

Short Communication

Small-area study of the incidence of neoplasms of the brain and central nervous system among adults in the West Midlands Region, 1974–86

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Summary This small-area study of incidence of cancers of the brain and central nervous system found evidence of trend ($P = 0.02$) of cancer risk with deprivation (8% higher risk in affluent areas), but no significant association with urban – rural status. Results were not indicative of a strong geographically determined risk at small-area level.

Keywords: small-area analysis; geographical; brain and central nervous system tumours; deprivation; urban–rural

The incidence of neoplasms of the brain and central nervous system (CNS) has increased in the last 20 years in many industrialized countries, especially at older ages (Greig et al, 1990; Davis et al, 1991). Brain tumours now account for approximately 1.2% of all deaths, with causes confirmed by autopsy, and 9% of primary neoplasms in adults (Rubenstein, 1972), but little is known of their aetiology (Brownson et al, 1990; Higginson and Muir, 1992). A number of descriptive studies have shown wide geographical variation in brain cancer incidence and mortality at both international (Davis et al, 1990) and regional levels (Gardner et al, 1983; Swerdlow and dos Santos Silva, 1993), suggesting, along with evidence from migrant studies (Cohen and Moden, 1969), the potential importance of environmental and lifestyle factors. There have also been concerns about alleged local clusters of brain cancer, related, for example, to putative electromagnetic (NRPB, 1992) or chemical (Wilkinson et al, 1997) exposures.

Area-based measures of urbanization (Greenberg, 1983) or deprivation (Carstairs and Morris, 1991; Elliott, 1996) are related to the occurrence of several cancers. Up to two- to threefold variation in incidence has been reported across small areas for some cancers, including lung and stomach (Elliott, 1996). Whereas those two cancers have higher incidence in more deprived areas, for childhood brain and CNS tumours, a trend of higher incidence in more affluent areas has been reported for Scotland, with a ratio of incidence rates of 1.4 between the most affluent areas and the most deprived (McKinney et al, 1994). Studies at individual level have also shown higher mortality from brain cancer associated with higher socioeconomic status, based on both social class gradients (OPCS, 1978; Pearce and Howard, 1986; Davey-Smith et al, 1991) and census-derived variables (Leon, 1988). In contrast, a study of incident brain cancer at all ages conducted across census enumeration districts in Great Britain found little evidence for a deprivation effect (Elliott, 1996).

The aims of the present study were to explore small-area variation in the incidence of adult brain and CNS tumours at the level of electoral ward in one English health region and to examine associations with socioeconomic deprivation and urbanization as proxies for environmental and lifestyle factors. A subsidiary aim was to help interpretation of alleged local clusters of these cancers by improving our understanding of background variability in incidence.

METHODS

Registrations of cancers of the brain and CNS (benign, malignant and unspecified) in the period 1974–86 among adults aged 15–64 years in the 832 wards of the West Midlands Region of England, UK (population 3.3 million), were extracted from the national dataset held by the Small Area Health Statistics Unit (Elliott et al, 1992), using residential postcodes to locate cases. Cases without a valid postcode were excluded, although the completeness of postcodes in the West Midlands Region is high (98.7%). International Classification of Disease codes were 191, 192 and 225 (eighth and ninth revision) and 287.5, 237.6 and 237.9 (ninth revision). Malignant tumours only (191, 192, 225) were also examined separately.

The number of cancer registrations and corresponding 1981 electoral ward populations were obtained by 5 year age group and sex. Wards were classified by measures of socioeconomic deprivation and urbanization. Deprivation was measured by the Carstairs score, a composite index based on four variables from the 1981 small-area census statistics: access to a car, unemployment, overcrowding and social class of head of household (Carstairs and Morris, 1991). Wards were then grouped into quintiles of Carstairs score.

Level of urbanization was derived from a measure developed by the former Office of Population Censuses and Surveys that categorizes wards into six groups based on land-use patterns (Craig, 1988). For the present study, these six groups were collapsed into two: the first comprised wholly urban, predominately urban and mixed urban–rural wards (68% of all wards); the second, wholly rural, predominantly rural and mixed rural–urban wards (32% of all wards).

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Table 1 Estimated age – sex adjusted relative risks and 95% confidence intervals for brain and central nervous system tumours by levels of deprivation and urbanization^a

	Relative risk (95% confidence interval)	Change in deviance when terms added individually to age–sex model
Quintile of deprivation		
1 (Least deprived)	1.00	As a linear term.
2	1.10 (0.95–1.26)	Deviance = 5.51 on 1 d.f., <i>P</i> = 0.02
3	0.96 (0.84–1.11)	
4	0.98 (0.86–1.11)	
5 (Most deprived)	0.92 (0.81–1.04)	As a categorical term: deviance = 9.30 on 4 d.f., <i>P</i> = 0.05
Urbanization		
Urban	1.00	Deviance = 0.84 on 1 d.f., <i>P</i> = 0.36
Rural	0.94 (0.83–1.07)	

^aDeprivation measured by Carstairs score; urbanization derived from OPCS urban–rural classification based on land use (see text).

Statistical methods

Poisson regression was used to examine the relationship between ward-level cancer incidence and age (categorized in 5-year bands), sex and either deprivation or urban–rural status. Statistical significance was assessed using likelihood ratio tests. The numbers of cancers predicted from the Poisson models were used as expected

values in the calculation of observed – expected (O/E) ratios in subsequent analyses, as detailed below. This method was suggested by Bithell et al (1994, 1995) as a flexible model-based alternative to standardization and gives covariate adjusted expected values for each ward based on regional rates.

Three sets of analyses were then carried out to investigate and quantify possible variability in risk across wards. Further details and discussion of the methods can be found in Elliott et al (1995). First, the Potthoff – Whittinghill (1966) test was used to investigate the presence of any residual extra-Poisson variability in the regression-based O/E ratios. This tests the hypothesis of homogeneity of risk against the alternative that the relative risks are drawn from a gamma distribution. Secondly, the Smans' rank-adjacency statistic, calculated by simulation (Smans and Esteve, 1992), was used to test for the presence of possible geographical autocorrelation, i.e. when areas with relatively high (or low) risks were found close together. It should be noted that spurious autocorrelation may be generated because of variability in the size of the underlying populations at risk and hence in the stability of the rates (Smans and Esteve, 1992). Finally, in order to remove the large component of random variability arising from these small populations with unstable rates, a set of 'smoothed' risks were calculated using empirical Bayes techniques (Clayton and Kaldor, 1987). The resulting smoothed O/E ratios are a compromise between the crude (unsmoothed) estimate for each ward and the overall mean for the region, with the degree of 'shrinkage to the mean' for each ward being determined by its population size. For presentational purposes, the unsmoothed and smoothed O/E ratios were then mapped.

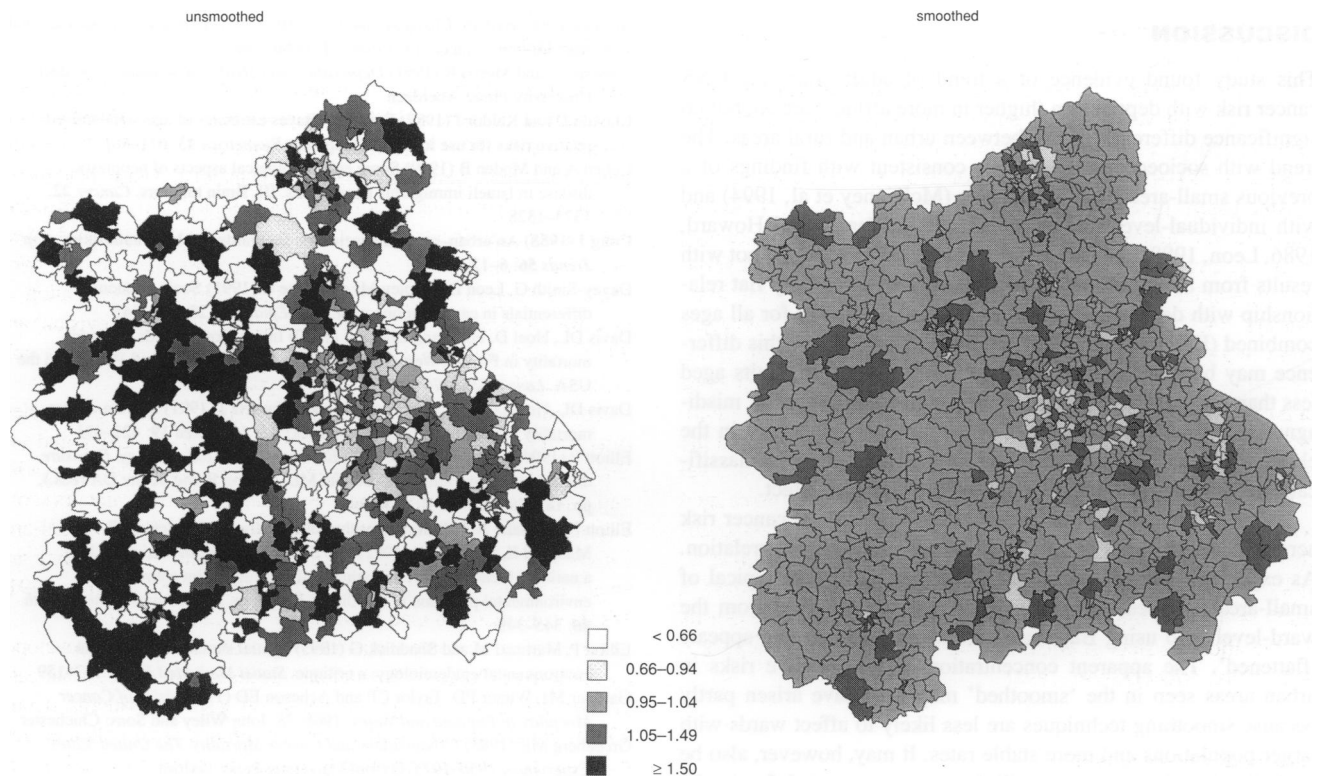


Figure 1 Age-, sex- and deprivation-adjusted relative risks of brain and central nervous system tumours for electoral wards in West Midlands region, age 15–64 years, 1974–86. Unsmoothed risks (left) and after map smoothing (right) using empirical Bayes method (see text)

RESULTS

Overall, there were 2934 postcoded registrations of malignant, benign and unspecified brain and CNS tumours (range 0–29 per ward) and 2086 malignant tumours (range 0–20 per ward) in the West Midlands Region during the period of the study. Results were broadly similar for the two diagnostic groups; only those for malignant, benign and unspecified tumours are shown here.

Results of the Poisson regression analysis, with adjustment for age, sex and either deprivation or urban–rural status, are shown in Table 1. A statistically significant ($P = 0.02$) inverse relationship was found between cancer risk and deprivation, measured as a continuous variable, and which was of borderline significance ($P = 0.05$) when deprivation was included as a categorical variable. There was an estimated 8% deficit of cases among the most deprived compared with the least deprived wards (Table 1). Urban–rural status did not add significantly to the regression model either without (Table 1) or with inclusion of deprivation (not shown) and was not included in the subsequent analyses (Table 1).

The Potthoff – Whittinghill test was suggestive of underlying heterogeneity of disease risk ($P = 0.04$), both without and with deprivation included in the calculation of expected values. There was no evidence of spatial autocorrelation using Smans test ($P = 0.18$ and $P = 0.33$ respectively).

Figure 1 shows maps of the relative risks in each ward, adjusted for age, sex and deprivation, before and after ‘smoothing’. Much of the (random) variability in the unsmoothed map is removed by smoothing, especially the high rates in the large rural areas which are based on only one or two cases. However, some low rates apparently persist in the more population-dense urban areas.

DISCUSSION

This study found evidence of a trend of adult brain and CNS cancer risk with deprivation (higher in more affluent areas), but no significance difference in risk between urban and rural areas. The trend with socioeconomic status is consistent with findings of a previous small-area study in children (McKinney et al, 1994) and with individual-level studies (OPCS, 1978; Pearce and Howard, 1986; Leon, 1988; and Davey-Smith et al, 1991) although not with results from a national study that reported an essentially flat relationship with deprivation across enumeration districts for all ages combined (Elliott, 1996). One possible explanation for this difference may be that the present study was restricted to adults aged less than 65 years to minimize the well-known problems of misdiagnosis and misclassification of these tumours, especially in the elderly (Annegers et al, 1980; Rees et al, 1993). Such misclassifications (if haphazard) would tend to dilute any true effect.

We only found weak evidence for heterogeneity of cancer risk across electoral wards and no evidence of spatial autocorrelation. As expected, after allowing for the sparseness of data typical of small-area analyses, extreme O/E ratios were removed from the ward-level map using Bayesian techniques, and the map appears ‘flattened’. The apparent concentration of low relative risks in urban areas seen in the ‘smoothed’ map may have arisen partly because smoothing techniques are less likely to affect wards with larger populations and more stable rates. It may, however, also be an indication that there are still factors unaccounted for in the model that vary geographically, albeit weakly. Further discussion of the problems in interpreting disease maps such as this can be

found elsewhere (Smans and Esteve, 1992; Elliott et al, 1995; Olsen et al, 1996).

In summary, while a trend of higher risk in more affluent areas was found, which is consistent with the results of individual-level studies, overall there was no indication of a strong geographically determined risk for brain and CNS tumours at the small-area level.

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