at the end of Tables 7 and 8 for physical and non-physical violence, respectively. They point out a positive endogeneity bias in the estimated coefficients of the employment statuses of the woman and of her partner. This suggests that unobserved factors that increase the risk of IPV are positively correlated with the employment statuses of each of them. We could think on several potential reasons behind this positive bias, such as that working women are more likely to report physical abuse than non-working women (as pointed out in the literature), or certain unobserved characteristics of each partner that make abuse more likely (such as unobserved family background) or that make them more likely to be employed. Last three columns in Tables 7 and 8 present the estimated coefficients for the two employment and the low income equations, that resemble usual specifications. It is worth noticing that the results for these equations, both for physical and non-physical abuse, are very similar and that the set of instruments used have a highly significant effect in explaining the probability of female employment, male employment and the probability of living in a low income household.

Our findings challenge absolute resource theories by which it is availability of resources, rather than who is the breadwinner, what matters for violence. Besides, the results are not fully supportive of relative resource theories either, by which, other things equal, woman's employment should improve her bargaining position, thus reducing the risk of violence, particularly when her partner's position is weaker (as when he is not employed). The fact that it is male employment what matters most for reducing domestic violence suggests that the significance of male employment status goes beyond its economic role, emphasizing its symbolic role in strengthening the masculinity in the relationship among partners.

These results are consistent with the interpretation that different family patterns, particularly equal vs traditional families, which entail symmetry or asymmetry in labor market participation, are associated with different risk of abuse. Anderson (1997) remarked that status incompatibilities among partners are differentially associated with IPV. Expectations about the relative statuses of male and female partners are based on what is normative within society, and deviations from the norm result in dissatisfaction for the male partner (Horning and McCulloch, 1981). As an extreme case, a situation where the male partner is not employed challenges the cultural depiction of the male partner as breadwinner, what might trigger the risk of violence.

The differential results between the univariate and the multivariate estimated effect of employment estimates are qualitatively similar for physical an non physical IPV. In this latter case, even though employment effects are not individually significant, they remain jointly significant. The marginal effects of employment statuses of each partner on the risk of non-physical IPV (reported in the second column of Table 9) are negative though only significant when the other partner is employed too. Interestingly, the effect of female employment looks substantially higher than the effect of male employment. Specifically, the risk of non-physical violence is 4.9 percentage points lower when the woman is employed, provided that her partner is employed too, whereas such risk is 2.7 percentage points lower when the male partner is employed, provided that the woman is employed too. These effects are quantitatively sizeable, considering that the sample rate of non-physical IPV in 2006 was 8.2%.

As to the effect of income, the other potentially endogenous variable in our model, the positive effect of living in a poor household on both physical and non-physical abuse that is found in the univariate estimations, disappears and becomes statistically non-significant once we account for the endogeneity problem. As with the employment variables, we also find that ignoring the endogeneity problem induces a positive bias, which is validated by the sign and significance of the correlation coefficient between the unobservables affecting the probability of experiencing abuse and the equation of low household income. It is worth noting that the correlation coefficient between the IPV and the low income equations are significantly positive for both types of abuse, suggesting that the unobservables that increase the risk of both types of IPV are positively associated with the unobservables that increase the risk of poverty.

Regarding the effect of the other variables of interest, we find that the risk of experiencing non-physical abuse increases with woman's age, but the effect of woman's age on physical abuse appears too be weaker. Interestingly, the education of the potential abuser seems to matter more than the education of the potential victim: we find that women whose partners are more educated have lower risk of any type of IPV, while the effect of woman's education seems less relevant for physical abuse. We find that differences in education among partners, that capture to some extent socioeconomic disparities among partners, have only a significant effect for non-physical abuse. In particular, if the woman has a higher level of education that her partner, the risk of non-physical abuse increases, while there is not a significant effect on physical abuse. Finally, household size has a significantly positive effect on both types of IPV. This variable provides an imperfect measure of children, who may be seen as a marriage-specific capital, which enhance the value of the relationship and may decrease the chance that the relationship ends (Becker, Landes and Michael, 1977). The theoretical effect of children in the risk of abuse is ambiguous, depending on their relative value to each partner and how it affects the marginal utility for violence of the male partner. The positive effect suggests two channels by which the risk of violence would rise. First, children increased the child care responsabilities for the woman, lowering her reservation utility level; second, children increased the stress within the household, contributing to rise the risk of violence.

6 Conclusions

In this paper, we have addressed the effect of employment statuses of the woman and her partner on the risk of intimate-partner violence (IPV), using cross-sectional data on Spanish women from the 1999, 2002 and 2006 VAW surveys. We have explicitly considered the separate effects of the employment status of each partner, as well as the interaction between them. We have analyzed separately physical and non-physical IPV. Our measures of employment status are two binary indicators that capture the employment status of each partner. Additionally, we have conditioned in income, using a binary variable on whether household income is substantially below average, and a set of sociodemographic characteristics, including woman's age, the levels of education of each partner, and characteristics of the location of residence and the household. Our results may be confronted to different theories about IPV that generate different predictions about the effect of the employment status of each partner.

Our estimation results confirm the need of accounting for endogeneity of employment statuses and income, both for physical and non-physical abuse. Particularly, the employment status of the partner plays a leading role in the risk of physical abuse. Women whose partner is employed are much less likely to experience physical abuse than women cohabiting with non-working partners; this effect is much stronger when the women is employed too. Quantitatively, the estimated marginal effects of the employment statuses of both partners on both types of IPV are sizeable, by which the risk of abuse is less than half among couples in which both partners are employed with respect to couples in which only the male partner works, and much lower with respect to couples in which the male partner does not work. These results seem to be in accordance with the instrumental use of violence by abusing partners predicted by backlash theories, by which violence is exerted when the male partner faces a relative improvement in the woman's position within the household in order to assert his position within the household.

Our results favor the interpretation that different family patterns, particularly equal vs traditional families, which entail symmetry or asymmetry in labor market participation, are associated with different risk of abuse (Anderson, 1997). A major finding is that the risk of physical violence is the lowest for those balanced couples in which both partners are employed. Laws and services devised to empower women and encourage men to value an equal partner are likely to be important steps towards abuse reduction (Coleman and Strauss, 1986). Additionally, as long as more children grow up in nontraditional households with working mothers (Nock, 2001), this will induce an intergenerational transmission of nontraditional gender patterns (Pollak, 2004; Bowlus and Seitz, 2006).

Besides, the risk of violence is strongly increased in those couples where the male partner is not employed, so that his traditional role as main breadwinner is challenged.

Programs designed to increase women's independence should give priority to such vulnerable relationships. Long run policies towards women's economic independence and gender equality must be complemented with short run interventions to increase female empowerment within the household, particularly reinforcing law and deterrence mechanisms and increasing the resources for shelters and help services for battered women. Also, promoting early intervention to target children who witness violence or experience abuse.

Our study is limited by the characteristics of the data. In particular, our data lacks of family background information on violence during partners' childhood. Availability of such information would help us to understand the forces driving female risk of victimization. In addition, it would shed light on the intergenerational effects of family patterns on the transmission of abuse in the long run, given the evidence that domestic violence as a child dramatically increases the odds that men become abusers (Bowlus and Seitz, 2006). Also, it would allow to gauge the extent of assortative mating on the basis of violence in the partners' families of origin (Pollak, 2004).

In addition, the use of cross-sectional data prevents us to understand the dynamics of employment and violence, for which longitudinal data would be needed. In particular, we could exploit past employment statuses to ascertain the long run and short run effects of employment statuses, and the persistence effects of employment on the risk of abuse (see Tauchen and Witte, 1995).

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Table 1
Indicators of IPV violence in Spain

Year	Homicides committed by current or former partner: All	Cohabiting	Without indictment	Number of Indictments*	Users of public help service to abused women	Number of emergency phone calls
1999	54					
2000	63					
2001	50					
2002	54			47,165		
2003	71	55		56,484		
2004	72	48		67,171		
2005	57	36		72,098	2,374	
2006	69	45	36	80,751	5,661	
2007	71	47	40	126,293	8,787	15,715
2008	76	40	45	142,125	12,274	74,951
2009	56	35	33	135,540	13,696	68,541
2010	73	46	41	134,105	8,830	67,696
2011	61	39	40	134,002	9,939	70,679
2012	52	37	38	128,477	9,405	55,810
2013	54	29	30	124,893	10,426	58,268
2014	54	35	30	126,742	10,502	68,612

Source: Spanish Monitoring Centre on Violence Against Women, Ministry of Justice and Ministry of the Interior.
*Until 2006, it only computed direct complaints to the police forces, but not in court.

Table 2
Categories of serious abuse in the Spanish VAW surveys

Behavior	Physical Abuse	Non-Physical Abuse
Stopped from seeing relatives, friends and neighbors		×
Prevented from fair share of household money		×
Insulted or threatened you	×	
Prevented from deciding by yourself		×
Forced to have sexual intercourse	×	
Deprived of your necessities		×
Scared you sometimes		×
Pushed you or hit you	×	
Scorned about your capacity		×
Criticized for the things you do		×
Despised for your beliefs		×
Disregarded for your work		×
Disrespected in front of your children		×

Table 3
Intimate Partner Violence (IPV) by year and type

		IPV Type		
	Obs.	Any	Physical	Non-Phys.
1999	9379	13.40	6.42	10.35
2002	9934	12.39	5.18	9.92
2006	13097	9.96	3.92	8.20
All	32410	11.70	5.03	9.35
Test for independence (χ^2_2)		69.1	71.8	35.3

Source: Own calculations from Spanish VAW Surveys, 1999, 2002, 2006.

Table 4
Intimate Partner Violence (IPV) by region

		IPV Type		
	Obs.	Any	Physical	Non-Phys.
Andalucia	5345	12.95	5.97	10.10
Aragon	1355	8.78	3.03	7.38
Asturias	958	11.27	3.86	9.60
Baleares	886	10.50	4.51	8.24
Canarias	1174	12.44	5.79	9.63
Cantabria	823	9.48	3.89	7.53
Castilla-La Mancha	1299	14.40	6.47	11.39
Castilla y León	1653	12.70	6.05	9.38
Cataluña	3841	10.34	4.17	8.49
Comunitat Valenciana	2609	11.73	4.98	9.08
Extremadura	847	13.93	6.61	10.39
Galicia	2154	11.23	4.83	9.24
Madrid	4868	11.89	4.87	9.98
Murcia	948	12.03	4.85	9.81
Navarra	843	9.96	3.91	7.59
Euskadi	1516	10.95	4.88	8.64
La Rioja	463	9.50	4.10	7.99
Ceuta & Melilla	828	13.29	6.16	10.39
Test for independence (χ^2_{17})		55.1	53.6	34.1

Source: Own calculations from Spanish VAW Surveys, 1999, 2002, 2006.

Table 5
Summary statistics for partners and household characteristics by IPV status

	No Abuse		IPV Type Any		Physical		Non-Phys.	
Woman age								
18-29 years	0.09	(0.29)	0.07§	(0.25)	0.07§	(0.25)	0.07§	(0.25)
30-49 years	0.59	(0.49)	0.53§	(0.50)	0.47§	(0.50)	0.54§	(0.50)
50-64 years	0.32	(0.46)	0.40§	(0.49)	0.46§	(0.50)	0.39§	(0.49)
Woman education								
Primary or less	0.59	(0.49)	0.71§	(0.44)	0.80§	(0.40)	0.73§	(0.44)
Secondary	0.21	(0.41)	0.16§	(0.36)	0.11§	(0.32)	0.15§	(0.35)
College	0.20	(0.40)	0.13§	(0.34)	0.09§	(0.28)	0.12§	(0.32)
Partner age								
18-29 years	0.05	(0.21)	0.03§	(0.17)	0.03§	(0.16)	0.03§	(0.17)
30-49 years	0.56	(0.50)	0.48§	(0.50)	0.43§	(0.50)	0.48§	(0.50)
50-64 years	0.32	(0.47)	0.39§	(0.49)	0.42§	(0.49)	0.39§	(0.49)
above 65 years	0.07	(0.26)	0.10§	(0.30)	0.12§	(0.32)	0.10§	(0.30)
Partner education								
Primary or less	0.59	(0.49)	0.71§	(0.45)	0.77§	(0.42)	0.70§	(0.46)
Secondary	0.21	(0.41)	0.15§	(0.36)	0.14§	(0.34)	0.16§	(0.37)
College	0.19	(0.39)	0.13§	(0.33)	0.09§	(0.29)	0.14§	(0.35)
Household size	3.55	(1.21)	3.69§	(1.34)	3.69§	(1.32)	3.69§	(1.34)
Married (Y/n)	0.95	(0.22)	0.95	(0.22)	0.95	(0.21)	0.94	(0.23)
Household income								
Below average	0.11	(0.31)	0.19§	(0.39)	0.23§	(0.42)	0.19§	(0.39)
Average	0.27	(0.45)	0.30§	(0.46)	0.30†	(0.46)	0.29†	(0.46)
Above average	0.62	(0.49)	0.51§	(0.50)	0.47§	(0.50)	0.52§	(0.50)
Large municipality	0.24	(0.42)	0.24	(0.43)	0.23	(0.42)	0.25*	(0.43)

Source: Own calculations from Spanish VAW Surveys, 1999, 2002, 2006.

Means and standard deviations (between parentheses).

*,†,§denote mean difference with respect to not abused significant at the 10, 5 and 1 percent, respectively.

Table 6
Summary statistics for partners' employment and IPV status

	No Abuse		IPV Type: Any		Physical		Non-Phys.	
Woman empl.	0.42	(0.49)	0.34§	(0.47)	0.31§	(0.46)	0.35§	(0.48)
Partner empl.	0.84	(0.37)	0.78§	(0.42)	0.74§	(0.44)	0.78§	(0.41)
Both employed	0.39	(0.49)	0.30§	(0.46)	0.26§	(0.44)	0.31§	(0.46)
Woman not empl., Partner empl	0.45	(0.50)	0.48§	(0.50)	0.48§	(0.50)	0.47†	(0.50)
Woman empl. Partner not empl.	0.03	(0.17)	0.04§	(0.20)	0.05§	(0.21)	0.04§	(0.20)
Both not empl.	0.13	(0.33)	0.18§	(0.39)	0.21§	(0.41)	0.18§	(0.38)

Source: Own calculations from Spanish VAW Surveys, 1999, 2002, 2006.

Means and standard deviations (between parentheses).

*,†,§denote mean difference with respect to not abused significant at the 10, 5 and 1 percent, respectively.

Table 7
Estimates for risk of Physical IPV

	Univariate	Tetravariate			
	Ipv Physical	Ipv Physical	Woman empl.	Partner empl.	Low income
Woman employed	$0.1359^{\dagger}$	-0.1267			
	(0.0662)	(0.1826)			
Partner employed	-0.0058	$-0.1585^{\dagger}$			
	(0.0375)	(0.0802)			
Woman empl × Partner empl.	$-0.1817^{\dagger}$	$-0.1818^{\dagger}$			
	(0.0715)	(0.0729)			
Household size	$0.0472^{\S}$	$0.0416^{\S}$	$-0.0761^{\S}$	$0.0179^{\dagger}$	$-0.0263^{\S}$
	(0.0097)	(0.0108)	(0.0066)	(0.0085)	(0.0084)
Woman Age 30-49	0.0147	-0.0094	0.0194		
	(0.0482)	(0.0487)	(0.0386)		
Woman Age 50-64	$0.1899^{\S}$	0.0960	-0.0641		
	(0.0513)	(0.0639)	(0.0611)		
Woman: Secondary	$-0.1309^{\S}$	-0.1002^{*}	$0.3804^{\S}$		$-0.3373^{\S}$
	(0.0445)	(0.0515)	(0.0221)		(0.0362)
Woman: College	$-0.1358^{\dagger}$	-0.0511	$0.9975^{\S}$		$-0.5360^{\S}$
	(0.0569)	(0.0811)	(0.0265)		(0.0501)
Partner: Secondary	$-0.1429^{\S}$	$-0.1542^{\S}$	0.0405^{*}	$0.1983^{\S}$	$-0.4602^{\S}$
	(0.0381)	(0.0397)	(0.0225)	(0.0289)	(0.0337)
Partner: College	$-0.2435^{\S}$	$-0.2607^{\S}$	0.0252	$0.4936^{\S}$	$-0.7294^{\S}$
	(0.0517)	(0.0537)	(0.0268)	(0.0316)	(0.0458)
ln(province GDP per capita)	-0.0815	-0.0818	$-0.1051^{\dagger}$	-0.0453	$-0.2459^{\S}$
	(0.0542)	(0.0607)	(0.0491)	(0.0531)	(0.0596)
prov. population density	0.0145	0.0104	-0.0195^{*}	0.0254^{*}	0.0059
	(0.0167)	(0.0168)	(0.0116)	(0.0152)	(0.0141)
Woman more educated	0.0258	0.0214			
	(0.0451)	(0.0447)			
Low income hhold	$0.2566^{\S}$	-0.0477	$-0.3871^{\S}$		
	(0.0339)	(0.1445)	(0.1230)		
Large municipality	0.0428	0.0383			
	(0.0292)	(0.0289)			

Source: Own calculations from Spanish VAW Surveys, 1999, 2002, 2006.

Standard errors in parentheses

*,†, § denote significance at the 10, 5 and 1 percent, respectively.

Table 7 (cont.)
Estimates for risk of Physical IPV

	Univariate	Tetravariate			
	Ipv Physical	Ipv Physical	Woman empl.	Partner empl.	Low income
Y3	$-0.0897^{§}$	$-0.0932^{§}$			
	(0.0248)	(0.0323)			
RNorMed	$-0.0945^{§}$	$-0.1078^{§}$			
	(0.0198)	(0.0260)			
Partner Age 30-49			$-0.1765^{§}$	-0.1193	$-0.1456^{\dagger}$
			(0.0476)	(0.0747)	(0.0584)
Partner Age 50-64			$-0.2322^{§}$	$-0.3088^{§}$	-0.0701
			(0.0536)	(0.0799)	(0.0665)
Partner Age 65+			$-0.3044^{§}$	$-0.5474^{§}$	-0.0927
			(0.0890)	(0.1046)	(0.0969)
Fem. Emp. rate			$1.4676^{§}$	$1.0019^{§}$	$-0.4213^{§}$
			(0.1148)	(0.1092)	(0.1114)
Fem. Unem. rate			-0.4568^{*}	$1.0429^{§}$	0.1227
			(0.2344)	(0.3086)	(0.2755)
Male Emp. rate			$0.3960^{§}$	$2.2619^{§}$	$-1.0646^{§}$
			(0.0903)	(0.0867)	(0.0941)
Male Unem. rate.			0.1810	$1.4885^{§}$	-0.0692
			(0.3109)	(0.4127)	(0.3888)
% Low income					$1.6177^{§}$
					(0.1788)
Wald tests of joint significance ($p-$value)					
Employment ind.	$0.0316^{\dagger}$	$0.0002^{§}$			
Woman Age	$0.0000^{§}$	$0.0403^{\dagger}$	0.0532^{*}		
Woman Educ.	$0.0073^{§}$	0.1037	$0.0000^{§}$		$0.0000^{§}$
Partner Age			$0.0003^{§}$	$0.0000^{§}$	$0.0079^{§}$
Partner Educ.	$0.0000^{§}$	$0.0000^{§}$	0.1963	$0.0000^{§}$	$0.0000^{§}$
Z Em + Unem.			$0.0000^{§}$	$0.0000^{§}$	$0.0000^{§}$
Cross-equations correlation coefficients					
Ipv Physical			0.1652^{*}	0.0736^{*}	0.1417^{*}
			(0.0968)	(0.0391)	(0.0764)
Woman empl.				0.0340^{*}	0.0581
				(0.0176)	(0.0657)
Partner empl.					$-0.3559^{§}$
					(0.0142)

Table 8
Estimates for risk of Non-Physical IPV

	Univariate	Tetravariate			
	Ipv Non-Phys	Ipv Non-Phys	Woman empl.	Partner empl.	Low income
Woman employed	0.0817	−0.1848			
	(0.0574)	(0.1588)			
Partner employed	−0.0276	−0.1252			
	(0.0322)	(0.0773)			
Woman empl × Partner empl.	−0.1164*	−0.0930			
	(0.0612)	(0.0623)			
Hhold size	0.0453§	0.0362§	−0.0761§	0.0178†	−0.0265§
	(0.0080)	(0.0090)	(0.0066)	(0.0085)	(0.0084)
Woman Age 30-49	0.0877†	0.0681*	0.0209		
	(0.0389)	(0.0396)	(0.0386)		
Woman Age 50-64	0.1676§	0.1156†	−0.0595		
	(0.0421)	(0.0539)	(0.0611)		
Woman: Secondary	−0.1250§	−0.1019†	0.3778§		−0.3342§
	(0.0342)	(0.0414)	(0.0221)		(0.0361)
Woman: College	−0.1945§	−0.1192*	0.9943§		−0.5355§
	(0.0432)	(0.0679)	(0.0266)		(0.0501)
Partner: Secondary	−0.0971§	−0.1249§	0.0366	0.1978§	−0.4629§
	(0.0301)	(0.0320)	(0.0224)	(0.0289)	(0.0338)
Partner: College	−0.0562	−0.0977†	0.0201	0.4942§	−0.7275§
	(0.0392)	(0.0417)	(0.0267)	(0.0316)	(0.0457)
ln(province GDP per capita)	0.0143	−0.0111	−0.1095†	−0.0455	−0.2422§
	(0.0443)	(0.0516)	(0.0490)	(0.0532)	(0.0595)
prov. population density	0.0151	0.0120	−0.0190*	0.0256*	0.0045
	(0.0140)	(0.0141)	(0.0115)	(0.0152)	(0.0142)
Woman more educated	0.1342§	0.1229§			
	(0.0355)	(0.0352)			
Low income hhold	0.2190§	−0.2037	−0.4410§		
	(0.0298)	(0.1332)	(0.1194)		
Large municipality	0.0781§	0.0731§			
	(0.0235)	(0.0232)			

Source: Own calculations from Spanish VAW Surveys, 1999, 2002, 2006.

Standard errors in parentheses

*,†, § denote significance at the 10, 5 and 1 percent, respectively.

Table 8 (cont.)
Estimates for risk of Non-Physical IPV

	Univariate	Tetravariate			
	Ipv Non-Phys.	Ipv Non-Phys.	Woman empl.	Partner empl.	Low income
Y3	$-0.0770^{\S}$	$-0.0725^{\S}$			
	(0.0263)	(0.0259)			
RNorMed	$-0.0827^{\S}$	$-0.0845^{\S}$			
	(0.0210)	(0.0211)			
Partner Age 30-49			$-0.1735^{\S}$	-0.1166	$-0.1434^{\dagger}$
			(0.0474)	(0.0747)	(0.0583)
Partner Age 50-64			$-0.2344^{\S}$	$-0.3062^{\S}$	-0.0703
			(0.0536)	(0.0799)	(0.0663)
Partner Age 65+			$-0.3135^{\S}$	$-0.5477^{\S}$	-0.1007
			(0.0890)	(0.1046)	(0.0966)
Fem. Emp. rate			$1.4700^{\S}$	$1.0046^{\S}$	$-0.4346^{\S}$
			(0.1147)	(0.1092)	(0.1111)
Fem. Unem. rate			-0.4544	$1.0497^{\S}$	0.1439
			(0.2342)	(0.3088)	(0.2747)
Male Emp. rate			$0.3670^{\S}$	$2.2572^{\S}$	$-1.0711^{\S}$
			(0.0903)	(0.0867)	(0.0937)
Male Unem. rate.			0.1848	$1.4956^{\S}$	-0.0760
			(0.3108)	(0.4130)	(0.3876)
% Low income					$1.6160^{\S}$
					(0.1782)
Wald tests of joint significance ($p-$value)					
Employment ind.	0.0449^{*}	$0.0030^{\S}$			
Woman Age	$0.0001^{\S}$	0.0991^{*}	$0.0002^{\S}$		
Woman Educ.	$0.0000^{\S}$	$0.0484^{\dagger}$	$0.0000^{\S}$		$0.0000^{\S}$
Partner Age			$0.0002^{\S}$	$0.0000^{\S}$	$0.0002^{\S}$
Partner Educ.	$0.0055^{\S}$	$0.0005^{\S}$	0.2636	$0.0000^{\S}$	$0.0000^{\S}$
Z Em + Unem.			$0.0000^{\S}$	$0.0000^{\S}$	$0.0000^{\S}$
Cross-equations correlation coefficients					
Ipv Non-Phys.			0.1564^{*}	0.0108	$0.2193^{\S}$
			(0.0843)	(0.0366)	(0.0709)
Woman empl.				0.0288	0.0879
				(0.0175)	(0.0641)
Partner empl.					$-0.3566^{\S}$
					(0.0142)

Table 9
Estimated marginal effects of woman and partner employment on IPV

	Physical	Non-Physical
Effect of woman employment		
Partner not employed	-0.0157	-0.0439
	(0.0221)	(0.0270)
Partner employed	-0.0254^{*}	$-0.0491^{\dagger}$
	(0.0140)	(0.0212)
Effect of partner employment		
Woman not employed	-0.0191^{*}	-0.0219
	(0.0102)	(0.0152)
Woman employed	$-0.0289^{\dagger}$	-0.0272^{*}
	(0.0131)	(0.0159)

Source: Own calculations from Spanish VAW Surveys, 1999, 2002, 2006.
Marginal effects evaluated for 2006, using the sample mean values at that year.
Standard errors, computed using the Delta method, in parentheses
*,†, §denote significance at the 10, 5 and 1 percent, respectively.