*The Economic Journal*, **126** (*November*), 1947–1979. Doi: 10.1111/ecoj.12246 © 2015 Royal Economic Society. Published by John Wiley & Sons, 9600 Garsington Road, Oxford OX4 2DQ, UK and 350 Main Street, Malden, MA 02148, USA.

# UNEMPLOYMENT AND DOMESTIC VIOLENCE: THEORY AND EVIDENCE*

*Dan Anderberg, Helmut Rainer, Jonathan Wadsworth and Tanya Wilson*

Does rising unemployment really increase domestic violence as many commentators expect? The contribution of this article is to examine how changes in unemployment affect the incidence of domestic abuse. Theory predicts that male and female unemployment have opposite-signed effects on domestic abuse: an increase in male unemployment decreases the incidence of intimate partner violence, while an increase in female unemployment increases domestic abuse. Combining data on intimate partner violence from the British Crime Survey with locally disaggregated labour market data from the UK's Annual Population Survey, we find strong evidence in support of the theoretical prediction.



During each global recession in recent decades there have been recurrent suggestions in the media that domestic violence increases with unemployment. In 1993, for example the British daily newspaper *The Independent* cited a senior police officer as saying of the increase in domestic violence:

> With the problems in the country and unemployment being as high as it is and the associated financial problems, the pressures within family life are far greater. That must exacerbate the problems and, sadly, the police service is now picking up the pieces of that increase. (Andrew May, Assistant Chief Constable South Wales, *The Independent*, 9 March 1993).

In a 2008 interview for *The Guardian*, the Attorney General for England and Wales argued that domestic violence will spread as the recession deepens:

> When families go through difficulties, if someone loses their job, or they have financial problems, it can escalate stress, and lead to alcohol or drug abuse. Quite often violence can flow from that. (Baroness Scotland of Asthal, *The Guardian*, 20 December 2008).

*Corresponding author: Dan Anderberg, Department of Economics, Horton Building H209, Royal Holloway, University of London, Egham, Surrey, TW20 0EX, UK. Email: dan.anderberg@rhul.ac.uk.

We express our thanks to the Editor, Frederic Vermeulen and two referees for their detailed and helpful comments and suggestions. The article also benefited from comments from seminar participants at University College Dublin, University of Linz, Royal Holloway University of London, IZA, University of Bamberg, CESifo, CEP, ESPE, WPEG, EEA-ESEM, University of Hamburg, ZEW, University of Reading, University of Sussex, University of St Andrews, CASE, RES, NIESR, as well as from Dan Hamermesh and Andy Dickerson. This work was based on data from the British Crime and Annual Population Surveys, produced by the Office for National Statistics (ONS) and supplied under Special Licence by the UK Data Archive. The data are Crown Copyright and reproduced with the permission of the controller of HMSO and Queen's Printer for Scotland. The use of the data in this work does not imply the endorsement of ONS or the UK Data Archive in relation to the interpretation or analysis of the data. This work uses research data sets which may not exactly reproduce National Statistics aggregates.

And in 2012, the executive director of a Washington-based law enforcement think tank expressed his concerns about rising domestic violence rates in a USA *Today* article:

> You are dealing with households in which people have lost jobs or are in fear of losing their jobs. That is an added stress that can push people to the breaking point. (Chuck Wexler, USA *Today*, 29 April 2012).

All these accounts are based on the same underlying logic and suggest that high unemployment may provide the 'trigger point' for violent situations in the home. Yet, from a research perspective, it is far from clear whether unemployment is the overwhelming determinant of domestic violence that many commentators *a priori* expect it to be.[1] Indeed, a basic intra-household bargaining model would suggest that what really matters is the gender-profile of unemployment: an increase in male unemployment and/or a decrease in female unemployment should improve females' relative bargaining power, thereby reducing violence against women in much the same way as a decrease in the gender wage gap (Aizer, 2010).[2]

However, recent empirical evidence also points to factors such as emotional cues (Card and Dahl, 2011) and alcohol (Angelucci, 2008) as potential triggers of partner abuse. Such findings cast doubts on a theory that portrays partner abuse as intentional and rational acts that occur as part of Pareto efficient bargained outcomes. Nevertheless, even in settings where abuse is not an intentionally chosen action, as long as potential abusers have some influence over the likelihood of violent conflicts one would expect the economic logic of the bargaining power argument to carry over.

To verify this, we first present a novel model of partner abuse with gender-specific unemployment risk and where marriage allows couples to diversify income risk partially through consumption sharing. The model, which has parallels to the behavioural framework provided by Card and Dahl (2011), has the innovative feature that a female does not know her husband's 'type' with regard to his predisposition to violence. For a given couple, the male partner may or may not have a violent predisposition and his spouse infers his true nature from his behaviour. In equilibrium, a male with a violent predisposition can either reveal or conceal his type. When a male with a violent predisposition faces a high unemployment risk, he has an incentive to conceal his true nature by mimicking the behaviour of non-violent men. This follows since his spouse, given his low expected future earnings, would have a strong incentive to leave him if she were to learn his violent nature. As a consequence, higher male unemployment is associated with a lower risk of violence. Conversely, when a female faces a high unemployment risk, her low expected future earnings would make her less inclined to leave her partner even if she were to learn that he has a violent nature. Anticipating this, a male with a violent predisposition has

[1] Specifically, we focus on violence against women perpetrated by their partners. While the term 'domestic violence' generally also includes violence between other individuals within households, we will refer to partner violence and domestic violence interchangeably.

[2] The bargaining model provided by Aizer (2010) can be readily extended to include unemployment risk to derive this empirical prediction. An Appendix containing such an extension is available online on the journal website.

weakened incentives to conceal his true nature, thus making higher female unemployment associated with a higher risk of violence. We, therefore, argue that, at a general level, a robust prediction from economic theory is that a high relative male unemployment rate should strengthen the relative economic position of females and lower partner violence against them.

Our empirical analysis combines high-quality individual-level data on intimate partner violence from the British Crime Survey (BCS) with local labour market data at the police force area (PFA) level from the UK's Annual Population Survey (APS). Our basic empirical strategy exploits the substantial variation in the change in unemployment across PFAs, gender and age groups associated with the onset of the late-2000s recession. Our main specification links a woman's risk of being abused to the unemployment rates for females and males in her local area and age group. We first use basic probit regressions to estimate the effects of total and gender-specific unemployment rates on both physical and non-physical abuse. The structure of our data allows us to control for observable socio-economic characteristics at the individual level as well as observable economic, institutional and demographic variables at the PFA level. In addition, we control for unobservable time-invariant area level characteristics and national trends in the incidence of abuse through the inclusion of area and time fixed effects. Finally, as our basic regressions suggest that unemployment matters for the incidence of abuse primarily because of the difference in unemployment rates by gender within areas and age groups, we instrument for the unemployment gender gap by exploiting differential trends in unemployment by industry and variation in initial local industry structure.

We find no evidence to support the common perception that domestic violence increases with the overall unemployment rate. This result parallels findings in previous studies suggesting near zero effects of total unemployment on domestic violence (Iyengar, 2009; Aizer, 2010). However, when we model the incidence of domestic violence as a function of gender-specific unemployment rates, as suggested by economic theory, we find that male and female unemployment have opposite-signed effects on domestic violence: while female unemployment increases the risk of domestic abuse, unemployment among males reduces it. The effects are also quantitatively important: the estimates imply that a 3.7 percentage point increase in male unemployment, as observed in England and Wales over the sample period, 2004–11, causes a decline in the incidence of domestic abuse by up to 12%. Conversely, the 3.0 percentage point increase in female unemployment observed over the same period causes an increase in the incidence of domestic abuse by up to 10%. Thus, our results provide strong support for the predictions arising from the theory. We perform a battery of robustness checks on our data and find that our results are maintained across various alternative specifications. We further note that the relationship between gender-specific unemployment and partner abuse is unique to this type of crime: for the same group of respondents we do not find the same relationship to the personal experience of theft and general violence.

The article contributes to a small but growing literature in economics on domestic violence. These studies can be divided into three broad categories. The first examines the relationship between the relative economic status of women and their exposure to domestic violence. Aizer (2010) specifies and tests a simple model where (some) males

have preferences for violence and partners bargain over the level of abuse and the allocation of consumption in the household.[3] The key prediction of the model is that increasing a woman's relative wage increases her bargaining power and monotonically decreases the level of violence by improving her outside option. Consistent with this prediction, Aizer (2010) presents robust evidence that decreases in the gender wage gap reduce intimate partner violence against women.

The second type of study investigates the effects of public policy on domestic violence. Iyengar (2009) finds that mandatory arrest laws have the perverse effect of increasing intimate partner homicides. She suggests two potential channels for this: decreased reporting by victims and increased reprisal by abusers. Aizer and Dal Bó (2009) find that no-drop policies, which compel prosecutors to continue with prosecution even if a domestic violence victim expresses a desire to drop the charges against the abuser, result in an increase in reporting. Additionally, they find that no-drop policies also result in a decrease in the number of men murdered by intimates suggesting that some women in violent relationships move away from an extreme type of commitment device, that is murdering the abuser, when a less costly one, that is prosecuting the abuser, is offered.

The third type of study focuses more closely on male motives for violence. Card and Dahl (2011) argue that intimate partner violence represents expressive behaviour that is triggered by payoff-irrelevant emotional shocks. They test this hypothesis using data on police reports of family violence on Sundays during the professional football season in the US. Their result suggests that upset losses by the home team (i.e. losses in games that the home team was predicted to win) lead to a significant increase in police reports of at-home male-on-female intimate partner violence. Bloch and Rao (2002) argue that some males use violence to signal their dissatisfaction with their marriage and to extract more transfers from the wife's family. They test their model using data from three villages in India. Pollak (2004) presents a model in which partners' behaviour with respect to domestic violence is transmitted from parents to children.

The remainder of the article is organised as follows. Section 1 establishes the theoretical prediction, linking gender-specific unemployment risks to the incidence of domestic violence against women. Section 2 describes the data that we use. Section 3 outlines the methodology we employ to test the main ideas behind the model and presents the results. Section 4 concludes.

## 1. Theory

The main empirical hypothesis that we will take to the data is that the gender-profile of unemployment should matter for the incidence of domestic abuse. In order to verify the generality of this prediction – which follows naturally from a standard intra-household bargaining model – we present a novel theory of domestic violence in which abuse is not an intentionally chosen action and where asymmetric information occurs. Indeed, our model is the first economic theory to examine domestic violence in a setting where wives do not have perfect information about their husbands' types. The

[3] Earlier studies that have also employed a household bargaining approach to analyse domestic violence include Tauchen *et al.* (1991) and Farmer and Tiefenthaler (1997).

model is based on the premise that marriage is a non-market institution that can provide some degree of insurance against income risk. A key feature of our framework is that a male may or may not have a violent predisposition and that his female partner infers his type from his behaviour. In equilibrium, a male with a violent predisposition can either reveal or conceal his type; his incentives for doing so depend on each partners' future earnings' prospects as determined by their idiosyncratic unemployment risks and potential wages.

### 1.1. *A Signalling Model with Forward-looking Males*

We consider a dynamic game of incomplete information involving two intimate partners: a husband ($h$) and a wife ($w$). The precise timing of the game is as follows:

(*i*) Nature draws a type for the husband from a set of two possible types $\theta \in \{N, V\}$. Type $V$ has a violent predisposition, while type $N$ has an aversion towards violence. The probability that $\theta = V$ is denoted $\phi \in (0,1)$.

(*ii*) The husband learns his type $\theta$ and chooses a behavioural effort from a binary set, $\varepsilon \in \{0,1\}$, which, along with his type, determines the probability that future conflictual interactions with his spouse escalate into violence. The probability of violence occurring is denoted by $\kappa(\theta, \varepsilon) \in [0,1]$. We assume that the behavioural effort $\varepsilon = 1$ reduces the risk of violence and that a husband of type $N$ is less prone to violence than a husband of type $V$. Hence $\kappa(\theta,1) < \kappa(\theta, 0)$ for each $\theta \in \{N,V\}$ and $\kappa(N,\varepsilon) < \kappa(V, \varepsilon)$ for each $\varepsilon \in \{0,1\}$. Making the effort $\varepsilon = 1$ costs the husband $\xi$ (measured in utility units). Effort $\varepsilon = 1$ can, therefore, be interpreted as a costly action for the husband that reduces the likelihood of him 'losing control' in a marital conflict situation. For example, he may voluntarily avoid criminogenic risk factors, such as excessive consumption of alcohol, or he may deliberately reduce his exposure to emotional cues (Card and Dahl, 2011).

(*iii*) The wife observes the husband's action $\varepsilon$ (but not his type $\theta$) and updates her beliefs about his type to $\hat{\phi}(\varepsilon)$. Given her updated beliefs, she then decides whether to remain married or whether to get divorced, a decision we denote by $\chi = \{m,d\}$. If the wife decides to terminate the relationship, each partner $i$ suffers a divorce cost $\alpha_i \geq 0$ (which may be emotional).

(*iv*) Nature decides on employment outcomes. Each partner $i$ ($i = h,w$) is employed or unemployed with probabilities $1 - \pi_i$ and $\pi_i$, respectively. If employed, partner $i$ earns income $y_i = \omega_i$. If unemployed, each individual has an income of $y_i = b$, which can be interpreted as an unemployment benefit.[4] We assume that $b < \omega_i$ for each partner $i$. If still married, each spouse obtains a monetary pay-off that depends on total household income $(y_i + y_j)$. Formally, the monetary pay-off of partner $i$ is

$$u_i^m = \lambda^m + v(y_i + y_j), \tag{1}$$

[4] The benefit income could be gender-specific, but we ignore this for notational simplicity.

where $\lambda^m$ is a constant and where $v(\cdot)$ is an increasing, strictly concave and continuously differentiable function. If divorced, each partner's monetary pay-off depends simply on his or her own income,

$$u_i^d = \lambda^d + v(y_i), \tag{2}$$

where $\lambda^d$ is a constant which satisfies $\lambda^d \geq \lambda^m$. Note that our assumptions on relative pay-offs are consistent with a fairly broad class of non-cooperative models of intra-household public good provision, the equilibria of which feature 'local income pooling' but also inefficiency to due to 'free-riding' (Warr, 1983; Bergstrom *et al.*, 1986; Browning *et al.*, 2010).[5]

(*v*) If still married, the couple encounters a conflict situation (e.g. heated disagreements) which escalates to violence with probability $\kappa(\theta,\varepsilon)$. The wife suffers additive disutility $\delta_w > 0$ if violence occurs. The husband's disutility from violence is type-dependent, $\delta_N > 0$ for a husband of type $N$ and $\delta_V = 0$ for a husband of type $V$.

We solve the model for a pure strategy perfect Bayesian equilibrium. Throughout, $(\varepsilon', \varepsilon'')$ denotes that a husband of type $V$ chooses $\varepsilon'$ and a husband of type $N$ chooses $\varepsilon''$. Similarly, $(\chi', \chi'')$ indicates that the wife plays $\chi'$ following $\varepsilon = 0$ and $\chi''$ following $\varepsilon = 1$.

### 1.2. *Equilibrium*

The wife rationally chooses whether or not to continue the marriage. Her expected pay-off from getting divorced is given by:

$$D(\pi_w) = \mathrm{E}(u_w^d|\pi_w) - \alpha_w, \tag{3}$$

where

$$\mathrm{E}(u_w^d|\pi_w) = \lambda^d + (1 - \pi_w)v(\omega_w) + \pi_w v(b). \tag{4}$$

The expected value to the wife of remaining married depends not only on the wife's own unemployment risk but also on the husband's unemployment probability and the perceived risk of domestic violence. Formally, the wife's expected pay-off from remaining married is given by:

[5] To see this, consider the simplest possible public good game in which the preferences of each individual $i$ are represented by the utility function $u_i(x_i, G) = \ln(x_i) + \ln(G)$, where $G$ is the amount of a 'household good', which is a pure public good to married spouses. Private consumption by individual $i$ is denoted by $x_i$. A married couple's total expenditure on the public good is the sum of individual contributions. These are given by $g_i$, and so $x_i = y_i - g_i$ and $G = g_h + g_w$. Each married individual chooses their contribution to the public good to maximise their utility, taking the contribution of their partner as given. Thus, individual $i$ chooses $g_i$ to maximise $\ln(y_i - g_i) + \ln(g_h + g_w)$. It is straightforward to show that both spouses will contribute to the household good for income shares satisfying $y_h/(y_h + y_w) \in (1/3, 2/3)$. In this case, $x_i^* = G^* = (y_h + y_w)/3$, that is there is 'local income pooling' with household demands for all goods only depending on aggregate household income and not on individual incomes. Thus, the indirect utility function of individual $i$ can be written as $u_i^m(x_i^*, G^*) = \lambda^m + v(y_h + y_w)$, where $\lambda^m = -2\ln(3)$ and $v(\cdot) = 2\ln(\cdot)$. A divorced individual chooses $g_i$ to maximise $\ln(y_i - g_i) + \ln(g_i)$, and so the indirect utility function of a divorced individual can be written as $u_i^d(x_i^*, g_i^*) = \lambda^d + v(y_i)$, where $\lambda^d = -2\ln(2)$ and $v(\cdot) = 2\ln(\cdot)$. Notice that this simple example implies our assumption that $\lambda^d > \lambda^m$, which follows from the fact that married spouses 'crowd out' each other's contributions to the public good, that is that their contributions are strategic substitutes.

$$M[\pi_h, \pi_w, \varepsilon, \hat{\phi}(\varepsilon)] = \mathrm{E}[u_w^m|(\pi_h, \pi_w)] - \delta_w\{[1 - \hat{\phi}(\varepsilon)]\kappa(N, \varepsilon) + \hat{\phi}(\varepsilon)\kappa(V, \varepsilon)\}, \quad (5)$$

where

$$\begin{aligned}\mathrm{E}[u_w^m|(\pi_h, \pi_w)] = \lambda^m + (1 - \pi_h)(1 - \pi_w)v(\omega_w + \omega_h) + \pi_h\pi_w v(2b) \\ + \pi_h(1 - \pi_w)v(\omega_w + b) + \pi_w(1 - \pi_h)v(b + \omega_h).\end{aligned} \quad (6)$$

Note that the wife's expected utility from remaining married is decreasing in her perceived probability that the husband has a violent predisposition, $\hat{\phi}(\varepsilon)$. The wife continues the partnership if and only if her expected value of remaining married exceeds the expected value of getting divorced. The key assumptions of the model are as follows (for expositional convenience, we suppress the arguments of the functions):

Assumption 1. *$M < D$ when $\pi_w = 0$, $\pi_h = 1$, $\varepsilon = 0$ and $\hat{\phi} = 1$.*
Assumption 2. *$M > D$ when $\pi_w = 1$, $\pi_h = 0$, $\varepsilon = 0$ and $\hat{\phi} = 1$.*
Assumption 3. *For any $(\pi_h, \pi_w) \in [0, 1]^2$ and $\varepsilon \in \{0,1\}$, $M > D$ when $\hat{\phi} = \phi$.*

The first two assumptions imply that the wife's tolerance of violence depends on her earnings prospects. To be more precise, suppose the wife observes the husband choosing $\varepsilon = 0$. Assumption 1 ('not-take-it-if-employed') then says that if the wife will be employed with certainty and the husband will be unemployed with certainty, and she knows that the husband has a violent predisposition, then she will choose to divorce the husband. This may be interpreted as implying that economically independent women leave their abusive partners. On the other hand, Assumption 2 ('accept-it-if-unemployed') implies that if the wife will be unemployed with certainty and the husband will be employed with certainty, and she knows that he has a violent predisposition, then she will not leave him. This captures the idea that women who are economically dependent on their abusers may be unable to leave them. Finally, Assumption 3 ('stay-if-no-new-info') says that if the wife retains her prior beliefs, then she will continue the relationship irrespective of their unemployment probabilities and the husband's action. It is, therefore, consistent with wife accepting to be in a partnership with the husband in the first place.

In addition, we make the following two-part assumption:

Assumption 4.

(*i*) $[\kappa(N, 0) - \kappa(N, 1)]\delta_N > \xi$, *and*

(*ii*) $\alpha_h > \kappa(N, 0)\delta_N$.

Part (*i*) implies that a husband with an aversion towards violence values the reduction in violence associated with making the effort $\varepsilon = 1$ more than its cost. Part (*ii*) is a sufficient condition to ensure that continued marriage is preferable to divorce for each type of husband $\theta \in \{N,V\}$ at any effort level $\varepsilon \in \{0,1\}$. Thus, the husband has no incentive to choose his behavioural effort in a way that triggers a divorce.

Next we define $\hat{\pi}_w(\pi_h)$ as the unemployment probability for the wife at which she, conditional on having observed the husband choosing $\varepsilon = 0$ and knowing that the husband has a violent predisposition, is indifferent between continued marriage and divorce. Formally, $\hat{\pi}_w(\pi_h)$ is implicitly defined through:

$$M[\pi_h, \hat{\pi}_w(\pi_h), 0, 1] = D[\hat{\pi}_w(\pi_h)]. \tag{7}$$

Equation (7) may fail to have a solution in the unit interval. However, the following Lemma tells us that it will do so for some values of $\pi_h$.

LEMMA 1. *There exist two values, $\pi'_h$ and $\pi''_h$, satisfying $0 \leq \pi'_h < \pi''_h \leq 1$ such that (7) has a solution $\hat{\pi}_w(\pi_h) \in [0,1]$ for every $\pi_h \in [\pi'_h, \pi''_h]$. Moreover, $\hat{\pi}_w(\pi_h)$ is differentiable at any $\pi_h \in (\pi'_h, \pi''_h)$ with $\partial\hat{\pi}_w(\pi_h)/\partial\pi_h > 0$. In addition, $\partial\hat{\pi}_w(\pi_h)/\partial\omega_w > 0$ and $\partial\hat{\pi}_w(\pi_h)/\partial\omega_h < 0$.*

*Proof.* See Appendix A.

Figure 1 illustrates a case where $\pi' > 0$ and $\pi''_h < 1$. The locus $\hat{\pi}_w(\pi_h)$ partitions the set of possible unemployment risk profiles, $(\pi_h, \pi_w) \in [0,1]^2$, into two non-empty subsets or 'regimes':

$$R_0 \equiv \{(\pi_h, \pi_w) | \pi_h \geq \pi''_h\} \cup \{(\pi_h, \pi_w) | \pi_w \leqslant \hat{\pi}_w(\pi_h)\}, \tag{8}$$

$$R_1 \equiv \{(\pi_h, \pi_w) | \pi_h < \pi'_h\} \cup \{(\pi_h, \pi_w) | \pi_w > \hat{\pi}_w(\pi_h)\}. \tag{9}$$

An increase in the husband's wage $\omega_h$ expands regime $R_1$ by shifting the locus $\hat{\pi}_w(\pi_h)$ downwards. In contrast, an increase in the wife's wage $\omega_w$ expands regime $R_0$ by shifting the locus upwards.

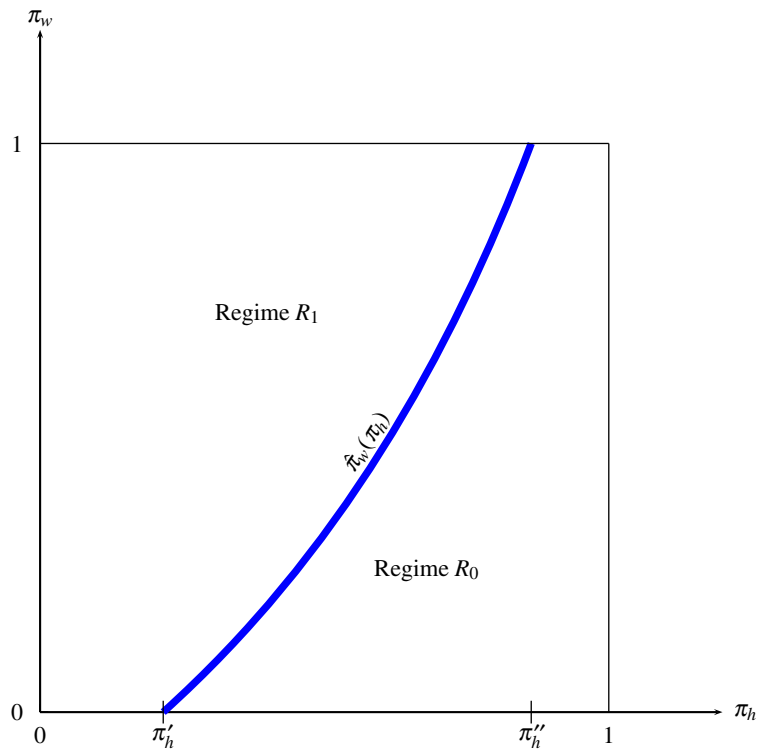


Fig. 1. *Critical Locus $\hat{\pi}_w(\pi_h)$ Separating Regime $R_1$ and Regime $R_0$*

The following Proposition shows that the nature of the game's equilibrium depends on which regime the couple's unemployment risk profile $(\pi_h, \pi_w)$ falls within. Since signalling games are prone to equilibrium multiplicity, we focus on pure strategy equilibria that satisfy the commonly used Cho–Kreps 'intuitive criterion' (Cho and Kreps, 1987).

PROPOSITION 1. *In each regime, there is a unique pure strategy perfect Bayesian equilibrium that satisfies the 'intuitive criterion'*

(*a*) *If* $(\pi_h, \pi_w) \in R_0$, *then*

$$[(\varepsilon', \varepsilon'') = (1, 1), (\chi', \chi'') = (d, m), \hat{\phi}(0) = 1, \hat{\phi}(1) = \phi]$$

*is a 'pooling' equilibrium.*

(*b*) *If* $(\pi_h, \pi_w) \in R_1$, *then*

$$[(\varepsilon', \varepsilon'') = (0, 1), (\chi', \chi'') = (m, m), \hat{\phi}(0) = 1, \hat{\phi}(1) = 0]$$

*is a 'separating' equilibrium.*

*Proof.* See Appendix A.

To see that this describes a perfect Bayesian equilibrium, consider each regime in turn, starting with $R_0$. Here a pooling equilibrium occurs where both types of husbands make the costly effort that reduces the risk of violence. A husband without a violent predisposition makes the effort since he values the reduction in the risk of violence that it generates more than the cost. A husband with a violent predisposition on the contrary makes the effort in order not to reveal his type as doing so would trigger a divorce. Central to the equilibrium are the wife's out-of-equilibrium beliefs and associated action: upon observing $\varepsilon = 0$, the wife would conclude that the husband has a violent predisposition and would choose divorce.

Consider then regime $R_1$. In this case the husband knows that the wife is economically vulnerable and would not leave him even if she were to believe that he has a violent predisposition. A husband with a violent predisposition therefore has no incentives to make the costly effort that would reduce the risk of violence. A husband without a violent predisposition again values the reduction in the risk of violence more than the cost of making the effort. The wife's belief updating follows Bayes' rule and her continuing of the partnership with either type of husband is rational given her relatively weak earnings prospects.

### 1.3. *Empirical Prediction*

The above results form the basis of our empirical predictions: men with a violent predisposition may strategically mimic the behaviour of non-violent men, thus concealing their type, when facing relatively weak earnings prospects (Regime $R_0$) in the form of relatively high unemployment risk and relatively low wages. In contrast, when men face relatively strong earnings prospects (Regime $R_1$) they will be less inclined to conceal any violent predisposition they may have. Noting that the

difference in the equilibrium probability of violence between Regime $R_1$ and $R_0$ is $\phi[\kappa(V,0) - \kappa(V,1)] > 0$ we arrive at the following central empirical prediction:

PREDICTION 1.

- *A higher risk of male unemployment and lower wages for men are associated with a lower risk of domestic violence.*
- *A higher risk of female unemployment and lower wages for women are associated with a higher risk of domestic violence.*

Thus, we build our empirical approach on the theoretical prediction that a woman's risk of being abused depends on gender-specific unemployment risks. In particular, in the empirical analysis we relate a woman's risk of experiencing domestic abuse to the local unemployment rates for males and females in her own age group.[6]

## 2. Data and Descriptive Statistics

### 2.1. *Domestic Abuse Data from the BCS*

We use data on the incidence of domestic abuse from the BCS. The BCS is a nationally representative repeated cross-sectional survey of people aged 16 and over, living in England and Wales, which asks the respondents about their attitudes towards and experiences of crime. The BCS employs two different methods of data collection with respect to domestic abuse. The first method, available from the survey's inception in 1981, is based on face-to-face interviews. However, the unwillingness of respondents to reveal instances of abuse to interviewers implies that this method significantly underestimates the true extent of domestic violence. To overcome such non-disclosure, a self-completion module on interpersonal violence (IPV), which the respondents complete in private by answering questions on a laptop, was introduced.[7] We use BCS data for the survey years 2004/5–2010/11, covering interviews conducted between April 2004 and March 2011, and base our analysis on data on domestic violence from the self-completion IPV module.[8]

The BCS data have several key strengths as a source of data on domestic abuse. The IPV module in particular is unique in an international context, insofar that through self-completion the respondent does not need to provide answers directly to an interviewer. Furthermore, to reassure the respondent of privacy, the BCS randomly selects only one person per household who is exposed to the survey only once,

[6] On a related note, Raphael and Winter-Ebmer (2001) find a negative association between female unemployment and state level rape rates which they suggest may be attributable to an increased number of encounters with potential perpetrators if an individual is working away from home. Rape perpetrated by a non-partner is, however, by definition not an outcome decided within a relationship and, therefore, is not applicable to our theoretical framework.

[7] The IPV module was first introduced in 1996. In 2001 it was used for a second time and the use of laptops was introduced. Since the 2004/5 survey the IPV module has been included on an annual basis, with a comparable set of questions.

[8] In the 2010–11 BCS survey, half of the sample were, in a trial, asked the same abuse questions but in a simplified sequential format. For consistency, we include in our sample only those respondents who were asked the abuse questions in the format consistent with the previous years' surveys.

implying that other household members including any partner will not know what questions the respondent has faced. In contrast the corresponding US survey, the National Crime Victimization Survey, administers the same set of questions to all household members every six months over a three year period, implying that the content of the questionnaire is common knowledge within the household.

Over our sample period, only 11% of those who report, in the IPV module, having been subjected to physical abuse by a partner also report being exposed to intra-household abuse in the general interviewer-based part of the BCS survey. Similarly, only 48% and 50% report having mentioned the abuse to a medical staff and to the police respectively. Hence, compared to alternative data from interviewer-based surveys, or data derived from police reports or hospital episodes statistics, the BCS IPV data are likely to provide substantially more comprehensive data on the incidence of domestic abuse. Furthermore, while police reports and hospital episode data can be used to measure incidence of (severe) domestic violence, such data generally cannot distinguish between multiple victims *versus* multiple events for the same victim. Finally, using micro-level data, obviously, allows us to control for individual level characteristics.

The BCS IPV module is answered by respondents aged 16–59 and we focus our analysis on intimate partner violence experienced by women.[9] Table 1 presents descriptive statistics of our sample.

In the IPV module, respondents are presented with a list of types of behaviour that constitute domestic abuse and are asked to indicate which, if any, they have experienced in the 12 months prior to the interview. Table 2 presents this list of behaviour types from which we, following the Home Office classification, construct two

Table 1

*Demographic Characteristics of the BCS Sample*

| Variable | Mean | SD | Variable | Mean | SD |
|---|---|---|---|---|---|
| Age | 38.93 | 11.67 | Qualification: Degree or above | 0.236 | 0.425 |
| Ethnicity: White | 0.928 | 0.258 | Qualification: High Education < Degree | 0.137 | 0.344 |
| Ethnicity: Asian | 0.028 | 0.165 | Qualification: A level | 0.150 | 0.357 |
| Ethnicity: Black | 0.023 | 0.150 | Qualification: GCSE grades A–C | 0.237 | 0.426 |
| Ethnicity: Other | 0.021 | 0.143 | Qualification: Other | 0.096 | 0.295 |
| Religion: None | 0.216 | 0.412 | Qualification: None | 0.143 | 0.350 |
| Religion: Christian | 0.740 | 0.439 | Single | 0.355 | 0.479 |
| Religion: Muslim | 0.017 | 0.128 | Married | 0.455 | 0.498 |
| Religion: Hindu | 0.009 | 0.092 | Separated | 0.046 | 0.209 |
| Religion: Sikh | 0.004 | 0.060 | Divorced | 0.125 | 0.331 |
| Religion: Jewish | 0.003 | 0.057 | Widowed | 0.019 | 0.136 |
| Religion: Buddhist | 0.005 | 0.069 | Cohabiting | 0.120 | 0.325 |
| Religion: Other | 0.008 | 0.087 | Long-standing illness | 0.179 | 0.383 |
| Number of children | 0.493 | 0.896 | Poor health | 0.031 | 0.174 |
| Children below five years | 0.110 | 0.313 | | | |
| Observations | | | 86,898 | | |

[9] While the IPV module is also completed by male respondents, abuse against men is less common, generally less violent, and with no apparent connection to labour market conditions.

Table 2

*Categories of Domestic Abuse*

| Behaviour | Physical abuse | Non-physical abuse |
|---|---|---|
| Prevented from fair share of household money | | × |
| Stopped from seeing friends and relatives | | × |
| Repeatedly belittled you | | × |
| Frightened you, by threatening to hurt you | | × |
| Pushed you, held you down or slapped you | × | |
| Kicked, bit or hit you | × | |
| Choked or tried to strangle you | × | |
| Threatened you with a weapon | × | |
| Threatened to kill you | × | |
| Used a weapon against you | × | |
| Used other force against you | × | |

binary indicators of abuse. The first, physical abuse, is a dummy variable indicating whether the respondent had any type of physical force used against them by a current or former intimate partner. The second, non-physical abuse, indicates whether the respondent was threatened, exposed to controlling behaviours or deprived of the means needed for independence by a current or former partner.

In our sample, 3.0% of women report episodes of physical abuse in the past 12 months and 4.4% declare having experienced non-physical abuse.[10] Figure 2 illustrates the extent to which the incidence of physical abuse in particular varies with the demographic characteristics of the respondents. In general, exposure to physical abuse declines with age and with academic qualifications acquired after compulsory education. It varies relatively little with religion and ethnicity, but increases with the number of children.[11] With respect to marital status, it should be noted that this refers to the respondent's formal status at the time of the interview, which is hence observed after the 12-month period to which the abuse questions refer. The high reported rate of abuse among separated and divorce women therefore suggests a 'reverse causality'. The high rate of incidence among singles also emphasises the fact that 'intimate partners' include current and past boyfriends.[12] Due to the highly endogenous nature of the respondent's current marital status, we do not make use of this information except as a final sensitivity check on our estimates.[13] Figure 3 shows the trends in physical and non-physical abuse which, if anything, suggests that the overall level of abuse is lower towards the end of our sample period than at the beginning. A corresponding decline has been observed over the same period in the most extreme form of violence against women: the rate of female homicide where the prime suspect is an intimate partner decreased by 6.3% between 2003–7 and 2007–11 (Smith *et al.*, 2012).

[10] The fraction of women reporting at least one of the two types of abuse was 5.7%.

[11] The relationship between physical violence and ethnicity is somewhat unexpected, given the data from the US where blacks are typically found to have a higher incidence (Aizer, 2010).

[12] For respondents who are not currently married we also use a cohabitation dummy to indicate that the respondent is currently living with a partner. The incidence of abuse among currently cohabiting respondents is about double that of currently married respondents.

[13] The same applies to any information we have on the individual's current employment status. Hence, we make no use of such information.

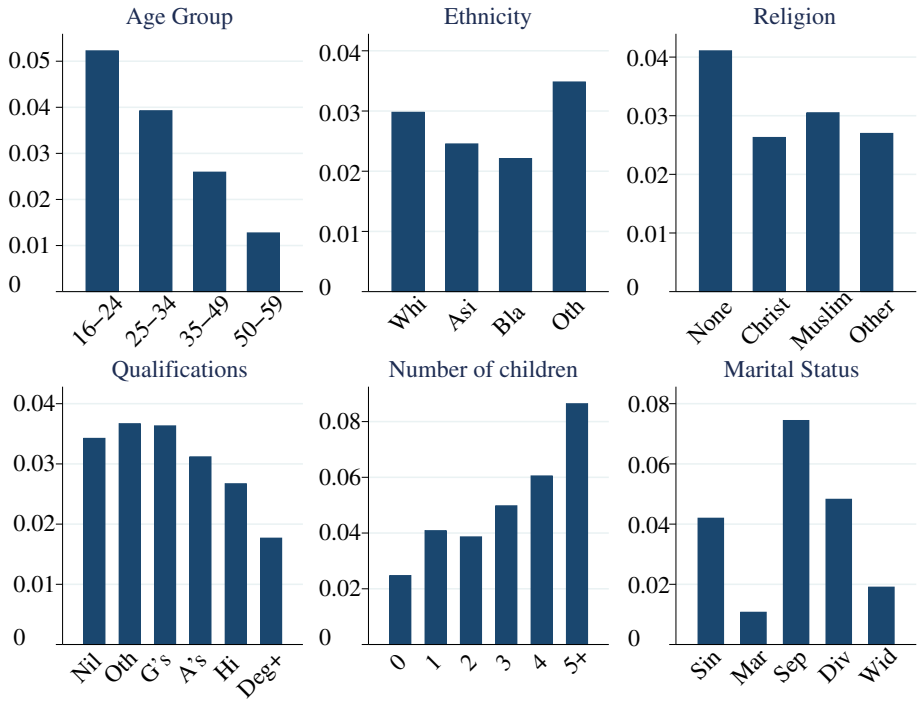


Fig. 2. *Incidence of Physical Abuse by Demographic Characteristics*

### 2.2. *Labour Market Data from the APS*

We merge our individual-level data from the BCS with labour market data from the APS. The APS combines the UK Labour Force Survey (LFS) with the English, Welsh and Scottish LFS boosts. Data sets are released quarterly, with each data set containing 12 months of data. This means that we can, for each respondent in the BCS, using the known interview date, match the 12-month period to which the IPV questions refer to a closely corresponding 12-month period of labour market data.[14] Each respondent is matched to local labour market conditions corresponding to the PFA of residence, of which there are 42 in our data.[15]

[14] For instance, any respondent interviewed in the first three months of 2005 is matched to the labour market data for the calendar year 2004, whereas a BCS respondent interviewed between April and June in 2005 is matched to labour market data for the period April 2004 to March 2005 etc.

[15] There are 43 PFAs in England and Wales. However, the City of London PFA is a small police force which covers the 'Square Mile' of the City of London. As this is a small area enclosed in the many times larger Metropolitan PFA we merge the two. This leaves us with 42 PFAs. They are Avon and Somerset, Bedfordshire, Cambridgeshire, Cheshire, Cleveland, Cumbria, Derbyshire, Devon and Cornwall, Dorset, Durham, Essex, Gloucestershire, Greater Manchester, Hampshire, Hertfordshire, Humberside, Kent, Lancashire, Leicestershire, Lincolnshire, City of London and Metropolitan Police District, Merseyside, Norfolk, Northamptonshire, Northumbria, North Yorkshire, Nottinghamshire, South Yorkshire, Staffordshire, Suffolk, Surrey, Sussex, Thames Valley, Warwickshire, West Mercia, West Midlands, West Yorkshire, Wiltshire, Dyfed-Powys, Gwent, North Wales and South Wales. The APS data are available in a finer geography and are hence aggregated up to the PFA level.

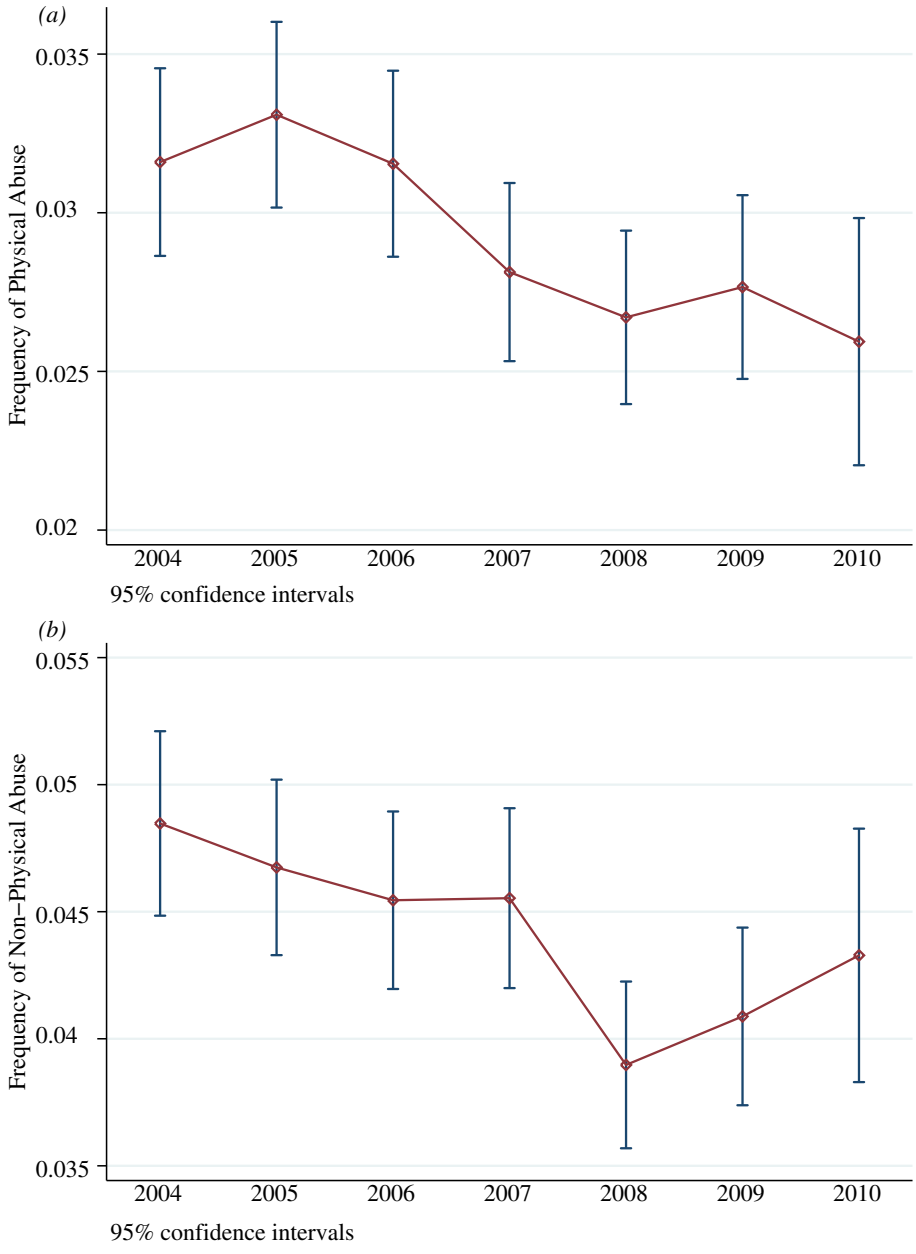


Fig. 3. *Trends in Domestic Abuse in England and Wales*

Our theory developed in the previous Section stresses the role of male and female unemployment risk for the incidence of domestic violence. In the empirical analysis, we relate the incidence of domestic violence to the observed unemployment rates for the respondent's female and male peers, as defined by age group and geographical area. Hence, we effectively interpret the observed unemployment rate not only as a

measure of the direct incidence of unemployment but also more broadly as an indicator for the perceived risk of unemployment. This interpretation is supported by the literature that documents workers' subjective unemployment expectations and relates it to the current level of unemployment. For instance for the US, Schmidt (1999) shows how workers' average beliefs about the likelihood of job loss in the next 12 months closely tracked the unemployment rate over the period 1977–96. The limited data that are available on unemployment expectations in the UK equally support the notion that individual expectations of future unemployment risk are positively associated with the current unemployment rate. The British Social Attitudes (BSA) survey has, in selected years, asked respondents:

(*i*) how 'secure' they feel in their jobs; and
(*ii*) whether they expect to see a change in the number of employees in their workplace.

Both variables saw changes with the onset of the latest recession. In 2005, 78% of respondents reported feeling secure in their jobs; in 2009–10, this figure had dropped to 73%. Similarly, while 16% of respondents reported expecting a reduction in the number of employees in the workplace in 2006–7, this number had increased to 26% in 2009–10.[16]

Table 3 presents basic descriptive statistics for local unemployment rates, broken down by gender and age group.[17] Figure 4 shows that the increase in the rate of unemployment (left-hand scale) associated with the latest recession was far from uniform across gender and age groups. In particular, the impact of the recession is reflected more strongly in male than in female unemployment. As a consequence, we observe a widening of the female–male unemployment gap (right-hand scale) in the

Table 3
*Summary Statistics for Local Unemployment Rates*

| Variable | Mean | SD | Min | Max |
|---|---|---|---|---|
| Total unemployment | 0.060 | 0.020 | 0.022 | 0.129 |
| *Unemployment by gender* | | | | |
| Male | 0.064 | 0.023 | 0.022 | 0.149 |
| Female | 0.054 | 0.018 | 0.014 | 0.103 |
| *Unemployment by age group* | | | | |
| Aged 16–24 | 0.150 | 0.045 | 0.029 | 0.283 |
| Aged 25–34 | 0.055 | 0.021 | 0.009 | 0.136 |
| Aged 35–49 | 0.039 | 0.016 | 0.010 | 0.104 |
| Aged 50–64 | 0.035 | 0.014 | 0.004 | 0.086 |

*Notes.* The Table provides averages over the time-interval January 2003–December 2010 based on data from the APS which are provided in overlapping 12-month periods: January–December, April–March, July–June, October–September. Reported standard deviations and minimum and maximum values are over 1,218 PFA-period observations.

[16] Using data from the Skills Surveys, Campbell *et al.* (2007) document a similar fall in the average individual expectations of job loss between 1997 and 2001, a period of declining unemployment.
[17] The age grouping used in our analysis follows that conventionally used by the Office for National Statistics.

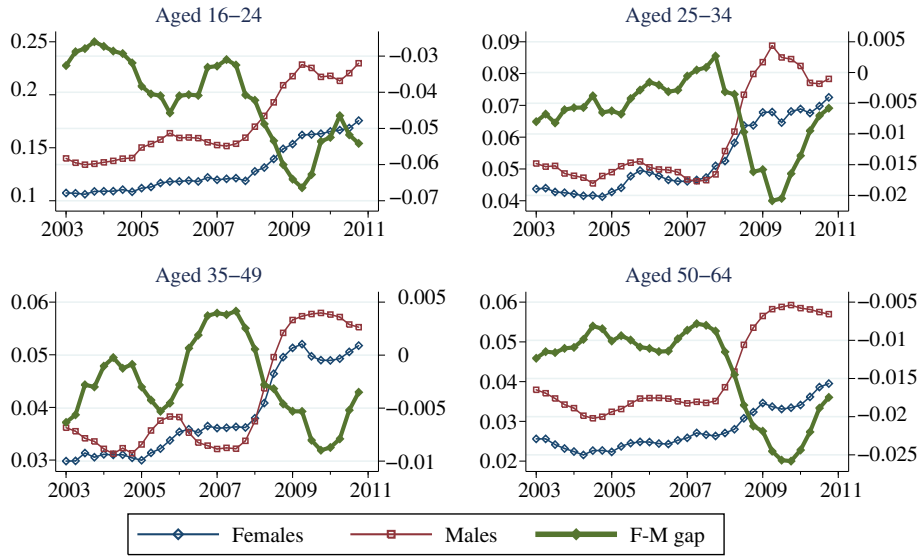


Fig. 4. *Gender-specific Unemployment Rates and the Female–Male Unemployment Gap by Age Group in England and Wales, 2003–2011*

latter part of the sample period. In addition to local unemployment, we also use the APS to construct measures of mean hourly real wages.

Figure 5 contrasts the change over the sample period from 2004/5 to 2010/11 in the incidence of physical abuse with corresponding changes in male and female unemployment rates across the 42 PFAs. The Figure highlights substantial spatial variation in the change in unemployment over the sample period. Moreover, the local changes in female unemployment are not obviously correlated with the corresponding local changes in male unemployment.[18] Inspection of the Figure further suggests that several PFAs in which men were relatively more affected by unemployment increases (e.g. the North–East) saw relative decreases in the incidence of physical abuse. Indeed, if anything, the Figure suggests a more positive association between relative increases in female unemployment and relative increases in abuse.

## 3. Empirical Specification and Results

### 3.1. *Baseline Specification*

This subsection presents our main analysis where we relate a female respondent's experience of domestic violence to the local level of unemployment. We focus in particular on the rates of female and male unemployment within the respondent's own age group as these are likely to be the most relevant for the respondent's own unemployment risk as well as that of her (potential) partners. As the APS data are released

[18] Indeed, relevant to our identification strategy, less than 40% of the variation in the gender unemployment gap is explained by area, period and age group effects. This figure increases to just under 50% when age-group-period and age-group-area fixed effects are introduced.

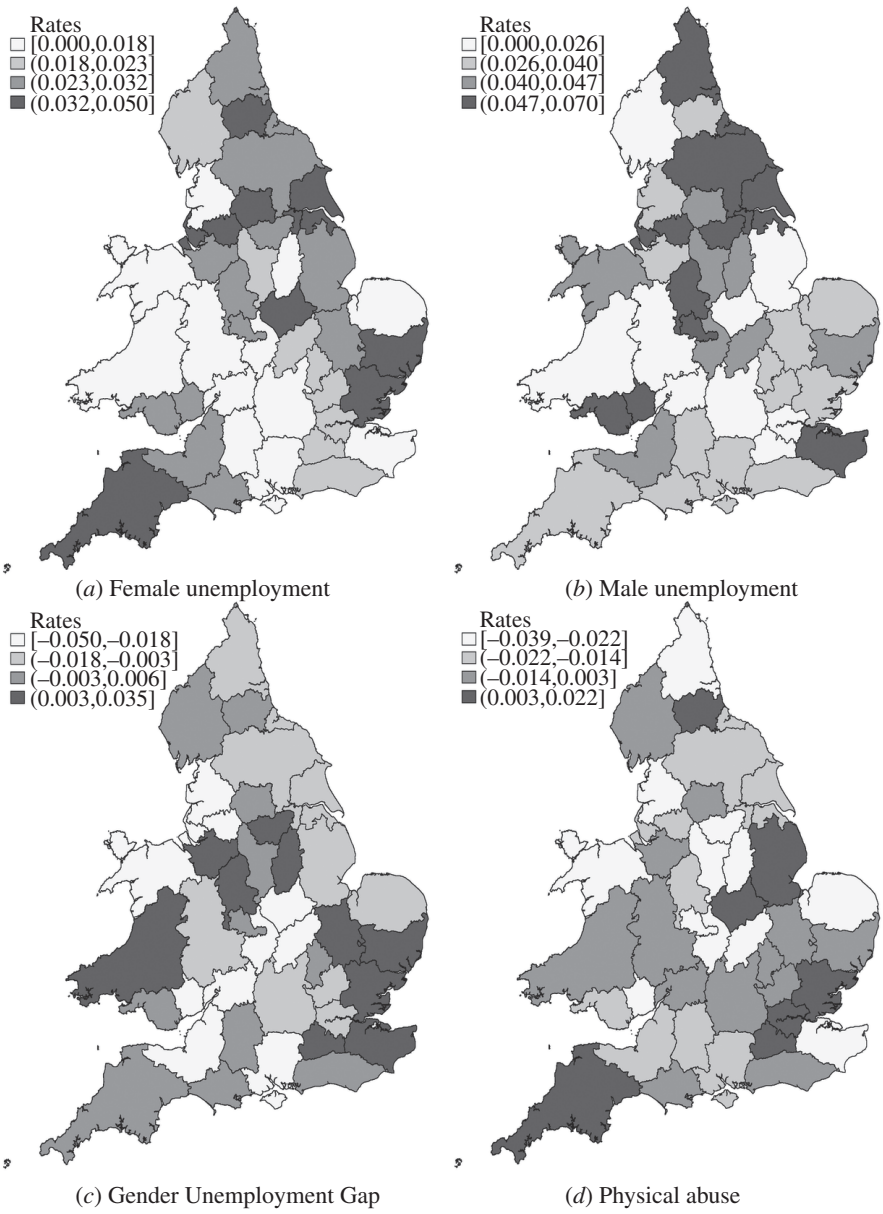


Fig. 5. *Change in Female and Male Unemployment, in the Female–Male Unemployment Gap and in Incidence of Physical Abuse across Police Force Areas in England and Wales, 2004–2011*

quarterly, with each data set containing 12 months of data, we define a 'period' variable, denoted $t$, where a given period contains the particular APS release and BCS data from the following three months. Constructed in this way, our data stretch over 28 periods.[19]

[19] See footnote 14 for further details.

As the outcome variables in our analysis are binary indicators of abuse, we estimate probit models. In particular, the basic model for the latent propensity for abuse against individual $i$ in PFA $j$ in period $t$ and within age group $g$ is given by

$$y^*_{ijtg} = \beta X_{ijtg} + \gamma^f UNEMPL^f_{jtg} + \gamma^m UNEMPL^m_{jtg} + \lambda_t + \alpha_j + \varepsilon_{ijtg} \quad (10)$$

where $X_{ijtg}$ includes demographic controls at the individual level, $UNEMPL^f_{jtg}$ and $UNEMPL^m_{jtg}$ are the female and male unemployment rate in $i$'s own age group in PFA $j$ during period $t$, and $\varepsilon_{ijtg}$ is a normally distributed random term.[20] The parameters $\lambda_t$ and $\alpha_j$ are fixed effects for time-periods and PFAs respectively, and thus control for the aggregate trend in the outcome variable and for factors affecting abuse that vary across areas but are fixed over time. Thus, our basic model identifies the impact of gender-specific unemployment on domestic abuse from variation in trends across PFAs.

### 3.2. *Baseline Results*

Our basic results for the probability of being a victim of physical abuse are provided in Table 4.[21] Specification (1) gives the average marginal effect of the total unemployment rate within the own age group on the incidence of physical abuse. The estimated model includes a set of individual-level demographic controls: age measured in years and dummy variables indicating ethnicity, qualification level and religious denomination, along with number of children and a dummy for the presence of at least one child under the age of five in the household. It further includes area and time fixed-effects. The marginal effect is small and insignificant.[22] This result parallels findings in previous studies (Iyengar, 2009; Aizer, 2010) suggesting near zero effects of total unemployment on domestic violence.

Specification (2) reports the estimated average marginal effect of each gender-specific unemployment rate within the own age group. The marginal effect of female unemployment in the own age group is positive and statistically significant. The magnitude of the coefficient suggests that a 1 percentage point increase in the own age female unemployment rate causes an increase in the likelihood of the respondent being a victim of physical abuse by 0.097 percentage points or little over 3% of the sample mean. We also see that the estimated average marginal effect of male unemployment is negative and statistically significant. The magnitude of the coefficient indicates that a 1 percentage point increase in male unemployment in the respondent's own age group causes a decline in the risk of physical abuse by 0.090 percentage points – again about 3% of the sample mean.

Controls for female and male unemployment within age groups other than the own are added in specification (3). We find that male and female unemployment within the own age group still have opposite-signed effects on the risk of physical abuse, while unemployment in age groups other than the own appears to have little impact. Our

[20] In subsection 3.3, we further include area-level controls.

[21] Estimates from linear probability models are very similar and are available on request from the corresponding author.

[22] A (non-reported) regression on aggregate unemployment – across genders and age groups – is also not significant but also has less precision due to low local variation from the national trend.

Table 4

*Impact of Unemployment on Physical Abuse – Main Specification*

| Specification | (1) | (2) | (3) | (4) | (5) | (6) | (7) | (8) | (9) |
|---|---|---|---|---|---|---|---|---|---|
| Unemployment in own age group | −0.026 | | | | | | | | 0.008 |
| | (0.018) | | | | | | | | (0.019) |
| Female unemployment in own age group | | 0.097** | 0.093** | 0.102** | 0.094** | 0.083** | 0.102** | 0.091** | |
| | | (0.027) | (0.027) | (0.028) | (0.027) | (0.029) | (0.028) | (0.035) | |
| Male unemployment in own age group | | -0.090** | −0.097** | −0.081** | −0.089** | −0.094** | −0.067* | −0.084* | |
| | | (0.021) | (0.022) | (0.027) | (0.021) | (0.029) | (0.029) | (0.037) | |
| Female unemployment in other age groups | | | −0.013 | | | | | | |
| | | | (0.065) | | | | | | |
| Male unemployment in other age groups | | | −0.047 | | | | | | |
| | | | (0.055) | | | | | | |
| Female real wage in own age group | | | | 0.005 | | | | | |
| | | | | (0.009) | | | | | |
| Male real wage in own age group | | | | −0.001 | | | | | |
| | | | | (0.006) | | | | | |
| Female–Male UE gap in own age group | | | | | | | | | 0.094** |
| | | | | | | | | | (0.022) |
| Area and time fixed effects | Yes | Yes | Yes | Yes | Yes | Yes | Yes | Yes | Yes |
| Linear age-in-years control | Yes | Yes | Yes | Yes | Yes | No | No | No | Yes |
| Age group fixed effects | No | No | No | No | No | Yes | Yes | Yes | No |
| Age group × period FEs | No | No | No | No | No | No | Yes | Yes | No |
| Age group × areas FEs | No | No | No | No | No | No | No | Yes | No |
| Other demographic controls | Yes | Yes | Yes | Yes | Yes | Yes | Yes | Yes | Yes |
| Area-specific linear trends | No | No | No | No | Yes | No | No | No | No |
| Observations | | | | | 86,731 | | | | |

*Notes.* Standard errors clustered on police force area and age group in parentheses. 'Other demographic controls' include dummies for ethnicity category, qualification level and religious denomination, number of children and a dummy to indicate the presence of at least one child under the age of five in the household. **Significant at 1%. *Significant at 5%.

theory suggests that potential wages of men and women might also matter for the incidence of abuse. Therefore, we add measures of local female and male mean hourly real wage rates within the own age group in specification (4). Controlling for wage-effects in this way leaves the marginal effects for male and female unemployment largely unchanged. The estimated wage effects are small and insignificant.[23] Specification (5) shows that our estimates are robust to the introduction of area-specific linear time trends.

Specifications (1)–(5) use the respondent's age-in-years as a control variable. This has the advantage of allowing us to use the exact information on the respondent's age. In contrast, our labour market variables are measured at the age group level. The next three specifications verify that our results are robust to alternative age controls. In specification (6), we replace the age-in-years variable with dummy variables indicating the respondent's age group, thus allowing for age group fixed effects along with the area and period fixed effects. In specification (7), we interact the age group dummies with the period dummies, thus allowing each age group to have a separate non-linear trend. In specification (8), we further interact the age group dummies with the area dummies, thus allowing each PFA to be associated with a separate fixed effect for each age group. Specification (8) is particularly restrictive as the fixed effects net out any changes in the unemployment rates across age groups over time as well as differences between age groups across areas. However, the point estimates remain of a similar magnitude to those in our preferred specification (2), although they lose some precision. This indicates that across all specifications our identification is primarily driven by age-area-time differences in unemployment rates.

An evident feature of the results in Table 4 is that the estimated effects of female and male unemployment are of very similar absolute magnitude but of opposite sign. This suggests that what matters for the incidence of abuse is not the overall level of unemployment but rather the unemployment gender gap. Hence, in specification (9), we report the estimated marginal effect of the linear difference between the female and male unemployment rates within the own age group and of the total unemployment rate in the own age group. The estimated effect of the unemployment gender gap is noticeably strong, whereas the estimated effect of the overall unemployment rate is not statistically significant. Specification (9) will also serve as the benchmark regression for our IV analysis below where we will also focus on the gender unemployment gap.[24]

Table 5 presents corresponding results for non-physical abuse. The estimated marginal effects for this alternative outcome variable are strikingly similar to those for physical abuse.

[23] In fact, the coefficient have the 'wrong' signs. In order to look further into this, we obtained alternative measures of local wages from the Annual Survey of Hours and Earnings (ASHE) which is based on a one per cent sample of individuals from National Insurance records. Using this alternative data source, the coefficient on wages has the expected sign but remains statistically insignificant.

[24] As for our non-reported demographic control variables two factors stand out. Women with academic qualifications at A-level or above are less at risk of abuse. In contrast, there is a strong and significant positive correlation between the number of children and the incidence of abuse. To the extent that children reduce their mother's earnings capacity, this result is in line with our theoretical prediction. More generally, in our model any increase in the gains from marriage over divorce for the wife – obtaining from children or any other source – will be exploited by an abusive husband and make violence more likely. However, it is also possible that children are a cause of extra stress within a partnership and that this provides a trigger for more violence. An Appendix with expanded versions of Tables 4 and 5 which include the coefficients on the demographic control variables is available online on this JOURNAL's website.

Table 5

*Impact of Unemployment on Non-physical Abuse – Main Specification*

| Specification | (1) | (2) | (3) | (4) | (5) | (6) | (7) | (8) | (9) |
|---|---|---|---|---|---|---|---|---|---|
| Unemployment in own age group | −0.012 | | | | | | | | 0.021 |
| | (0.023) | | | | | | | | (0.024) |
| Female unemployment in own age group | | 0.102** | 0.108** | 0.110** | 0.104** | 0.078* | 0.093* | 0.087 | |
| | | (0.037) | (0.038) | (0.038) | (0.037) | (0.040) | (0.042) | (0.048) | |
| Male unemployment in own age group | | −0.081** | −0.074* | −0.061 | −0.085** | −0.090* | −0.066 | −0.077 | |
| | | (0.030) | (0.032) | (0.037) | (0.031) | (0.039) | (0.039) | (0.048) | |
| Female unemployment in other age groups | | | 0.031 | | | | | | |
| | | | (0.080) | | | | | | |
| Male unemployment in other age groups | | | 0.035 | | | | | | |
| | | | (0.068) | | | | | | |
| Female real wage in own age group | | | | −0.002 | | | | | |
| | | | | (0.010) | | | | | |
| Male real wage in own age group | | | | 0.008 | | | | | |
| | | | | (0.007) | | | | | |
| Female–Male UE gap in own age group | | | | | | | | | 0.093** |
| | | | | | | | | | (0.032) |
| Area and time fixed effects | Yes | Yes | Yes | Yes | Yes | Yes | Yes | Yes | Yes |
| Linear age-in-years control | Yes | Yes | Yes | Yes | Yes | No | No | No | Yes |
| Age group fixed effects | No | No | No | No | No | Yes | Yes | Yes | No |
| Age group × period FEs | No | No | No | No | No | No | Yes | Yes | No |
| Age group × areas FEs | No | No | No | No | No | No | No | Yes | No |
| Other demographic controls | Yes | Yes | Yes | Yes | Yes | Yes | Yes | Yes | Yes |
| Area-specific linear trends | No | No | No | No | Yes | No | No | No | No |
| Observations | 86,731 | | | | | | | | |

*Notes.* See Table 4. **Significant at 1%. *Significant at 5%.

Table 6 breaks the estimated effect of the gender unemployment gap down by population subgroup in three dimensions. The top panel shows that the relationship is apparent for all bar the eldest age group. That nothing is found in the oldest age group is not entirely surprising, given the low incidence of domestic violence reported in this age group.

The lower left panel in Table 6 splits the respondents into those with 'high' educational attainment (A-level or above) *versus* those with 'low' attainment (GCSE level or below).[25] One may argue that individuals' with lower qualifications are more at risk of unemployment and that, as a consequence, they may be more affected by gender unemployment gap in terms of the incidence of abuse.[26] While the point estimate is higher for low qualified women, the difference in the estimated effects is not statistically significant.

One may similarly argue that female unemployment is less relevant when the labour force participation (LFP) rate is relatively low. To consider this, we calculate the average female LFP over the sample period for each PFA-age group cell and partition the cells into those with above *versus* below median female LFP rate. Estimates by subgroup are reported in the lower right panel of Table 6. Again, while not statistically significantly different, the point estimates suggest that the effect of the gender unemployment gap on the incidence abuse is, if anything, stronger when the female LFP is higher.

The observed relationship between the gender-profile of unemployment and intimate partner violence can be expected to be particular to this outcome and not hold for general victim experience of crime. To verify this, we replace our main outcome variables with other reported crime outcomes. The BCS respondents are asked whether, over the past 12 months, they have experienced theft from their person

Table 6

*Impact of Unemployment Gender Gap on Abuse by Population Subgroup*

| | Age group | | | |
|---|---|---|---|---|
| | 16–24 | 25–34 | 35–49 | 50–59 |
| | 0.082** | 0.122** | 0.128* | −0.047 |
| | (0.030) | (0.044) | (0.063) | (0.096) |
| | Own qualification | | Female LFP in cell | |
| | Low | High | Low | High |
| | 0.114* | 0.089** | 0.075** | 0.137** |
| | (0.053) | (0.024) | (0.025) | (0.045) |
| Observations | 86,731 | | | |

*Notes.* The Table reports average marginal effects from three probit estimations of the impact of the unemployment gender gap on physical abuse, with the same set of controls as in specification (2) in Table 4. **Significant at 1%. *Significant at 5%.

[25] The 'high' qualifications are effectively those that require undertaking post-compulsory education.

[26] However, noting that the earnings drop associated with unemployment tends to be larger among individuals with higher qualifications, the effect could in principle go in the either direction.

or been a victim of a violent assault.[27] The results of this analysis are displayed in Table 7. For both theft and violence we find, in line with the literature (Raphael and Winter-Ebmer, 2001; Öster and Agell, 2010), that the probability of reporting having been a victim of crime increases with total unemployment. Moreover, unlike domestic abuse where there can be expected to be a direct power relationship between the perpetrator and the victim, these outcomes if anything increase with both the male and the female rate of unemployment.

To summarise, consistent with the literature, we find no evidence to support the view that total unemployment increases domestic abuse. Instead, our results suggest that male and female unemployment have distinct impacts on the incidence of domestic abuse: increases in male unemployment are associated with declines in domestic abuse while increases in female unemployment have the opposite effect. These findings are consistent with economic theory. The magnitude of the estimated relationships imply:

(*i*) that a 3.7 percentage point increase in male unemployment, as observed in England and Wales between 2004 and 2011, causes a decline in the incidence of domestic abuse of between 10.1% and 12.1%; and

(*ii*) that the 3.0 percentage point increase in female unemployment over the sample period causes an increase in the incidence of domestic abuse of between 9.1% and 10.3%.

Table 7

*Impact of Unemployment on Experience of Crime*

| Specification | Theft from person | Theft from person | Violence against person | Violence against person |
|---|---|---|---|---|
| Total unemployment in own age group | 0.099** (0.017) | | 0.035* (0.015) | |
| Female unemployment in own age group | | 0.042 (0.028) | | 0.039 (0.028) |
| Male unemployment in own age group | | 0.056** (0.020) | | 0.004 (0.021) |
| Area and time fixed effects | Yes | Yes | Yes | Yes |
| Linear age-in-years control | Yes | Yes | Yes | Yes |
| Other demographic controls | Yes | Yes | Yes | Yes |
| Observations | 86,725 | 86,725 | 86,726 | 86,726 |

*Notes.* See notes to Table 4. **Significant at 1%. *Significant at 5%.

[27] In both outcomes, the victim is present at the time of the crime, so gender is readily identifiable. In the case of theft, as this crime is mainly an opportunist event, the gender of either victim or perpetrator should however play only a minor role. Considering violence, one might expect that in cases of affray the victim and assailant often share the same gender. The exact questions answered by the respondents were: 'Was anything you were carrying stolen out of your hands or from your pockets or from a bag or case?' and 'Has anyone, including people you know well, deliberately hit you with their fists or with a weapon of any sort or kicked you or used force or violence in any other way?'.

### 3.3. *Extended Results: Area Level Controls*

Our estimates in the previous subsection would be biased if there were omitted variables that are correlated with local unemployment and that affect the incidence of domestic abuse. For example, a positive effect of unemployment on crime in general may trigger a response by the criminal justice system, such as increased police efforts or higher incarceration rates. If the response by the criminal justice system reduces domestic abuse by increasing deterrence, omitting controls related to the general level of criminal activity and the judiciary biases the estimated effect of unemployment on domestic abuse. Similarly, assuming that the consumption of alcohol and drugs is correlated with unemployment and also affects domestic abuse, omitting these factors from the regression again biases the estimates.[28] Additionally, selective migration might confound our estimates. For example, employment-driven migration of low-skilled men from areas with high local unemployment to areas with low local unemployment creates a downward bias (due to 'compositional effects') if low-skilled males have a higher propensity to abuse their partners than high-skilled males. To mitigate such omitted-variables bias, we now control extensively for observable institutional and demographic covariates at the PFA-level.

The results for physical abuse are shown in panel (*a*) of Table 8. Specification (2) repeats our basic specification from Table 4 for convenience. In specification (10), we add a set of controls that capture the general level of criminal activity and the potential response by the criminal justice system to it. In particular, we include *per capita* measures of violent and non-violent crimes. We include *per capita* measures of police force manpower and a proxy for the 'efficiency' of the criminal justice system: the average time from charge to magistrate court appearance. Overall, the inclusion of these crime-related controls leaves our key estimates unchanged. This suggests that variation in overall crime rates and policing and criminal justice efforts do not confound our estimated effects of unemployment on domestic abuse.

Specification (11) includes a measure of the hospitalisation rate for alcohol-related conditions as well as a *per capita* measure of drugs possession.[29] Adjusting for the cyclical consumption of criminogenic commodities in this way does not alter our main finding that male and female unemployment have opposite-signed effects on the incidence of physical abuse. In specification (12), we account for the possibility of skill-selective migration by including the qualification distribution in the respondent's own age group. Specification (13) controls directly for area-level migration by including the number of in and out-migrants as a percentage of the PFA population in the respondent's own age group. In each case, the estimated marginal effects of gender-specific unemployment remain largely unaffected.

The two remaining specifications provide additional robustness checks. Specification (14) shows that our results are robust to the introduction of controls for the average own age group female and male unemployment rates in neighbouring PFAs. Specification (15) shows that our main findings remain intact also when we include

[28] The association between business cycles and alcohol consumption is not clear cut. For instance, Dee (2001) notes that average drinking is generally pro-cyclical but finds that binge-drinking is counter-cyclical.

[29] Information on hospitalisation rates for alcohol-related conditions in particular is only available for England. This accounts for the drop in the number of observations in this particular specification.

Table 8

*Impact of Unemployment on Physical Abuse and Non-physical Abuse – Additional Controls*

| Specification | (2) | (10) | (11) | (12) | (13) | (14) | (15) |
|---|---|---|---|---|---|---|---|
| *Panel (a): Physical abuse* | | | | | | | |
| Female unemployment in own age group | 0.097** | 0.096** | 0.102** | 0.087** | 0.097** | 0.107** | 0.092** |
| | (0.027) | (0.027) | (0.028) | (0.027) | (0.027) | (0.028) | (0.026) |
| Male unemployment in own age group | −0.090** | −0.088** | −0.107** | −0.086** | −0.089** | −0.069** | −0.109** |
| | (0.021) | (0.021) | (0.021) | (0.025) | (0.021) | (0.026) | (0.021) |
| *Panel (b): Non-physical abuse* | | | | | | | |
| Female unemployment in own age group | 0.102** | 0.101** | 0.105** | 0.091* | 0.104** | 0.109** | 0.092* |
| | (0.037) | (0.038) | (0.038) | (0.039) | (0.037) | (0.039) | (0.037) |
| Male unemployment in own age group | −0.081** | −0.080** | −0.090** | −0.077* | −0.083** | −0.073* | −0.104** |
| | (0.030) | (0.031) | (0.031) | (0.034) | (0.031) | (0.037) | (0.030) |
| Local area crime-related controls | No | Yes | No | No | No | No | No |
| Local area drugs and alcohol | No | No | Yes | No | No | No | No |
| Local area qualifications distribution | No | No | No | Yes | No | No | No |
| Selective migration | No | No | No | No | Yes | No | No |
| Unemployment in neighbouring areas | No | No | No | No | No | Yes | No |
| Health and marital status | No | No | No | No | No | No | Yes |
| Observations | 86,731 | 86,731 | 80,011 | 86,731 | 86,731 | 86,731 | 86,674 |

*Notes.* Standard errors clustered on police force area and age group in parentheses. All specifications include area and time fixed effects, linear age-in-years control and other demographic controls (see notes to Table 4). Local area crime related-controls include police force manpower per 10,000 capita, violent and non-violent crimes per 10,000 capita, and average time from charge to magistrate court appearance. Local area drugs and alcohol include the number of arrests for drugs possession per 10,000 capita and the number of alcohol-related hospitalisations per 10,000 capita. Selective migration includes the number of in and out-migrants as a percentage of the PFA population in the respondent's own age and gender group. For a detailed description of controls used in this section, see Appendix B. **Significant at 1%. *Significant at 5%.

controls that capture a respondent's marital and health status (measured at the time of the interview and hence after the period to which the abuse information pertains).

Panel (*b*) of Table 8 provides the corresponding extended results for non-physical abuse. Again, the general conclusion is that the estimated effects of unemployment by gender are robust to the inclusion of further controls. The results presented in this subsection thus suggest that our initial finding that female unemployment increases domestic abuse, while male unemployment reduces it, is robust to including a wide variety of observable institutional and demographic covariates at the PFA level.

### 3.4. *Instrumental Variables Estimation*

The analysis so far has treated the local unemployment variables as exogenous regressors. Concerns about potential omitted variables motivated our use of additional regressors in subsection 2.3. However, this may not have entirely solved the potential issue of omitted variables and would not address any potential problem of simultaneity. Solving these problems requires constructing measures of local labour market conditions that do not reflect characteristics of female and male workers, which could be affected by violence itself, or unobservables that might be correlated with violence. Hence, as a final robustness check, we also consider an instrumental variables approach. Building on the work of Bartik (1991) and Blanchard and Katz (1992), we interact the initial local industry composition of employment with the corresponding national industry-specific trends in unemployment.

Specifically, we use APS data on local PFA industry composition by gender and age group at baseline, defined as the calendar year 2003, which we combine with APS data on industry unemployment rates by gender and age group at the national level over the sample period.[30] For each PFA, gender, age group and time period we construct an industry-predicted unemployment rate as follows,

$$\widehat{UNEMPL}^{h}_{jtg} = \sum_{k} \psi^{h}_{jgk} UNEMPL^{h}_{ktg}, \tag{11}$$

where $\psi^{h}_{jgk}$ is the share of industry $k$ among employed individuals of gender $h$ and age group $g$ in PFA $j$ at baseline, and where $UNEMPL^{h}_{ktg}$ is the unemployment rate, at the national level, in industry $k$ among individuals of gender $h$ and age group $g$ in time period $t$. Hence, (11) is a weighted average of the national industry-specific unemployment rates where the weights reflect the baseline local industry composition in the relevant gender and age group. The weights are thus fixed over time and do not reflect local sorting into industries over the sample period.

[30] Eight industries are used in the analysis based on a condensed version of the UK Standard Industrial Classification of Economic Activities, SIC (2007): 'agriculture, forestry, fishing, mining, energy and water supply', 'manufacturing', 'construction', 'wholesale, retail & repair of motor vehicles, accommodation and food Services', 'transport and storage, information and communication', 'financial and insurance activities, real estate activities, professional, scientific & technical activities, administrative & support services', 'public admin and defence, social security, education, human health & social work activities', 'other services'. The 'industry unemployment rate' is defined as the unemployed by industry of last job as percentage of economically active by industry.

Our approach draws on recent work by Albanesi and Sahin (2013) who, using US data, show how the gender gap in unemployment tends to vary over the business cycle. In particular, they find that unemployment rises more for men than for women during recessions, and also decreases more for men in subsequent recoveries. The authors also explore the role played by gender differences in industry structure. Specifically with respect to the recession in the late 2000s, Albanesi and Sahin show how gender differences in industry composition explain around half of the difference in the observed unemployment growth. Based on this observation, and on our previous finding that unemployment appears to matter for the incidence of domestic abuse only in the form of the unemployment gender gap, our IV analysis is focused on estimating models where the incidence of domestic violence is related to the female–male unemployment gender gap. We instrument for the actual gender gap using the corresponding industry-predicted gender gap in unemployment.

Table 9 presents the results for two different specifications, each estimated as a basic probit and as an IV probit model. Specification (1) in Table 9 includes the same controls as in specification (2) in Table 4. Hence, the difference is that here we include the unemployment rates in the own age group in the form of the gender gap rather than as levels. Specification (2) in Table 9 includes the same controls as in specification (5) in Table 4. The probit estimated average marginal effects of the gender unemployment gap on physical and non-physical abuse reported in columns (1*a*) and (2*a*) are naturally in line with the corresponding estimates in Tables 4 and 5.

Turning to the IV probit estimates, panel (*a*) of Table 9 confirms that our instrument is indeed a strong and relevant predictor of the gender unemployment gap in the own age group. More precisely, the estimates show that the actual variation in gender unemployment gap trends across PFAs and age groups is strongly positively

Table 9

*Impact of Unemployment on Physical Abuse – Instrumental Variables Estimation*

| Specification | (1*a*) Probit | (1*b*) IV Probit | (2*a*) Probit | (2*b*) IV Probit |
|---|---|---|---|---|
| *Panel (a): Gender unemployment gap in own age group* | | | | |
| Predicted unemployment gender gap in own age group | | 1.733** | | 1.723** |
| | | (0.106) | | (0.102) |
| *Panel (b): Physical abuse* | | | | |
| Gender unemployment gap in own age group | 0.090** | 0.104* | 0.089** | 0.105* |
| | (0.021) | (0.049) | (0.021) | (0.049) |
| *Panel (c): Non-physical abuse* | | | | |
| Gender unemployment gap in own age group | 0.081** | 0.083 | 0.084** | 0.081 |
| | (0.031) | (0.062) | (0.031) | (0.063) |
| Area and time fixed effects | Yes | Yes | Yes | Yes |
| Linear age-in-years control | Yes | Yes | Yes | Yes |
| Other demographic controls | Yes | Yes | Yes | Yes |
| Area-specific linear time trends | No | No | Yes | Yes |
| Observations | 86,731 | | | |

*Notes.* See notes to Table 4. **Significant at 1%. *Significant at 5%.

related to the corresponding variation in the unemployment gap trends predicted using local variation in industry structure at baseline.

The IV probit estimated average marginal effects of the gender unemployment gap on the incidence of domestic abuse are reported in columns (1*b*) and (2*b*). For physical abuse we find that, for both specifications, the IV estimated marginal effects are slightly larger than, but not statistically significantly different from, the corresponding probit estimated effects. Each estimated marginal effect is also statistically significant. For non-physical abuse, the IV probit estimated average marginal effects of the gender unemployment gap are also very similar to the basic probit estimated effects. However, due to lower precision, they are not statistically significant. Overall, we view our IV estimates as evidence that our basic probit estimates do not exaggerate the impact of unemployment on domestic abuse.

## 4. Concluding Comments

This article examines the effect of unemployment in England and Wales on partner abuse against women. The geographical variation in unemployment in these countries induced by the Great Recession provides an interesting context in which to look at domestic abuse. Our empirical approach was motivated by a theoretical model in which partnership provides insurance against unemployment risk through the pooling of resources. The key theoretical result is that an increased risk of male unemployment lowers the incidence of intimate partner violence, while an increased risk of female unemployment leads to a higher rate of domestic abuse. We demonstrate that this prediction accords well with evidence from the BCS matched to geographically disaggregated labour market data. In particular, our empirical results suggest that a 1 percentage point increase in the male unemployment rate causes a decline in the incidence of physical abuse against women of around 3%, while a corresponding increase in the female unemployment rate has the opposite effect. Moreover, our results also rationalise findings in previous studies of near zero effects of the overall rate of unemployment on domestic violence.

Overall, our theoretical model and empirical results contrast the conventional wisdom that male unemployment in particular is a key determinant of domestic violence. Quite the contrary, latent abusive males who are in fear of losing their jobs or who have lost their jobs may rationally abstain from abusive behaviour, as they have an economic incentive to avoid divorce and the associated loss of spousal insurance. However, when women are at a high risk of unemployment, their economic dependency on their spouses may prevent them from leaving their partners. This in turn might prompt male partners with a predisposition for violence to reveal their abusive tendencies. Thus, high female unemployment leads to an elevated risk of intimate partner violence. From a policy perspective, it is therefore conceivable that policies designed to enhance women's employment security could prove an important contributor to domestic violence reduction.

## Appendix A. Proofs

*Proof of Lemma 1.* We start by noting that, due to the functional form, $M(\pi_h, \pi_w, \varepsilon, \hat{\phi})$ is a continuously differentiable function of $(\pi_h, \pi_w, \hat{\phi})$ and $D(\pi_w)$ is a continuously differentiable function of $\pi_w$. Differentiating yields that $\partial M/\partial \pi_h < 0$, $\partial D/\partial \pi_h = 0$, $\partial M/\partial \pi_w < 0$, and $\partial D/\partial \pi_w < 0$, and, importantly,

$$\frac{\partial(M-D)}{\partial \pi_h} < 0 \quad \text{and} \quad \frac{\partial(M-D)}{\partial \pi_w} > 0, \tag{A.1}$$

where the latter inequality follows from concavity of $v(\cdot)$. Hence, an increase in the wife's unemployment risk makes marriage more attractive to her, as the loss in earnings associated with unemployment has a larger negative impact on her utility when she does not have access to her partner's income.

Next, we define

$$\pi_h' \equiv \begin{cases} 0 & \text{if } M(0,0,0,1) \leq D(0) \\ \sup\{\pi_h \in [0,1] | M(\pi_h, 0, 0, 1) \geq D(0)\} & \text{if } M(0,0,0,1) > D(0) \end{cases} \tag{A.2}$$

and

$$\pi_h'' \equiv \begin{cases} 1 & \text{if } M(1,1,0,1) \geq D(1) \\ \inf\{\pi_h \in [0,1] | M(\pi_h, 1, 0, 1) \leq D(1)\} & \text{if } M(1,1,0,1) < D(1) \end{cases}. \tag{A.3}$$

Consider the case where $M(0,0,0,1) > D(0)$, the second case in (A.2). By Assumption 1, $M(1,0,0,1) < D(0)$. Hence, it follows that $\pi_h' \in (0,1)$ and is the unique critical value for $\pi_h$ at which $M = D$ given $\pi_w = 0$ (and $\varepsilon = 0$ and $\hat{\phi} = 1$). Similarly, consider the case where $M(1,1,0,1) < D(1)$, the second case in (A.3). By Assumption 2, $M(0,1,0,1) > D(1)$. Hence, it follows that $\pi_h'' \in (0,1)$ and is the unique critical value for $\pi_h$ at which $M = D$ given $\pi_w = 1$ (and $\varepsilon = 0$ and $\hat{\phi} = 1$). Next, we verify that $\pi_h' < \pi_h''$. This follows trivially if $\pi_h' = 0$ and/or $\pi_h'' = 1$. Hence, consider the case where $\pi_h' > 0$ and $\pi_h'' < 1$ (as in Figure 1). Note that since, per definition of $\pi_h'$, $M(\pi_h', 0, 0, 1) = D(0)$, and using (A.1) it follows that $M(\pi_h', 1, 0, 1) > D(1)$ and hence that $\pi_h'' > \pi_h'$.

Next, we verify that (7) has a solution in the unit interval if and only if $\pi_h \in [\pi_h', \pi_h'']$. Consider the case where $\pi_h' > 0$. Then, $M(\pi_h, \pi_w, 0, 1) > D(\pi_w)$ at any $(\pi_h, \pi_w) \in [0, \pi_h') \times [0,1]$, implying that (7) does not have a solution in the unit interval. Similarly, consider the case where $\pi_h'' < 1$. Then, $M(\pi_h, \pi_w, 0, 1) < D(\pi_w)$ for any $(\pi_h, \pi_w) \in (\pi_h'', 1] \times [0,1]$, implying that (7) does not have a solution in the unit interval. Thus, (7) can have a solution in the unit interval only if $\pi_h \in [\pi_h', \pi_h'']$. Consider then some $\pi_h \in (\pi_h', \pi_h'')$. By definition of $\pi_h'$ and $\pi_h''$ if follows that $M(\pi_h, 0, 0, 1) < D(0)$ and $M(\pi_h, 1, 0, 1) > D(1)$. It then follows from continuity of the value functions and (A.1) that (7) has a unique solution we denote by $\hat{\pi}_w(\pi_h) \in (0,1)$.

Implicitly differentiating (7) yields that

$$\frac{\partial \hat{\pi}_w}{\partial \pi_h} = -\frac{\partial(M-D)/\partial \pi_h}{\partial(M-D)/\partial \pi_w} > 0, \tag{A.4}$$

where the sign follows from (A.1).

The sign of the derivatives of $\hat{\pi}_w(\pi_h)$ with respect to the partners' wages follow in a similar way from the observation that

$$\frac{\partial(M-D)}{\partial \omega_h} > 0 \quad \text{and} \quad \frac{\partial(M-D)}{\partial \omega_w} < 0, \tag{A.5}$$

where the latter inequality follows due to concavity of $v(\cdot)$.

*Proof of Proposition 1.* We first define the husband's expected utility in the case of divorce,

$$D(\pi_h, \varepsilon) \equiv \mathrm{E}(u_h^d | \pi_h) - \alpha_h - \xi\varepsilon, \tag{A.6}$$

where $\mathrm{E}(u_h^d|\pi_h)$ is defined analogously to (4). The husband's expected utility from continued marriage on the other hand is type-dependent,

$$M(\pi_h, \pi_w, \varepsilon; \theta) = \mathrm{E}\left[u_h^m|(\pi_h, \pi_w)\right] - \delta_\theta \kappa(\theta, \varepsilon) - \xi\varepsilon, \tag{A.7}$$

where $\mathrm{E}\left[u_h^m|(\pi_h, \pi_w)\right]$ is defined analogously to (6). In particular, we obtain that a husband of type $N$ ranks the possible outcomes with respect to marriage and behavioural effort in the following way:

$$M(\pi_h, \pi_w, 1; N) > M(\pi_h, \pi_w, 0; N) > D(\pi_h, 0) > D(\pi_h, 1). \tag{A.8}$$

To see this, note that the first inequality follows from part (*i*) of Assumption 4, the second inequality follows from part (*ii*) of Assumption 4 and the third inequality is trivial. In contrast, a husband of type $V$ ranks the possible outcomes in the following way:

$$M(\pi_h, \pi_w, 0; V) > M(\pi_h, \pi_w, 1; V) > D(\pi_h, 0) > D(\pi_h, 1). \tag{A.9}$$

The first inequality follows from the assumption that $\delta_V = 0$. The second inequality follows from the fact that $\alpha_h > \xi$ which is implied by the combination of parts (*i*) and (*ii*) of Assumption 4.

The key difference between (A.8) and (A.9) is that a husband of type $V$ does not value the reduction in the risk of violence associated with the effort $\varepsilon = 1$ whereas a husband of type $N$ values it more than its cost.

There are four possible pure strategy profiles that the husband can adopt:

*Strategy profile (1):* separation with $(\varepsilon', \varepsilon'') = (0, 1)$;
*Strategy profile (2):* separation with $(\varepsilon', \varepsilon'') = (1, 0)$;
*Strategy profile (3):* pooling with $(\varepsilon', \varepsilon'') = (1, 1)$; and
*Strategy profile (4):* pooling with $(\varepsilon', \varepsilon'') = (0, 0)$.

We will consider each possible pure strategy profile within each regime.

*Regime* $R_1$. Given that $(\pi_h, \pi_w) \in R_1$, the wife obtains a higher expected pay-off from marriage than from divorce with any husband of type $\theta$ and any effort choice $\varepsilon$ by the husband. We now consider the four possible pure strategy profiles in turn:

*Strategy profile* (1). Bayesian updating implies that $\hat{\phi}(0) = 1$ and $\hat{\phi}(1) = 0$, and the wife rationally chooses to remain married at either choice of $\varepsilon$, $\chi' = \chi'' = m$. According to (A.8) and (A.9), each type of husband obtains his most preferred outcome and hence has no incentive to deviate, confirming that this is a PBE.

*Strategy profile* (2). Bayesian updating implies that $\hat{\phi}(0) = 0$ and $\hat{\phi}(1) = 1$, and the wife rationally chooses to remain married at either choice of $\varepsilon$, $\chi' = \chi'' = m$. In this case neither type of husband obtains his most preferred outcome and, since the wife responds to either choice of $\varepsilon$ by continuing the marriage, each type of husband would have an incentive to deviate.

*Strategy profile* (3). Bayesian updating implies that $\hat{\phi}(1) = \phi$, while $\hat{\phi}(0)$ is not determined by Bayesian updating. Irrespective of how the wife updates her beliefs at $\varepsilon = 0$, she rationally chooses to remain married at either choice of $\varepsilon$, $\chi' = \chi'' = m$. Given this, a husband of type $V$ would be better off deviating to $\varepsilon = 0$.

*Strategy profile* (4). Bayesian updating implies that $\hat{\phi}(0) = \phi$, while $\hat{\phi}(1)$ is not determined by Bayesian updating. Irrespective of how the wife updates her beliefs at $\varepsilon = 1$, she rationally chooses to remain married at either choice of $\varepsilon$, $\chi' = \chi'' = m$. Given this, a husband of type $N$ would be better of deviating to $\varepsilon = 1$.

*Regime* $R_0$. In this regime, the wife's decision whether or not to remain married depends on her beliefs and on the husband's observed effort.

*Strategy profile* (1). Bayesian updating implies that $\hat{\phi}(0) = 1$ and $\hat{\phi}(1) = 0$. The wife then (by Assumptions A1 and A3) continues the marriage if and only if the husband makes the effort

$\varepsilon = 1$, that is $\chi'' = m$ and $\chi' = d$. A type $V$ would then be better of deviating to $\varepsilon = 1$ as by doing so he would avoid triggering divorce.

*Strategy profile* (2). Bayesian updating implies that $\hat{\phi}(0) = 0$ and $\hat{\phi}(1) = 1$. Given these updated beliefs, the wife rationally responds (by Assumption 3) to $\varepsilon = 0$ by continuing the marriage, that is $\chi' = m$. This then cannot be an equilibrium since a type $V$ husband could then deviate to $\varepsilon = 0$ and obtain is his most preferred outcome.

*Strategy profile* (3). Bayesian updating implies that $\hat{\phi}(1) = \phi$ and, by Assumption 3, the wife rationally responds to $\varepsilon = 1$ by continuing the marriage, $\chi'' = m$. Note that $\hat{\phi}(0)$ is not determined by Bayesian updating. Suppose that the wife, at $\varepsilon = 0$, believes that the husband is of type $V$, that is $\hat{\phi}(0) = 1$. She would then rationally respond to $\varepsilon = 0$ by choosing divorce, $\chi' = d$. Given this, and given the preference orderings in (A.8) and (A.9), neither husband type has any incentive to deviate. Note also that the out-of-equilibrium belief $\hat{\phi}(0) = 1$ satisfies the Choo–Kreps 'intuitive criterion'. For a husband of type $N$, $\varepsilon = 0$ is equilibrium dominated as this type, by choosing $\varepsilon = 1$, obtains his most preferred outcome in equilibrium. In contrast, a husband of type $V$ would benefit if the wife were to respond to $\varepsilon = 0$ by continuing the marriage.

*Strategy profile* (4). Bayesian updating implies that $\hat{\phi}(0) = \phi$ but does not determine $\hat{\phi}(1)$. Given this, and by Assumption 3, the wife rationally continues the marriage upon observing $\varepsilon = 0$, that is $\chi' = m$. Next, note that by (A.8) for a husband of type $N$ in particular to prefer to choose $\varepsilon = 0$ it must be that the wife responds to $\varepsilon = 1$ by divorcing, that is $\chi'' = d$. Hence, for this to be a PBE, $\hat{\phi}(1)$ must be such that the wife prefers divorce upon observing $\varepsilon = 1$. In particular, from Assumption 3 it must be that $\hat{\phi}(1) > \phi$. Such a PBE however does not satisfy the 'intuitive criterion'. For a husband of type $V$, $\varepsilon = 1$ is equilibrium dominated as this type, by choosing $\varepsilon = 0$, obtains his most preferred outcome in equilibrium. In contrast, a husband of type $N$ would benefit from deviating if the wife were to respond to $\varepsilon = 1$ by continuing the marriage. Hence, by the 'intuitive criterion' the wife's out-of-equilibrium beliefs must be $\hat{\phi}(1) = 0$, contradicting that she would choose $\chi'' = d$.

## Appendix B. Variable Descriptions

The following variables are used in subsection 3.3:

1 *Magistrate court timeliness:* This is a measure of the duration from first listing of an offence to completion, for defendants in indictable cases in magistrates courts, and hence captures the 'efficiency' of the criminal justice system, post-arrest. The data are released on an annual basis from the Ministry of Justice, and is at the local justice area (LJA) geography which coincides with the PFAs we use in the analysis.
2 *Police force manpower:* This variable refers to overall police manpower per 10,000 capita at PFA level. It is comprised of the number of (full-time equivalent) police officers, police community support officers and police staff. These data are released annually by the Home Office.
3 *Violent crime rate:* This is the number of recorded violent crimes per 10,000 capita at PFA level. The data are from the Home Office.
4 *Non-violent crime rate:* This is the number of recorded non-violent crimes per 10,000 capita at PFA level. The data are from the Home Office.
5 *Alcohol hospitalisations:* This is the number of alcohol hospitalisations per 10,000 capita at PFA level. This is from the local alcohol profiles for England data sets, available from the North West Public Health Observatory data, which are part of Public Health England. Note that these data are not available for the four Welsh PFAs. We aggregated the data up to PFA level from local authority level.

6 *Internal migration:* These are the number of in and out-migrants as a percentage of the PFA population in each age/gender group. The statistics are compiled using the data series 'Internal Migration by Local Authorities in England and Wales' which are released annually by the Office for National Statistics (ONS) to coincide with the mid-year population estimates. The data have received the 'National Statistics' accreditation and are understood to be the best official source of information on internal migration in England and Wales. The data are available by gender and in five-year age groups at local authority level. Here, we aggregated up to PFA level and using the APS defined age grouping.

7 *Drugs possession:* This is the number of arrests for possession per 10,000 capita at PFA level. These data are from the quarterly Home Office Offences Tables.

The data in (1)–(6) come from annual Tables, so have been interpolated to produce data at the period frequency.

*Royal Holloway, University of London*
*University of Munich and Ifo Institute*
*Royal Holloway, University of London*
*Royal Holloway, University of London*

*Submitted: 18 March 2014*
*Accepted: 18 November 2014*

Additional Supporting Information may be found in the online version of this article:

**Appendix C.** Complete Set of Estimated Marginal Effects.
**Appendix D.** A Simple Model of Household Bargaining Under Uncertainty.
**Data S1.**

## References

Aizer, A. (2010). 'The gender wage gap and domestic violence', *American Economic Review*, vol. 100(4), pp. 1847–59.

Aizer, A. and Dal Bó, P. (2009). 'Love, hate and murder: commitment devices in violent relationships', *Journal of Public Economics*, vol. 93(3), pp. 412–28.

Albanesi, S. and Sahin, A. (2013). 'The gender unemployment gap: trend and cycle', mimeo, Federal Reserve Bank of New York.

Angelucci, M. (2008). 'Love on the rocks: alcohol abuse and domestic violence in rural Mexico', *B.E. Journal of Economic Analysis and Policy: Contributions to Economic Analysis and Policy*, vol. 8(1), Article 43.

Bartik, T.J. (1991). *Who Benefits from State and Local Economic Development Policies?* Kalamazoo, MI: W. E. Upjohn Institute for Employment Research.

Bergstrom, T.C., Blume, L. and Varian, H. (1986). 'On the private provision of public goods', *Journal of Public Economics*, vol. 29(1), pp. 25–49.

Blanchard, O.J. and Katz, L.F. (1992). 'Regional evolutions', *Brookings Papers on Economic Activity*, vol. 23(1), pp. 1–76.

Bloch, F. and Rao, V. (2002). 'Terror as a bargaining instrument: a case study of dowry violence in rural India', *American Economic Review*, vol. 92(4), pp. 1029–43.

Browning, M., Chiappori, P.A. and Lechene, V. (2010). 'Distributional effects in household models: separate spheres and income pooling', Economic Journal, vol. 120(545), pp. 786–99.

Campbell, D., Carruth, A., Dickerson, A. and Green, F. (2007). 'Job insecurity and wages', Economic Journal, vol. 117(518), pp. 544–66.

Card, D. and Dahl, G. (2011). 'Family violence and football: the effect of unexpected emotional cues on violent behavior', *Quarterly Journal of Economics*, vol. 126(1), pp. 103–43.

Cho, I.K. and Kreps, D.M. (1987). 'Signaling games and stable equilibria', *Quarterly Journal of Economics*, vol. 102(2), pp. 179–221.

Dee, T.S. (2001). 'Alcohol abuse and economic conditions: evidence from repeated cross-sections of individual-level data', *Health Economics*, vol. 10(3), pp. 257–70.
Farmer, A. and Tiefenthaler, J. (1997). 'An economic analysis of domestic violence', *Review of Social Economy*, vol. 55(3), pp. 337–58.
Iyengar, R. (2009). 'Does the certainty of arrest reduce domestic violence? Evidence from mandatory and recommended arrest laws', *Journal of Public Economics*, vol. 93(1), pp. 85–98.
Öster, A. and Agell, J. (2010). 'Crime and unemployment in turbulent times', *Journal of the European Economic Association*, vol. 5(4), pp. 752–75.
Pollak, R.A. (2004). 'An intergenerational model of domestic violence', *Journal of Population Economics*, vol. 17(2), pp. 311–29.
Raphael, S. and Winter-Ebmer, R. (2001). 'Identifying the effect of unemployment on crime', *Journal of Law and Economics*, vol. 44(1), pp. 259–83.
Schmidt, S. (1999). 'Long-run trends in workers beliefs about their own job security: evidence from the General Social Survey', *Journal of Labor Economics*, vol. 17(S4), pp. 127–41.
Smith, K., Osborne, S., Lau, I. and Britton, A. (2012). 'Homicides, firearm offences and intimate violence 2010/11', Home Office Statistical Bulletin, vol. 2.
Tauchen, H.V., Witte, A.D. and Long, S.K. (1991). 'Violence in the family: a non-random affair', *International Economic Review*, vol. 32(2), pp. 491–511.
Warr, P.G. (1983). 'The private provision of a pure public good is independent of the distribution of income', *Economics Letters*, vol. 13(2-3), pp. 207–11.