Common mental disorders and the built environment in Santiago, Chile*

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Background There is growing research interest in the influence of the built environment on mental disorders.

Aims To estimate the variation in the prevalence of common mental disorders attributable to individuals and the built environment of geographical sectors where they live.

Method A sample of 3870 adults (response rate 90%) clustered in 248 geographical sectors participated in a household cross-sectional survey in Santiago, Chile. Independently rated contextual measures of the built environment were obtained. The Clinical Interview Schedule was used to estimate the prevalence of common mental disorders.

Results There was a significant association between the quality of the built environment of small geographical sectors and the presence of common mental disorders among its residents. The better the quality of the built environment, the lower the scores for psychiatric symptoms; however, only a small proportion of the variation in common mental disorder existed at sector level, after adjusting for individual factors.

Conclusions Findings from our study, using a contextual assessment of the quality of the built environment and multilevel modelling in the analysis, suggest these associations may be more marked in non-Western settings with more homogeneous geographical sectors.

Declaration of interest None.

There is growing interest in investigating whether contextual variables, such as those representing the built environment, can influence the prevalence of common mental disorders after accounting for individual variables (Weich, 2005). The built environment encompasses all those aspects of our habitat that are created or modified by people, such as homes, schools, parks and roads (Srinivasan et al, 2003). Up to now most studies of the built environment and mental illness have used indirect environmental measures such as individuals’ perceptions (Ellaway et al, 2001) or aggregated data (Ross, 2000; Weich et al, 2003). A few studies have assessed the built environment directly and independently (Weich et al, 2002; further details available from the authors), finding little or no variation in prevalence of common mental disorders across small or medium-sized areas (Weich, 2005).

We are unaware of any other Latin American study assessing the contextual effect of the built environment directly and using multilevel models to investigate its association with mental illness. We tested the hypothesis that contextual measures reflecting the quality of the built environment in Santiago, Chile would be associated with common mental disorders independent of individuals’ characteristics.

METHOD

Santiago, the capital of Chile, has a population of approximately 6 million people, representing 42% of the total Chilean population. Although Chile is considered to be a middle-income country, with a gross per capita annual income of £2600, it is one of the ten most unequal countries in the world in terms of income (World Bank, 2001). The city is geographically extended, with a mean population density of 392 persons per square kilometre. Geographically the city is neatly compartmentalised according to socio-economic groups, with wealthier people living mainly in the eastern suburbs and the poorest in the southern and northern fringes. Most of the city has basic amenities, including electricity, sanitation and drinkable water, but there are visible differences between more and less affluent sectors. Houses in the wealthier areas are bigger and of better quality. Facilities including better roads and pavements (sidewalks), green areas, overall cleanliness and abundant shops are noticeable. However, crime seems to be present in all sectors, and crime reported to police is especially common in wealthier sectors of the city because criminals target these areas and people report more incidents for various reasons, including private insurance claims.

Sampling strategy

A cross-sectional survey was conducted between 1996 and 1998. The sampling framework was the adult population, restricted to ages 16–64 years, living in private households in the Greater Santiago area. The sampling strategy involved a three-stage design, which included all the 35 boroughs of Greater Santiago, 248 sectors and 4300 households randomly selected with a probability proportional to the size of the sampling units. The number of households within each sector varied from 26 to 5. One person per household was chosen randomly for interview using Kish tables (Kish, 1965). Individuals from sectors with fewer than five observations were excluded from this analysis. Responses were obtained from 3870 households (response rate 90%). Further details of the sampling design can be found elsewhere (Araya et al, 2001).

Mental health, social and demographic questionnaire

Psychiatric symptoms were assessed with the Revised Clinical Interview Schedule (CIS–R; Lewis et al, 1992), a structured and detailed psychiatric interview used extensively in primary care and community studies in Chile and elsewhere. This interview has 14 items assessing the severity of the most common psychiatric symptoms. Each item is given a score, which can then be summed to yield a total score. This continuous measure, reflecting psychiatric symptom severity, was used as our main outcome. The mean weighted $x$ across all

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16 sections of the CIS–R was 0.87
(s.d. = 0.08). The validity and reliability of
the CIS–R are comparable with the Composite
International Diagnostic Interview
(CIDI; Lewis et al., 1992; Andrews & Peters,
1998; Brugha et al., 2005).

The gender and age of respondents
were recorded. Individuals also answered
questions regarding their socio-economic
status including marital status, educational
level and monthly per capita income. The
latter was estimated as the sum of net
monthly salaries and other income (pensions,
dividends, interests or rents) contributed
by each household member, divided by
the number of residents regardless of age.
Interviewers rated the quality of housing
through visual inspection as luxurious,
good, average, poor or very poor. In order
to do this, interviewers used the same
criteria and received the same training as
for the Chilean national census. The pres-
ence of a self-reported physical disease
was ascertained from the response to an
open-ended question: ‘Do you suffer from
any physical problem or disability at
present?’ Two independent physicians
assessed if the physical problem would
require medical attention, in which case
they classified it according to the bodily
system involved. The self-reported number
of friends or relatives who could provide
emotional or practical support if needed
was determined with a single, open-ended
question. The number of alcohol units con-
sumed daily was entered as a continuous
variable. All violent crimes reported to the
local police station in each one of the sec-
tors were added and this figure was divided
by the population of the sector to create an
index of violence. The following borough
variables were also included: education
and health budget per capita, and number
of social organisations divided per popu-
lation size in the borough.

**Built Environment Assessment Tool**

The Built Environment Assessment Tool
(PEAT; Dunstan et al., 2005) used to measure area
characteristics in Wales. Some additional
items were included because they were
thought to be important locally, such as
the presence of stray dogs or bad odours.
The final instrument (available from the
authors) included items relating to the
following characteristics, with the number
of items in parentheses: public lighting (2),
state of roads (6), sidewalks (4), public
green areas (5), green elements on side-
walks and front gardens (4), dirtiness (1),
traffic and noise (2), bad odours (1), general
maintenance of properties (1), general use
of the sector (4), empty sites (2), external
facilities: primary and secondary schools,
of the sector (4), empty sites (2), external
facilities: primary and secondary schools,
(1), presence of homeless people (1), presence
of stray dogs (1), access to properties (2),
balconies (1), street signs (1), public transport
(4), security and safety devices (6), and a list of facilities including:

- **Essential facilities**: primary and secondary schools, other type of schools,
creches, primary care clinics, hospitals and private clinics;
- **Leisure facilities**: public gymnasium, swimming pool, football pitch, sports
club, cultural centre, library, community centre, corner shops, pharmacies,
cinemas, theatre, restaurants, coffee shops, bars;
- **Other facilities**: petrol station, petrol station shops and kiosks.

Some of the items required dichoto-
mous ratings, for example the presence of
lamp-posts; others had a range of possible
values that could be ordered from high to
low, such as the level of maintenance of
front gardens. Ratings for all items were
converted into a score between 0 and 1,
with the value of 1 always representing
either a more desirable feature (e.g. cleaner
roads) or more of that particular item (e.g.
more essential facilities). For example, the
test for ‘beauty of front garden’ was initially
coded 1, 2 or 3 and subsequently recoded as
0, 0.5 or 1, with 1 representing ‘very
beautiful’ front gardens. The presence of
any of the listed facilities in the sector was
rated as ‘1’ and individual scores were
summed to generate three facility indices.

Statistical analyses

**Composite score for the BEAT scale**

All variables with 95% or more of respon-
dents in one category were eliminated.
Subsequently we performed factor analysis
with varimax rotation of all the remaining
items to assess if and how these items
loaded into common factors. Cronbach’s α
was estimated for all the items in the scale
and for the items within each one of the
newly derived factors after the factor analy-
sis. All variables were entered, including
the facilities indices, into the factor analysis
model and those with loadings lower than
individual and finally borough-level variables to the model and to note changes in the components of variance and coefficients for the sector-level factors. We investigated whether there were any differential associations between factor 1 and CIS–R for categories of selected individual variables by fitting appropriate interaction terms in the regression models.

RESULTS

Development of the Built Environment Assessment Tool

Twenty-seven items were removed from the original list for two reasons: 14 did not show sufficient ability to discriminate (95% or more of the answers fell into one category) and the remainder had loadings below 0.4 after varimax rotation or high uniqueness values so that they did not fit well with any of the factors. Among the items left out were several questions on the type of roads, external beautification of properties, parking on sidewalks, predominant type of properties, protection on balconies, security fences, presence of taxis, bad odours, the size of green areas on sidewalks, and presence of vagrants. Twenty-five items remained, including three representing the sum of the list of facilities. These were grouped into four main factors:

(a) general quality of the area;
(b) facilities, noise and traffic in the area;
(c) public green areas;
(d) empty sites.

These four factors all had eigenvalues over 1 and together explained 90% of the total variance (Table 1). The mean

<table>
<thead>
<tr>
<th>Items</th>
<th>Factor number</th>
<th>Factor 1 General quality</th>
<th>Factor 2 Facilities</th>
<th>Factor 3 Green areas</th>
<th>Factor 4 Empty sites</th>
</tr>
</thead>
<tbody>
<tr>
<td>Proportion variance, %</td>
<td>50.1</td>
<td>20.3</td>
<td>12.6</td>
<td>8.2</td>
<td></td>
</tr>
<tr>
<td>Eigenvalues</td>
<td>6.2</td>
<td>2.5</td>
<td>1.5</td>
<td>1.0</td>
<td></td>
</tr>
<tr>
<td>Loadings presented after varimax rotation</td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>Width of sidewalks</td>
<td>1</td>
<td>0.45</td>
<td>-0.24</td>
<td>-0.18</td>
<td>0.29</td>
</tr>
<tr>
<td>General maintenance of sidewalks</td>
<td>1</td>
<td>0.69</td>
<td>-0.012</td>
<td>0.07</td>
<td>-0.04</td>
</tr>
<tr>
<td>Additional features on sidewalks</td>
<td>1</td>
<td>0.71</td>
<td>-0.02</td>
<td>0.07</td>
<td>0.12</td>
</tr>
<tr>
<td>State of front gardens</td>
<td>1</td>
<td>0.70</td>
<td>0.03</td>
<td>0.03</td>
<td>-0.01</td>
</tr>
<tr>
<td>Trees on sidewalks</td>
<td>1</td>
<td>0.54</td>
<td>-0.07</td>
<td>-0.01</td>
<td>0.16</td>
</tr>
<tr>
<td>Size of trees on sidewalks</td>
<td>1</td>
<td>0.44</td>
<td>-0.17</td>
<td>-0.13</td>
<td>0.16</td>
</tr>
<tr>
<td>Green area on sidewalk</td>
<td>1</td>
<td>0.74</td>
<td>-0.15</td>
<td>-0.04</td>
<td>0.27</td>
</tr>
<tr>
<td>Dirtiness of street</td>
<td>1</td>
<td>0.63</td>
<td>-0.00</td>
<td>0.18</td>
<td>-0.08</td>
</tr>
<tr>
<td>General maintenance of properties</td>
<td>1</td>
<td>0.67</td>
<td>-0.11</td>
<td>0.17</td>
<td>-0.07</td>
</tr>
<tr>
<td>Type of properties</td>
<td>1</td>
<td>0.41</td>
<td>-0.14</td>
<td>0.08</td>
<td>0.18</td>
</tr>
<tr>
<td>Stray dogs</td>
<td>1</td>
<td>0.57</td>
<td>-0.08</td>
<td>0.19</td>
<td>0.14</td>
</tr>
<tr>
<td>Signs for orientation</td>
<td>1</td>
<td>0.41</td>
<td>-0.37</td>
<td>-0.10</td>
<td>0.05</td>
</tr>
<tr>
<td>Other public signs</td>
<td>1</td>
<td>0.67</td>
<td>-0.16</td>
<td>0.07</td>
<td>0.12</td>
</tr>
<tr>
<td>Security badges on houses</td>
<td>1</td>
<td>0.41</td>
<td>-0.29</td>
<td>-0.24</td>
<td>0.25</td>
</tr>
<tr>
<td>Guards</td>
<td>1</td>
<td>0.42</td>
<td>-0.24</td>
<td>-0.08</td>
<td>0.33</td>
</tr>
<tr>
<td>Level of traffic</td>
<td>2</td>
<td>-0.05</td>
<td>0.80</td>
<td>0.01</td>
<td>-0.17</td>
</tr>
<tr>
<td>Noise of traffic</td>
<td>2</td>
<td>0.02</td>
<td>0.71</td>
<td>-0.02</td>
<td>-0.06</td>
</tr>
<tr>
<td>Bus stop</td>
<td>2</td>
<td>0.08</td>
<td>0.54</td>
<td>-0.18</td>
<td>0.09</td>
</tr>
<tr>
<td>Essential facilities</td>
<td>2</td>
<td>0.27</td>
<td>0.41</td>
<td>-0.18</td>
<td>-0.03</td>
</tr>
<tr>
<td>Leisure facilities</td>
<td>2</td>
<td>0.12</td>
<td>0.52</td>
<td>-0.15</td>
<td>-0.01</td>
</tr>
<tr>
<td>Other facilities</td>
<td>2</td>
<td>0.14</td>
<td>0.47</td>
<td>-0.11</td>
<td>0.16</td>
</tr>
<tr>
<td>Public green areas</td>
<td>3</td>
<td>-0.11</td>
<td>-0.12</td>
<td>0.78</td>
<td>0.12</td>
</tr>
<tr>
<td>State of public green areas</td>
<td>3</td>
<td>0.28</td>
<td>-0.10</td>
<td>0.74</td>
<td>0.09</td>
</tr>
<tr>
<td>Presence of empty sites</td>
<td>4</td>
<td>0.08</td>
<td>0.03</td>
<td>0.08</td>
<td>0.80</td>
</tr>
<tr>
<td>Empty sites occupied illegally</td>
<td>4</td>
<td>0.11</td>
<td>0.07</td>
<td>0.08</td>
<td>0.82</td>
</tr>
</tbody>
</table>
Cronbach’s $z$ for the items in the scale was 0.87. There were small differences between these values of $z$ and those generated using only the items within each factor. Lower $z$ values for the items contained in factors with lower eigenvalues are explained because the first factor with the highest eigenvalue contains the most items, thus making the greatest contribution to the overall variation in scores. Kappa coefficients for items between pairs of interviewers fluctuated from 0.69 to 0.92, with 78% of the estimated $\kappa$ coefficients above 0.85 and full agreement for 70% of the items. Simply summing items to get a total factor score assumes equal weighting of each item and that ‘non-loading’ items are not important. For this reason we compared weighted (according to eigenvalues) and unweighted scores for each of the factors. We found high correlations between these two different ways of scoring the factors (correlation coefficients for factors 1, 2, 3 and 4 were 0.99, 0.87, 0.92 and 0.88 respectively), and therefore we decided to use the simple, unweighted scores in all analyses. Correlation coefficients between the four sector-level factors are presented in Table 2.

**Characteristics of the sample surveyed**

A total of 3870 interviews were completed to give a response rate of 90%. These 3870 individuals clustered into 248 sectors and 35 boroughs. A total of 488 individuals (12.6%) were excluded because of missing data or because they lived in a sector with fewer than five respondents, leaving 3382 observations from 210 sectors within 31 boroughs for analysis. Excluded individuals were no different to those included in terms of age ($P=0.56$), gender ($P=0.17$) or marital status ($P=0.59$), but had lower median income (in Chilean pesos, CLP62500 v. CLP100000, $P<0.0001$), were less likely to be educated to university level (20 v. 36%, $P<0.001$), were more likely to live in very poor or poor quality housing (22 v. 15%, $P<0.001$), have fewer supportive individuals (3.7 v. 4.2, $P=0.01$) and have lower alcohol consumption (1.5 v. 1.8, $P=0.005$). Excluded individuals also had higher mean CIS–R scores (8.5 v. 7.2, $P<0.001$). There was no evidence that excluded sectors were any different from those included in terms of the four factors generated from the BEAT scores or violent incidents reported to police. The number of individuals per sector in the final data-set ranged from 5 to 26 and the number of sectors per borough from 2 to 26. Characteristics of the individuals, sectors and boroughs are presented in Table 2.

**Variance components null model for common mental disorder**

Mean CIS–R score for the total sample was 7.19 (s.d.=8.00, range 0–49). We estimated that approximately 5.6% (95% CI 1.8–9.4) of the residual variation in total CIS–R score lies at the borough level, 3.8% (95% CI 1.8–5.7) at the sector level and 90.6% (95% CI 86.2–95.1) at the individual level (Table 4). In view of the non-normal distribution of CIS–R scores we also undertook all multilevel modelling using log-transformed scores and the results in all these models were almost identical to those presented here (further details available from the authors).

**Effect of including individual, sector and borough characteristics in the model**

Sector-, individual- and borough-level fixed effects were added to the null model in a cumulative manner (Table 4). Inclusion of the sector-level exposures reduced the total residual variation. The estimated percentage of residual variation at borough level decreased to 0.13%, whereas the percentage residual variation at sector (3.55%) and individual (96.32%) levels remained similar and increased respectively. Additional inclusion of individual-level variables reduced the overall variation further, with none of the residual variation now explained at the borough level. Estimates remained unchanged after the addition of borough-level variables. Estimated associations between sector-level factors and mental health were therefore based on a simpler, two-level model that included only sector (level 2) and individual (level 1) variables.

**Effect of neighbourhood quality on common mental disorders**

Table 5 shows the crude and adjusted associations between each of the sector-level factors (rescaled so that the possible range for each is 0–10) and CIS–R total score. After adjusting for other sector-level and individual-level variables, factor 1 (overall quality of the built environment) was inversely associated with total CIS–R score: that is, there was strong evidence that individuals living in sectors with more desirable features such as better roads or more green areas had better mental health, after taking into account individual characteristics. There was also a significant association with factor 4 in the adjusted model only, but this was in the opposite direction; higher factor scores were associated with higher CIS–R scores.

We tested interactions between factor 1 and the following individual variables: gender, income and education. The only significant interaction was for gender and factor 1 (–0.32, 95% CI −0.58 to −0.05, $P=0.02$) in which male respondents living in less desirable areas had significantly lower CIS–R scores than female respondents. No significant interaction was found for income (0.004, 95% CI −0.001 to 0.008, $P=0.10$) or education (secondary, 0.23, 95% CI −0.26 to 0.72; university, 0.16, 95% CI −0.35 to 0.68; overall $\chi^2=0.91$, d.f.=2, $P=0.63$).

Although our primary interest was to investigate the association of these factors with mental health, which was best represented by the continuous distribution in CIS–R total scores, we also explored associations with the most common ICD–10 disorders (World Health Organization, 1992), anxiety and depressive disorders, using logistic regression models. There were 154 (46.6%) cases of depression and 309 (9.1%) of anxiety. There was no evidence of any association with depression for factors 1, 2, or 4 (factor 1, OR=1.00,
Table 3 Characteristics of individuals, sectors and boroughs in the sample

<table>
<thead>
<tr>
<th>Variable</th>
<th>Individuals (n = 3382)</th>
</tr>
</thead>
<tbody>
<tr>
<td>Age, years: mean (s.d.)</td>
<td>36.9 (13.8)</td>
</tr>
</tbody>
</table>
| Income per capita in household: median (IQR)
1. Chilean pesos. | 100 000 (50 000–250 000) |
| Number of supportive people: median (IQR) | 3 (2–5) |
| Units of alcohol consumed daily: mean (s.d.) | 1.8 (2.2) |
| Gender, n (%) |  
| Male | 1358 (40) |
| Female | 2024 (60) |
| Self-rated presence of disease, n (%) |  
| No | 2782 (82) |
| Yes | 600 (18) |
| Education level, n (%) |  
| Primary | 559 (17) |
| Secondary | 1607 (48) |
| University | 1216 (36) |
| Marital status, n (%) |  
| Married/cohabiting | 1904 (56) |
| Widowed | 127 (4) |
| Separated | 259 (8) |
| Single | 1092 (32) |
| Housing type, n (%) |  
| Very poor | 135 (4) |
| Poor | 357 (11) |
| Average | 1474 (44) |
| Good | 1188 (35) |
| Luxurious | 228 (7) |
| Sectors (n = 210) |  
| Scores: mean (s.d.) |  
| Factor 1: General quality (possible range 0–15) | 8.97 (2.75) |
| Factor 2: Facilities (possible range 0–6) | 2.43 (0.55) |
| Factor 3: Green areas (possible range 0–2) | 0.97 (0.77) |
| Factor 4: Empty sites (possible range 0–2) | 1.66 (0.51) |
| Episodes of violent crime reported to local police: median (IQR) | 1.8 (0.7–5.5) |

Boroughs (n = 31)

| Scores: mean (s.d.) |  
| Education budget per capita: median (IQR)
1. Chilean pesos. | 146 404 (115 264–224 365) |
| Health budget per capita: median (IQR)
1. Chilean pesos. | 17 037 (14 558–22 536) |
| Number of social organisations per population: mean (s.d.) | 0.0018 (0.0013) |

IQR, interquartile range.
1. Chilean pesos.

Table 4 Components of variance in total Revised Clinical Interview Schedule score (as a continuous variable) at the individual, sector and borough level: multilevel modelling

| Variance (s.e.) | Null model | Model 1 (null+sector variables)
1. Sector variables were age, gender, presence of disease, income, education, marital status, housing type, number of supportive individuals and alcohol use. Model 2 also included the sector-level variable episodes of violent crime reported to local police. | Model 2 (model 1+individual variables)
2. Individual variables were age, gender, presence of disease, income, education, marital status, housing type, number of supportive individuals and alcohol use. Model 2 also included the sector-level variable episodes of violent crime reported to local police. | Model 3 (model 2+borough variables)
3. Borough variables were education budget per capita, health budget per capita and number of social organisations. |
<table>
<thead>
<tr>
<th></th>
<th></th>
<th></th>
<th></th>
</tr>
</thead>
<tbody>
<tr>
<td>Level 3 (borough)</td>
<td>3.59 (1.25)</td>
<td>0.08 (0.21)</td>
<td>0.00 (0.00)</td>
</tr>
<tr>
<td>Level 2 (sector)</td>
<td>2.42 (0.64)</td>
<td>2.15 (0.60)</td>
<td>0.51 (0.36)</td>
</tr>
<tr>
<td>Level 1 (individual)</td>
<td>58.28 (1.46)</td>
<td>58.33 (1.46)</td>
<td>51.29 (1.28)</td>
</tr>
</tbody>
</table>

DISCUSSION

This is the first large population-based study of common mental disorders and the built environment of small geographical sectors of a Latin American city using a directly assessed contextual measure and multilevel modelling in the analysis. We found strong evidence of an association between the quality of the built environment in these sectors and common mental disorders, before and after adjusting for individual variables. However, in line with previous reports, the contribution of these sectors to the total variance in common mental disorders was small and most of it was explained by individual factors. None the less, these results represent probably some of the most persuasive evidence found so far to establish an association between the quality of our surrounding built environment and the presence of common mental disorders.

The Built Environment Assessment Tool

We developed and tested a quick and reliable method to assess the built environment using a walk-through method. The great majority of the items reflected the built environment of small geographical sectors of a Latin American city using a directly assessed contextual measure and multilevel modelling in the analysis. We found strong evidence of an association between the quality of the built environment in these sectors and common mental disorders, before and after adjusting for individual variables. However, in line with previous reports, the contribution of these sectors to the total variance in common mental disorders was small and most of it was explained by individual factors. None the less, these results represent probably some of the most persuasive evidence found so far to establish an association between the quality of our surrounding built environment and the presence of common mental disorders.
environment, but there were a few – such as the presence of stray dogs – that represented the observable residential environment rather than something built. We had previous experience with developing a similar instrument for a study in South Wales (Dunstan et al., 2005) and we studied carefully other similar instruments (Sampson et al., 1997; Cohen et al., 2000; Weich et al., 2002; Hembree et al., 2005). Our measure showed good interrater reliability and internal consistency; in the absence of a gold standard, however, it is difficult to assess its criterion validity. Overall, the psychometrics of this tool are comparable with those found for the two similar tools developed in the UK (Weich et al., 2001; Dunstan et al., 2005).

**Variance in mental disorders according to geographical aggregation**

Studies in Western countries using multi-level models to estimate area-level variation in common mental disorders have found little or no variation at higher levels of aggregation, after accounting for individual differences (Weich, 2005). The contribution of smaller area effects to the total variance in common mental disorders usually fluctuates between 0.5% and 4% before adjusting for individuals’ characteristics, and drops to levels between 0% and 1% after adjustment (Weich, 2005). Our findings are in keeping with the higher end of previous estimates; we found that 3.8% of the variance in common mental disorders was explained at the small sector level in the unadjusted models, reducing to nearly 1% in the adjusted models.

As eloquently argued by Weich (2005), there may be a number of reasons to explain this lack of positive findings. For instance, sectors that are large or heterogeneous tend to yield negative results. However, in a previous UK study in South Wales using small geographical units (postcode with approximately 150 people) we also found little variation at this level (further details available from the authors). It is possible that using geographical units identified on the basis of an arbitrary geographic classification, which might not reflect neighbourhood unity, may influence the results. In this respect, Reijneveld et al. (2000) found that the clustering of common mental disorders was higher at neighbourhood level (sectors with similar types of building delineated by natural boundaries) than at postcode level using arbitrary geographical boundaries. We tried to deal with both of these possibilities, so we used small geographical sectors of approximately 300 people that were sufficiently homogeneous in terms of their neighbourhood. Our study did not measure the outcome (CIS–R) aggregated at household level and so it is possible that some of the variance found at higher (sector) or lower (individual) levels might reflect variance present at household level.

In our previous study in Wales we found that 37% (95% CI 27–49) of variance existed at household level (further details available from the authors). Although the CIS–R variance at borough level appears greater than at sector level in Table 4, the confidence intervals of these estimates (5.6%, 95% CI 1.8–9.4% at the borough level and 3.8%, 95% CI 1.8–5.7 at the sector level) show that one cannot reach this conclusion. More importantly, once adjustment for other variables are introduced this borough variance comes close to nil but the sector variance remains only slightly attenuated. Yet the variance we found at borough level in adjusted models is considerable in comparison with other studies. The most likely explanation is that Santiago, like other cities in Latin America, is quite compartmentalised in terms of the quality and socio-economic status of the

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**Table 5**  Association between sector-level factors according to Built Environment Assessment Tool and Revised Clinical Interview Schedule total score as a continuous variable: two-level regression models

<table>
<thead>
<tr>
<th></th>
<th>Unadjusted</th>
<th>Adjusted1</th>
</tr>
</thead>
<tbody>
<tr>
<td><strong>Beta (95% CI)</strong></td>
<td></td>
<td></td>
</tr>
<tr>
<td><strong>Sector level</strong></td>
<td></td>
<td></td>
</tr>
<tr>
<td>Factor 1</td>
<td>0.96 (-1.14 to -0.77)</td>
<td>-0.30 (-0.49 to -0.11)</td>
</tr>
<tr>
<td>Factor 2</td>
<td>-0.18 (-0.73 to 0.37)</td>
<td>-0.04 (-0.38 to 0.31)</td>
</tr>
<tr>
<td>Factor 3</td>
<td>-0.14 (-0.25 to -0.03)</td>
<td>-0.01 (-0.09 to 0.06)</td>
</tr>
<tr>
<td>Factor 4</td>
<td>0.12 (-0.04 to 0.28)</td>
<td>0.17 (0.06 to 0.28)</td>
</tr>
<tr>
<td><strong>Individual level</strong></td>
<td></td>
<td></td>
</tr>
<tr>
<td>Age</td>
<td>-0.02 (-0.03 to 0.004)</td>
<td>-0.08 (-0.10 to -0.06)</td>
</tr>
<tr>
<td>Female gender</td>
<td>3.03 (2.50 to 3.55)</td>
<td>3.16 (2.61 to 3.70)</td>
</tr>
<tr>
<td>Income2</td>
<td>-0.03 (-0.04 to -0.03)</td>
<td>-0.01 (-0.02 to -0.005)</td>
</tr>
<tr>
<td>Education</td>
<td></td>
<td></td>
</tr>
<tr>
<td>Primary Reference</td>
<td>-2.24 (-3.00 to -1.49)</td>
<td>-1.42 (-2.16 to -0.67)</td>
</tr>
<tr>
<td>Secondary Reference</td>
<td>-4.41 (-5.22 to -3.59)</td>
<td>-1.86 (-2.74 to -0.97)</td>
</tr>
<tr>
<td>University</td>
<td>0.02 (0.04 to 0.03)</td>
<td>0.01 (0.02 to 0.005)</td>
</tr>
<tr>
<td><strong>Marital status</strong></td>
<td></td>
<td></td>
</tr>
<tr>
<td>Married Reference</td>
<td>0.37 (-1.02 to 1.77)</td>
<td>0.29 (-1.06 to 1.64)</td>
</tr>
<tr>
<td>Separated</td>
<td>2.36 (1.36 to 3.37)</td>
<td>1.53 (0.58 to 2.47)</td>
</tr>
<tr>
<td>Single</td>
<td>-0.60 (-1.18 to -0.02)</td>
<td>-1.26 (-1.90 to -0.61)</td>
</tr>
<tr>
<td><strong>Housing</strong></td>
<td></td>
<td></td>
</tr>
<tr>
<td>Very poor Reference</td>
<td>-2.06 (-3.60 to -0.53)</td>
<td>-1.58 (-3.02 to -0.14)</td>
</tr>
<tr>
<td>Poor</td>
<td>-3.79 (-5.16 to -2.43)</td>
<td>-2.51 (-3.81 to -1.22)</td>
</tr>
<tr>
<td>Average</td>
<td>-6.40 (-7.80 to -5.01)</td>
<td>-3.37 (-4.76 to -1.98)</td>
</tr>
<tr>
<td>Good</td>
<td>-8.98 (-10.67 to -7.30)</td>
<td>-4.25 (-6.03 to -2.48)</td>
</tr>
<tr>
<td>Luxurious</td>
<td>-0.33 (-0.40 to -0.27)</td>
<td>-0.24 (-0.30 to -0.18)</td>
</tr>
<tr>
<td>Social support (no. of supportive people)</td>
<td>0.11 (-0.01 to 0.23)</td>
<td>0.39 (0.27 to 0.52)</td>
</tr>
<tr>
<td>Alcohol use</td>
<td>3.88 (3.20 to 4.56)</td>
<td>3.75 (3.08 to 4.41)</td>
</tr>
</tbody>
</table>

1. Adjusted for other sector-level factors and sector-level variable episodes of violent crime reported to local police, and individual-level variables age, gender, presence of disease, income, education, marital status, housing type, number of supportive individuals and alcohol use.
2. Per 10 000 pesos.
3. Value from Wald test.
geographical areas, with little variation within but more variation between boroughs.

Previous studies have been criticised because they tend to rely entirely on brief psychiatric self-reported questionnaires to measure the outcome. Our study used a detailed structured psychiatric interview to overcome this limitation. So, as it stands, we have to conclude that there seems to be little variation in prevalence of common mental disorder explained at area level and much of this variance resides at individual level. However, even if small sectors contribute little to this overall variance, is it still possible that some features of these areas may be associated with common mental disorders?

Quality of residential environment and common mental disorder

There have been only a handful of mental health studies that have used truly contextual and independent measures of the built environment throughout the world. In the UK there have been only two such studies. Weich et al (2002) found a significant association between the prevalence of depression and properties with predominantly deck access (OR = 1.28, 95% CI 1.03–1.58) and of recent construction (OR = 1.43, 95% CI 1.06–1.91). It is worth noting that this was a cross-sectional study in which no multilevel modelling was used in the analysis (Weich et al., 2002). Using a similar contextual assessment of the built environment as in the study we report here and multilevel modelling in the analysis we did not find any significant association between the total score of an index depicting quality of the residential environment and the prevalence of common mental disorders in our study in South Wales (further details available from the authors). It must also be borne in mind that both Weich et al. (2002) and our previous study used brief questionnaire measures of mental disorder and studied smaller samples than in this study (Weich et al., 76 sectors, n = 1887; our previous study, 51 sectors, n = 1500). A larger number of sectors could help to improve the accuracy of the estimates and provide greater power to test smaller effects.

We found strong evidence of an association (P ≤ 0.05) between two factors of our index of quality of the built environment (BEAT) and common mental disorders, after adjusting for individual differences. These factors represented almost two-thirds of the total variance in the quality of the built environment and thus one can confidently conclude that they are good indicators of the built environment in the city of Santiago. We used a similar method as in the South Wales study (REAT; Dunst et al., 2005), but there were some differences that might help explain the discrepant results. The REAT assessed mainly the more private built environment such as houses, gardens or housing density. The BEAT assessed extensively other aspects of the built environment such as roads, pavements and public facilities. The BEAT provided a total score reflecting the quality of the residential environment, whereas we used four factors with their corresponding individual scores. Although it may seem intuitive that a better built environment might help us feel better, the precise mechanism by which the built environment influences our mental health is still a matter of conjecture.

Why did only two factors show significant associations in our study? The first factor represented the largest proportion of the variance and it was the most comprehensive indicator of the quality of the neighbourhood and built environment. Although the relative contribution of this factor to change in CIS–R score is approximately ten times smaller than that associated with individual variables such as being female, it is a factor amenable to change and it is widely spread. Interestingly, the only significant interaction across levels showed that women were more affected (higher CIS–R scores) than men when living in less desirable areas. This would be in keeping with our hypotheses because the women – especially those who did not work outside the home – probably spent more time in the areas studied than men and were therefore more exposed.

We found, rather surprisingly, that factor 4 (empty sites) was associated in the opposite direction: fewer sites were associated with higher CIS–R scores. However, factor 4 was not a key indicator of the area environment, contributing only 8% of the variance in our factor analysis, and in the unadjusted model (see Table 5) this association was not significant at a 5% level. Our assumption was that fewer empty sites, especially if they were unoccupied, would be a good feature of the sector; however, it is possible that our assumptions were baseless and that empty sites in Santiago may not represent abandoned, derelict places where rubbish accumulates, as in other settings. We expected that factor 2, representing 20% of the total variance, would be significantly associated with CIS–R scores. However, this factor was a rare combination of essential and leisure facilities and noise and traffic in the area. Our assumption here was that an increased number of facilities would represent an asset for the locality, but it may be that more facilities bring more noise and traffic to the area and that this is more important. Nevertheless, overall it is reassuring that the strongest and clearest association is for the best and most comprehensive indicator of the quality of the built environment. When we explored associations of these factors with ICD–10 categorical disorders the results were puzzling. We found that there was no association between these disorders and factor 1, representing the overall quality of the neighbourhood. Even more surprisingly, individuals who were depressed were more likely to live in areas with more public green areas, an association that we did not find when using CIS–R total scores. More in keeping with the other results, individuals living in areas with fewer empty sites were less likely to have an anxiety disorder, an association that we found for CIS–R total scores but in the opposite direction. It is difficult to find a reasonable explanation for these disparate findings, especially for those related to depressive disorders. However, our interest was to focus on population changes in mental health (symptom scores) rather than concentrate on specific subgroups, mainly because the former approach would be more informative for public health decision-makers (Rose, 1993).

Santiago is fairly well compartmentalised according to socio-economic grouping. Wealthy people live in areas completely removed from the areas where poorer people live, something not always found in UK cities with a much more mixed socio-economic distribution within geographical sectors. This clear and distinct geographical distribution might have helped reduce 'contamination' and accentuated the differences between the sectors selected in our clustered sampling strategy. We selected the sectors in our sample to represent an adequate spread of neighbourhood deprivation, so we expected this would ensure an adequate spread of residential quality. We think that a drop of one point in the total CIS–R score attributable to living in the sectors with better built environment quality is a meaningful change, bearing in mind the large proportion of people who might potentially benefit from interventions.
to reduce this difference. When a common threshold of common mental disorder case-ness with the CIS–R (≥ 12) is used, those living in areas with better built environment are approximately 20% less likely to meet caseness criteria than those living in areas with poorer built environments. Lev-enthal & Brooks-Gunn (2003) found that families who moved from a very poor neighbourhood to a non-poor neighbour-hood showed better mental health than control families who did not move. A similar issue related to mobility is whether or not individuals with poorer mental health may selectively move to more deteriorated areas rather than poorer areas making individuals unhappier (causation v. selection). Unfortu-nately the design of our study does not allow adequate testing of this theory, and the stability of residence was not recorded.

Strengths and limitations

Our study benefited from using a truly con-textual and independent assessment of the built environment rather than measures derived from aggregating individual data. The small size of our surveyed areas ensured reasonable homogeneity within sectors. We used multilevel modelling to account for the hierarchical structure of the data. The study was large but its unique setting means its results are not necessarily generalisable to other cities throughout the world. Our independent measures at the highest level concentrated on the physical aspects of the environment, mostly because we thought that these could be measured reliably. Of course, the quality of the built environment also reflects something of the psychosocial environment, but we did not include these aspects in this study. This study should be taken as an invitation to explore this field further.

The assessment of the geographical sec-tors was undertaken almost 4 years after we finished the survey of the individuals. Although it is possible that the conditions in those neighbourhoods could have changed in the interim period, we did not find evidence that sectors had experienced major structural changes during the interval according to a survey of local government authorities (Secretaria Regional Ministerial de Planificacion y Coordinacion, 2005). A few sectors with a larger proportion of socially disadvantaged individuals were excluded from the analysis. The main reason for sector exclusion was the small number of people in the sector or the lack of data.

Common mental disorders are more preva-lent among socially deprived individuals; thus our estimates may be an underrepre-sentation of the true association. Finally, this is a cross-sectional study and as such we cannot infer the direction of causality. Equally, this kind of design cannot account for factors related to selective migration or population instability.

In conclusion, measuring the impact of the quality of neighbourhoods on mental health and understanding the complex interrela-tionships between individuals’ characteristics and their local environment are challenges that should be confronted, so that appropriate and effective inter-ventions can be implemented to improve the mental health of the population.

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Common mental disorders and the built environment in Santiago, Chile

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