Geographical variation in rates of common mental disorders in Britain: prospective cohort study

SCOTT WEICH, LIZ TWIGG, GLYN LEWIS and KELVYN JONES

Background There is little geographical variation in the prevalence of the common mental disorders. However, there is little longitudinal research.

Aims To estimate variance in rates of common mental disorders at individual, household and electoral ward levels prospectively.

Method A 12-month cohort study of 7659 adults aged 16–74 years in 4338 private households, in 626 electoral wards. Data were collected as part of the British Household Panel Survey. Common mental disorders were assessed using the 12-item General Health Questionnaire (GHQ). Ward-level socio-economic deprivation was measured using the Carstairs index.

Results Less than 1% of total variance, in onset and maintenance of common mental disorders and change in GHQ score between waves, occurred at ward level. However, 12% of variance, which is a statistically significant difference, was found at household level (a much smaller geographical unit) and this difference remained after further analyses.

Conclusions Ward level socio-economic deprivation does not influence the onset and maintenance of common mental disorders in Britain but local factors at the household level do. Reasons for this remain unclear.

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Cross-sectional studies suggest little geographical variation in the prevalence of the most common mental disorders of anxiety and depression after adjusting for individual characteristics (McCulloch, 2001; Pickett & Pearl, 2001; Wainwright & Surtees, 2003; Weich *et al*, 2003*a*). However, concluding that 'place doesn't matter' runs counter to the intuitive importance of location (Dorling, 2001; MacIntyre *et al*, 2002). Differential effects of place on the onset and outcome of common mental disorders may not be apparent in cross-sectional studies. Evidence that socio-economic adversity is associated with episode maintenance (Lorant *et al*, 2003; Hauck & Rice, 2004) suggests a longer episode duration in socio-economically deprived areas. Place effects also may vary with individual circumstances (Weich *et al*, 2003*b*). We aimed to estimate the variance in onset and maintenance of common mental disorders at individual, household and electoral ward levels, and also to test the hypothesis that ward-level socio-economic deprivation is associated with episode maintenance, after controlling for individual and household characteristics.

METHOD

Data were gathered during the first two waves of the British Household Panel Survey (BHPS), which was initially undertaken in 1991. The BHPS is an annual survey of individuals aged 16 years and over in a representative sample of private households in England, Wales and Scotland. First-wave members were selected via a two-stage, stratified clustered probability sample. Efforts are made to re-interview all original sample members in each subsequent year. Individuals aged 16–74 years at wave 1 who completed the 12-item General Health Questionnaire (GHQ; Goldberg & Williams, 1988) at both waves 1 and 2 were included in this analysis. The BHPS coordinators provided permission and facilitated the linkage of BHPS data to other geographically referenced datasets via each individual's electoral ward of residence at wave 1. This process did not threaten the anonymity of sample members.

Onset and maintenance of episodes of common mental disorders

Information on common mental disorders was gathered using the GHQ (Goldberg & Williams, 1988). Designed for case finding in community settings, with a sensitivity and specificity of about 80%, it has been widely validated against standardised clinical interviews. We followed evidence that common mental disorders may be represented validly as a single dimension encompassing comorbid anxiety and depression (Stansfeld & Marmot, 1992; Krueger, 1999; Vollebergh *et al*, 2001; Kendell & Jablensky, 2003).

We used the 'GHQ method' to identify the cases (Goldberg & Williams, 1988). Each GHQ item has four response categories. For example, responses to the question, 'Have you recently been unhappy and depressed?' are 'not at all', 'no more than usual', 'rather more than usual' and 'much more than usual'. The GHQ is scored in two ways, scoring each item either by the 'GHQ method' as present or absent (one point for either of the latter two responses, and zero otherwise), or by the Likert method (responses coded in order as 0, 1, 2 or 3; Goldberg & Williams, 1988). Those scoring 3 or more (out of 12) by the GHQ method were classified as cases (Goldberg & Williams, 1988; Weich & Lewis, 1998). Likert scores (range 0–36) more closely approximated a normal distribution and were used when the GHQ score was treated as a continuous outcome. 'Episode onset' describes non-cases at wave 1 on the GHQ who met the case criteria for common mental disorders at wave 2. 'Episode maintenance' describes individuals who met the case criteria at both waves.

Individual- and household-level risk factors

Age, gender, marital status, ethnicity, education, employment status, financial strain and number of current physical health problems were included as potential individual-level confounders of associations between area-level exposures and common mental disorders.

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There is significant variation in rates of common mental disorders between households, even after taking into account individual-level confounders (Weich *et al*, 2003*a*). Some exposures can be assigned only to households, such as overcrowding, household type, housing tenure and structural housing problems. This is not so for others, particularly income, for which data are commonly aggregated at household level (Weich *et al*, 2001). For occupational social class, stronger associations with rates of common mental disorders have been found between the social class of the head of the household than with individual social class, particularly among women (Weich & Lewis, 1998; Weich *et al*, 2003*b*). Household characteristics were assessed at wave 1, and included structural housing problems, household income, car access, tenure, social class (by head of household), overcrowding (more than two household members per bedroom) and household type (based on household composition). Structural housing problems were defined as any major problem or two or more minor problems from a list comprising damp, condensation, leaking roof and/or rot in wood. The BHPS data-set includes net income data that have been validated against official UK income distribution figures (Jarvis & Jenkins, 1995). Low income was defined as household income below half the median income for the sample.

Spatial scale

There were three potential 'area' levels above household level within this data-set: electoral ward, postcode sector (the primary sampling unit for the BHPS) and region. Electoral wards (2400 addresses on average, with mean population=5222, s.d.=3899) are currently the smallest geographical area at which BHPS data are available. Sensitivity analyses were undertaken by substituting each of the other two geographical levels for wards. The BHPS investigators and authors therefore agreed a method for matching respondents and characteristics of electoral wards, without disclosure of information that might permit identification of respondents.

Area-level socio-economic deprivation

The assessment of area-level exposures was limited by the absence of validated contextual measures and a dearth of evidence about which of the large number of compositional measures were likely to be associated with the prevalence of common mental disorders. We therefore chose, *a priori*, to use the Carstairs index of socio-economic deprivation (Morris & Carstairs, 1991), based on data collected in the 1991 census. The Carstairs index is based on *Z* scores of four person-level (compositional) variables for each ward: male unemployment, households with no car, overcrowding (more than one person per room) and head of household in Registrar General's social class IV or V. The BHPS investigators rounded *Z* scores to integer values and truncated the tails of the resulting distribution to protect respondents' identities.

Statistical analysis

Multi-level models were developed using MLwiN software (Goldstein *et al*, 1998). Null, random effects models were derived for persons nested in households, with households nested within wards (Snijders & Bosker, 1999). Individual-, household- and ward-level exposures were added subsequently. We analysed the onset of episodes separate from episode maintenance, using multi-level logistic regression. For binomial distributions, variance in the intercept term is neither constant across groups nor independent of mean values within the groups. A number of alternative approaches to ascertaining variance of the intercept term at higher levels can be used, including model linearisation using first-order Taylor expansion or simulation methods (Goldstein *et al*, 2002). We used a logit model based on the notion of a continuous latent variable in which a threshold defines the binary outcome (see Snijders & Bosker, 1999, p. 223). We assumed an underlying standard logistic distribution for the binary outcome (onset or not and maintenance or not, across two waves) at the individual level (level 1). This is justified by the threshold nature of the GHQ scoring method, but might be less suitable for discrete outcomes such as mortality.

Level 1 variance on this latent variable was the standardised logistic variance of $\pi^2/3=3.29$. When unexplained random variance at level 2 was indicated as r_0^2, the proportion of the total unexplained variance at this level was estimated (from a two-level null random intercept model) as $r_0^2/(r_0^2+3.29)$. In each of the logistic models, the constant term is the logit ($\log_e$ of the odds) of a person in the base (reference) category being either an individual experiencing episode 'onset' or episode 'maintenance'. The proportion of each onset or maintenance group was estimated from the constant term in the null model, which is equal to $\ln[p/(1-p)]$. Parameters were estimated using second-order Taylor expansion with predictive quasi-likelihood. Markov chain Monte-Carlo methods may improve the accuracy of such estimates but the method is computationally intensive and was used here only in the discussion of higher level variation. Statistical significance of individual fixed estimates was tested using a Wald test against a χ^2 distribution. Because difficulties are encountered when variances are close to zero, 95% interval estimates (the 'credible interval') derived from Markov chain Monte-Carlo procedures are reported for random model parameters.

The GHQ scores at wave 2 were also analysed as a continuous outcome, using hierarchical linear regression and controlling for GHQ score at wave 1. Finally, the stability of GHQ scores across waves was assessed using the intra-class correlation coefficient.

RESULTS

A total of 9518 individuals aged 16–74 years participated in the BHPS at wave 1. Of these, 8980 (94%) completed the GHQ at wave 1 and 7659 also did so at wave 2 (85% of those who completed the GHQ at wave 1, and 80% of the total baseline sample). The baseline prevalence of common mental disorders in the study sample was 24.6%. For episode onset analyses, 5809 individuals were nested within 3679 households, within 615 wards. For episode maintenance analyses, 1850 individuals were nested within 1566 households, within 511 wards.

Onset and maintenance of episodes of common mental disorders

In the null model, the rate of episode onset was 14.3% (95% CI 13.3–15.3) across all households and wards. As Table 1 shows, the estimated variance at the household level (13.9%) was statistically significant, but that at ward level (0.2%) was not. These variances were largely unchanged after adjusting for characteristics of individuals, households and wards (Table 1), or for GHQ score at baseline.

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Table 1 Variance (standard error), credible interval and percentage of total unexplained variance in the onset and maintenance of episodes of common mental disorders at the individual, household and electoral ward levels, for null and adjusted models

	Episode onset (n=5809)				Episode maintenance (n=1850)			
	Variance (s.e.)	Credible interval	P	Percentage of unexplained variance	Variance (s.e.)	Credible interval	P	Percentage of unexplained variance
Null model								
Individual level	3.29			85.9	3.29			87.5
Household level	0.53 (0.19)	0.23–0.86	0.005	13.9	0.45 (0.38)	0.00–1.29	0.23	12.0
Ward level	0.01 (0.02)	0.00–0.05	0.45	0.2	0.02 (0.03)	0.00–0.13	0.49	0.5
Model 2								
Individual level	3.29			85.2	3.29			65.2
Household level	0.55 (0.28)	0.03–0.97	0.05	14.2	1.73 (0.72)	0.38–2.92	0.02	34.3
Ward level	0.02 (0.03)	0.00–0.07	0.50	0.6	0.03 (0.05)	0.00–0.13	0.47	0.5
Model 3								
Individual level	3.29			82.3	3.29			66.6
Household level	0.69 (0.23)	0.33–1.07	0.002	17.3	1.62 (0.72)	0.10–2.84	0.03	32.8
Ward level	0.02 (0.026)	0.00–0.08	0.42	0.4	0.03 (0.06)	0.00–0.16	0.54	0.6

Model 2 is the null model plus individual and household-level characteristics; model 3 comprises model 2 plus area-level deprivation (Carstairs) scores.

A different pattern was observed for episode maintenance, the rate of which was 54.3% (95% CI 51.8–56.8) over 1 year. In the null model, neither variance at the household (12.0%) nor ward level (0.5%) was statistically significant. However, adjusting for individual and household characteristics resulted in an almost fourfold increase in the variance in episode maintenance at the household level (estimated variance=1.73, credible interval =0.38–2.92). Most of this increase in variance occurred on adjusting for individual-level variables (estimated variance=1.15, s.e.=0.56, credible interval=0.36–2.20), before household characteristics were introduced into the model. The adjusted variance at household level was statistically significant and was not altered on further adjusting for ward characteristics. None of these findings differed substantially when postcode sectors were substituted for wards, or when wards with five or fewer respondents were excluded.

General Health Questionnaire score as a continuous outcome

The intra-class correlation coefficient for GHQ score at waves 1 and 2 was +0.44. Multi-level analyses using GHQ score at wave 2 as a continuous outcome measure, adjusted for GHQ score at wave 1, confirmed previous findings. In the null model, 0.2% of the total (unexplained) variance in GHQ scores at wave 2 occurred at the ward level, compared with 87.5% and 12.3% at the individual and household levels, respectively. Ward-level variance was not statistically significant. Total variance in GHQ scores was reduced by 1.9% when individual- and household-level characteristics were included, and by a further 0.1% when ward-level exposures were introduced (Table 2).

Associations with ward-level deprivation

Maintenance, but not episode onset, was increased to a statistically significant degree among those living in wards with Carstairs scores in the highest (most deprived) quintile, compared with the lowest quintile wards, before adjusting for individual and household characteristics (odds ratio=1.28, 95% CI 1.01–1.62; P=0.04). However, none of these associations reached statistical significance after adjustment for lower-level variables, and there were no statistically significant overall trends with increasing ward-level deprivation (e.g. test for trend in unadjusted odds ratios for episode maintenance by Carstairs quintile χ^2=5.7, d.f.=4, P=0.22). The interaction between wave 1 case status and Carstairs score (by quintile) was not statistically significant in the unadjusted model (χ^2=3.26, d.f.=4, P=0.52), or in the fully adjusted model (with all individual and household variables) (χ^2=5.63, d.f.=4, P=0.23). No statistically significant association was found between area-level deprivation (Carstairs score) and GHQ score at wave 2 (adjusted for wave 1 GHQ score), even before adjusting for other potential confounders (regression coefficient, B, for top *v.* bottom Carstairs quintile=0.245, s.e.=0.16).

We found no evidence of statistically significant interactions between Carstairs scores (by quintile) and employment status in associations with either the onset (unadjusted χ^2=12.9, d.f.=8, P=0.11; adjusted χ^2=12.2, d.f.=8, P=0.14) or maintenance (unadjusted χ^2=7.7, d.f.=8, P=0.46; adjusted χ^2=8.0, d.f.=8, P=0.43) of episodes of common mental disorders.

DISCUSSION

Geographical variation in rates of common mental disorders

The view that place does not affect individual health is counter-intuitive. This study is one of the first to estimate variance in rates of common mental disorders prospectively. Such research is vital for establishing whether the lack of significant area-level variance in common mental disorders reported in cross-sectional studies might

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Table 2 Variance (standard error), credible interval and percentage of total unexplained variance in GHQ score at wave 2 (as a continuous measure, and adjusted for GHQ score at wave 1) at the individual, household and electoral ward levels, for null and adjusted models

	GHQ score at wave 2 (n=7659)[1]			
	Variance (s.e.)	Credible interval	P	Percentage of unexplained variance
Null model				
Individual level	17.29 (0.41)	16.62–17.97	<0.001	87.5
Household level	2.44 (0.35)	1.86–3.01	<0.001	12.3
Ward level	0.04 (0.04)	0.00–0.11	0.32	0.2
Total variance	19.77			
Model 2				
Individual level	16.99 (0.42)	16.31–17.70	<0.001	87.6
Household level	2.37 (0.36)	1.78–2.96	<0.001	12.2
Ward level	0.04 (0.06)	0.00–0.17	0.45	0.2
Total variance	19.40			
Model 3				
Individual level	16.98 (0.43)	16.29–17.70	<0.001	87.6
Household level	2.35 (0.36)	1.75–2.95	<0.001	12.1
Ward level	0.05 (0.06)	0.00–0.17	0.41	0.3
Total variance	19.38			

Model 2 is the null model plus individual and household-level characteristics; model 3 comprises model 2 plus area-level deprivation (Carstairs) scores.
GHQ, General Health Questionnaire.
1. Adjusted for GHQ score at wave 1.

mask differential effects of place on the onset and outcome of episodes of these disorders.

We found little evidence that episode maintenance was greatest in the most deprived wards. Although episode maintenance was more common (to a statistically significant degree) in the most deprived wards (by Carstairs score quintile), this association failed to reach statistical significance after adjusting for individual- and household-level characteristics. There was no statistically significant interaction between ward Carstairs score and baseline case status. These findings were confirmed when change in GHQ score between waves was modelled.

In null models, 0.5% or less of the variation in episode onset and maintenance occurred at electoral ward level. This is almost the same as was estimated for the cross-sectional prevalence of common mental disorders (Weich *et al*, 2003*a*). In contrast to our cross-sectional analyses, we found no evidence of statistically significant variation in the effects of area-level deprivation on common mental disorders with employment status at baseline. These findings confirm cross-sectional studies showing that variation in common mental disorders across areas the size of electoral wards is modest (Lewis & Booth, 1992; Duncan *et al*, 1995; McCulloch, 2001; Pickett & Pearl, 2001; Weich *et al*, 2003*b*).

Household-level effects

These results highlight the importance of modelling household as a distinct level, something that many studies overlook (McCulloch, 2001; Silver *et al*, 2002; Wainwright & Surtees, 2003). Our estimates of standard errors for variance at area level were less prone to bias than those arising from studies in which individual- and household-level exposures were conflated. Although the estimated proportion of variance in episode onset and maintenance at household level (12–14%) appeared greater than for prevalence (8%) (Weich *et al*, 2003*a*), credible intervals (equivalent to confidence limits) were considerably larger in analyses stratified by baseline case status. Differences in sample sizes may also explain why household-level variance reached statistical significance for episode onset but not episode maintenance in the null model. The former was unaffected by adjustment for the characteristics of households and individual household members.

Intriguingly, between-household variance in episode maintenance increased after adjusting for individual characteristics in particular. This was not the case for either episode onset or prevalence of common mental disorders (Weich *et al*, 2003*a*). This finding was verified using Markov chain Monte-Carlo methods. Thus, the effect of household on episode maintenance becomes more apparent after adjusting for characteristics of individual household members. This is analogous to the finding that variance in house prices across counties of southern England increases when house size is specified (Jones & Bullen, 1993). We found that the sharpest increase in household-level variance occurred on including financial strain (using individual responses) in the fixed part of the model. The effects of household-level factors emerged more clearly after controlling for factors associated with between-individual variation in episode maintenance.

These findings are consistent with evidence of spousal similarity in depressive symptoms (Dufouil & Alpérovitch, 2000). Intra-household factors subsequent to the onset of anxiety or depression in one or more members warrant closer scrutiny. Transient affective changes in one household member may have relatively little effect on the mental health of others, or indeed may even lead to 'resilient' coping and caring. If two or more household members experience an episode of common mental disorders, recovery does not appear to occur at random, but rather tends to happen (or not) synchronously, irrespective of individual and household social and economic circumstances. Other unmeasured factors in this study were life events (including resolution events), which often have consequences for all household members.

Limitations of this study

Measuring the common mental disorders

The study was limited by use of the GHQ rather than a standardised clinical interview.

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However, traditional objections to findings not based on clinical diagnostic categories are lessened by evidence that common mental disorders are most validly represented as a single dimension encompassing comorbid anxiety and depression (Krueger, 1999; Vollebergh *et al*, 2001; Kendell & Jablensky, 2003). The GHQ has been widely used in general population samples and is robust to retest effects (Pevalin, 2000). Nevertheless, associations between poverty and common mental disorders are generally larger in studies using clinical interviews (Meltzer *et al*, 1995). Because the GHQ is sensitive to recent change in psychological functioning, 'false positives' might have included individuals with mild or transient psychological disturbance. By contrast, individuals with chronic symptoms of anxiety and depression may be classed as non-cases (false negatives). This misclassification should have biased associations towards the null. Although physical ill-health also leads to 'false positives', study findings were adjusted for the number of current physical health problems. Those in lower occupational grades (Stansfeld *et al*, 1995) may underreport psychiatric symptoms on the GHQ compared with responses to a standardised clinical interview. This should have reduced individual-level variance in rates of common mental disorders.

The study was also limited by the absence of data on the duration of episodes of common mental disorders. Participants were interviewed on two occasions, separated by 12 months. 'Episode onset' was defined as the presence of common mental disorders at wave 2 among participants who did not meet criteria for caseness at wave 1. This definition refers to a specific episode of disorder occurring during the course of the study, irrespective of previous history. Most 'onset' episodes were likely to have been relapses, rather than first inceptions. 'Episode maintenance' was defined as the proportion of cases at wave 1 that also met criteria for caseness at wave 2. This may be viewed as implying continuous comorbidity throughout the year. Some wave 1 cases may have remitted and then relapsed between assessments, and a proportion of 'onset' cases could have experienced multiple episodes between assessments. Episodes that began and remitted between waves may have been missed among those identified as non-cases at both waves. However, the moderately high intra-class correlation between individual GHQ scores at waves 1 and 2 (r=+0.44) is consistent with limited intra-participant fluctuation in case status between waves.

Choice of spatial scale

The geographical scale at which contextual factors might have an impact on mental health remains unknown (Mitchell, 2001; MacIntyre *et al*, 2002), and previous studies have been undertaken at scales ranging from the household (Weich *et al*, 2003*a*,*b*) to UK electoral ward (average population 5500) (McCulloch, 2001) or equivalent (Reijneveld *et al*, 2000), to UK regions and US states. Two previous studies that found modest but statistically significant associations between area deprivation and depressive symptoms (Ross, 2000), and between deprivation, residential mobility and schizophrenia, major depression and substance misuse (Silver *et al*, 2002), after controlling for individual-level risk factors, were both conducted at the level of US Census tracts (average population 4000).

'Neighbourhoods' are difficult to define (Burrows & Bradshaw, 2001), and wards may be too large to detect contextual influences. This is consistent with statistically significant associations between common mental disorders and features of the built environment in small areas, after adjusting for residents' characteristics (Halpern, 1995; Weich *et al*, 2002). We had no alternative to using wards, to protect respondents' anonymity. Although residents may not equate wards with 'neighbourhoods', they are more than arbitrary administrative boundaries. Nevertheless, our findings could be consistent with substantial area-level variation at smaller spatial levels. The variance observed at the household level in this study may have been due to exposures operating at a spatial level between ward and household.

Measures of place

There is a dearth of contextual measures of place. We were also restricted in the number of area-level measures, to protect respondents' anonymity. Although the Carstairs index measures socio-economic deprivation, it may not capture aspects of the social environment with the greatest impact on mental health. We cannot exclude associations with other factors associated with place, such as residential mobility or social disorganisation (Silver *et al*, 2002).

Implications

Differences in rates of common mental disorders across electoral wards in the UK are negligible compared with the variation between individuals and households. However, these findings fail to explain why deprived persons continue to be clustered in deprived places. Geographical mobility in relation to mental health may be important but remains poorly understood. Restriction to one spatial level above household (electoral wards) and one compositional measure of place means that these findings do not wholly preclude the utility of area-based policies in reaching those at highest risk of common mental disorders (Joshi, 2001). Our findings support evidence from cross-sectional research concerning the importance of household as a determinant of mental health.

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CLINICAL IMPLICATIONS

■ There is no statistically significant geographical variation in prospectively ascertained rates of the most common mental disorders of anxiety and depression in Britain. Individual-level differences continue to dominate patterns of variance in these conditions.

■ There is substantial and statistically significant between-household variance in episode onset and maintenance and in cross-wave change in General Health Questionnaire (GHQ) score. This is not explained by the socio-economic or demographic characteristics of household members.

■ Although area-level effects appear modest, the most deprived individuals and households continue to be clustered together. Interventions delivered in specific places may still have a role in reaching individuals and households at risk of common mental disorders.

LIMITATIONS

■ Common mental disorders were assessed using a self-report symptom checklist (GHQ) rather than a standardised clinical interview.

■ There were no interval data on psychiatric morbidity between assessments.

■ Area effects were assessed at the level of electoral ward in the absence of robust evidence concerning the spatial scale at which place affects mental health.

SCOTT WEICH, MD, Division of Health in the Community, Warwick Medical School (LWMS), University of Warwick; LIZ TWIGG, PhD, Institute for the Geography of Health, Department of Geography, University of Portsmouth; GLYN LEWIS, PhD, Division of Psychiatry, University of Bristol; KELVYN JONES, PhD, School of Geographical Sciences, University of Bristol

Correspondence: Scott Weich, Division of Health in the Community, Warwick Medical School (LWMS), University of Warwick, Coventry CV4 7AL, UK. Tel: +44 (0)2476 574708; e-mail: s.weich@warwick.ac.uk

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