Has the economic crisis widened the intra-urban socioeconomic inequalities in mortality? The case of Barcelona, Spain

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Abstract

Background: There is considerable evidence demonstrating socioeconomic inequalities in mortality; some of which focuses on intra-urban inequalities. However, all the studies assume that the spatial variation of inequalities is stable over the time. We challenge this assumption and propose two hypotheses: i) have spatial variations in socioeconomic inequalities in mortality at an intra-urban level changed over time? and ii) as a result of the economic crisis, has the gap between such disparities widened?. In this paper our objective is to assess the effect of the economic recession on the spatio-temporal variation of socioeconomic inequalities in mortality in Barcelona (Catalonia, Spain).

Methods: We used a spatio-temporal ecological design to analyse mortality inequalities at the small area level in Barcelona. Mortality data and socioeconomic indicators corresponded to the years 2005 and 2008-2011. We specified spatio-temporal ecological mixed regressions for both men and women, using two indicators; neighbourhood and year. We allowed the coefficients of the socio-economic variables to differ according to the levels and explicitly took into account spatio-temporal adjustment.

Results: For men and women both absolute and, above all, relative risks for mortality, have increased since 2009, In relative terms this means that the risk of dying has increased much more in the most economically deprived neighbourhoods than in the more affluent ones.

Conclusion: Although the geographical pattern in relative risks for mortality in neighbourhoods in Barcelona remained very stable between 2005 and 2011, socioeconomic inequalities in mortality at an intra-urban level have surged since 2009 onwards.

What is already known on this subject?

- There are a considerable number of studies which concentrate on socioeconomic inequalities in health outcomes and mortality in particular.
- There is evidence of socioeconomic inequalities in mortality, even at intra-urban level, that is to say most (socioeconomically) deprived groups of the population have higher mortality.
- Although, because of little evidence, results are not conclusive it would seem that inequalities have increased over time on a country level but have remained stable or have even decreased at the intra-urban level

What this study adds?

- Both absolute and, above all, relative risks for mortality for men and women, have increased from 2009 onwards.
- In relative terms the risk of dying increased much more in the most economically deprived neighbourhoods than the most affluent.
- The economic crisis could have the cause of the widening gap of socioeconomic inequalities in mortality at an intra-urban level in Barcelona.

1.- Introduction

Nowadays, there are a considerable number of studies which concentrate on socioeconomic inequalities in health outcomes and mortality in particular^[1-4]. Meanwhile, there are studies examining intra-urban socioeconomic inequalities in mortality^[5-10], (albeit less numerous they are increasing in number), with most of them encompassed by the MEDEAt^[4,11-15] and now, the IneqCities^[16,17] projects. All of these studies find evidence of such inequalities, that is to say most (socioeconomically) deprived groups of the population have higher mortality^[4]. There are fewer studies that analyze the temporal trends of these inequalities, either at the country level^[18,19], or the intra-urban level^[10,13,14,20]. Although, because of little evidence, results are not conclusive it would seem that inequalities have increased over time on a country level^[18,19] but have remained stable^[13,14] or have even decreased^[10,20] at the intra-urban level.

However, none of the studies that analyze the temporal trends of socioeconomic inequalities in mortality have considered the very likely spatiotemporal interaction of these inequalities. That is to say, they all assume that the spatial variation of inequalities is stable over the time. Our first hypothesis in this work is that spatial variations in socioeconomic inequalities at intra-urban level have in fact changed over time.

Our second hypothesis is that, as a result of the economic crisis, socioeconomic inequalities in mortality at intra-urban level have widened in Barcelona, Spain. Although there is no 'official' definition of an economic recession, the most frequently used defines a recession as 'a decline in the Gross Domestic Product (GDP) for two or more consecutive quarters'^[21]. This definition does not take into account changes in variables other than GDP. Furthermore, using a quarter as reference period makes it difficult to determine with any precision the beginning and the end of the recession. At any rate, according to this definition, the Spanish economy went into recession in the first quarter of 2009; after the GDP fell over two consecutive quarters.

Given these two hypotheses, our objective in this paper is to assess the effect of the economic crisis on the spatio-temporal variation of socioeconomic inequalities in mortality in small areas in Barcelona, Spain.

2.- Methods

Design

We use a small area spatio-temporal ecological design to analyse mortality inequalities in Barcelona.

Study population and information sources

The population studied was that of those residing in the neighbourhoods of Barcelona. According to the National Statistical Institute of Spain, the population of Barcelona (January 1, 2012) was 1,620,943 inhabitants, 769,819 men (47.49%) and 851,124 women (52.51%)^[22]. It is the second most populated city

in Spain, after Madrid, and the eleventh of the European Union. The density of population is quite high, 15,903.5 hab./km². It is a city with a population pretty aging (20.90% of the population aged 65 years or more; 17.36% of men and 24.10% of women). The Barcelona city council divided for administrative and statistical purposes the city into 73 neighbourhoods^[23], and these were used as the units of analysis. In 2011, the median of habitants per neighbourhood was 20,299 (10,112 men and 10,187 women) with an interquartile range equal to 10,923 - 31,526^[24]. Also in 2011, the neighbourhood with less inhabitants was 'La Clota', with 466 inhabitants, and most 'La nova Esquerra de l'Eixample' with 58,204 inhabitants^[24].

Mortality data and socioeconomic indicators corresponded to the years 2005 and 2008-2011. Mortality data for the year 2005 were obtained from the Barcelona register under the IneqCities project. The remaining data, including cartography, were obtained from the OpenDataBCN website of the Barcelona City Council^[24].

Variables

As response variables we included total male and female mortality (crude death rates) by neighbourhoods. As denominators we used sex and age group, also by neighbourhood (source of mortality and denominators: Barcelona register).

As explanatory variables we included disposable household income and the percentage of foreigners from low income countries (all by neighbourhood) (source: OpenDataBCN website^[24]). Analyses were conducted separately for men and women^[25].

Disposable household income is an index (Barcelona = 100) obtained from OpenDataBCN website, and constructed, elsewhere^[26], from five socioeconomic indicators: i) unemployment rate (computed as unemployed over resident population aged 16-65 years); ii) the percentage of resident population (per neighbourhood) aged 25 years or more with a university degree; iii) cars per 1000 over total resident population; iv) cars more than 16 horsepower (hp) but less than two years old over the total number of cars less than two years old; and v) private home resale prices^[26].

As an index, disposable household income does not capture all the variability contained in socioeconomic indicators that compose it. In order to reduce the variability not captured by the index, that is to say, to capture those aspects of deprivation that were not captured by disposable household income, we included two additional explanatory variables: unemployment rate and percentage of foreigners from low income countries. In fact, on the one hand, these indicators have been used for collecting deprivation in both MEDEA^[4,11-13] and IneqCities^[17] projects. On the other hand, these variables are the only components of the deprivation indices used in those projects, for which information is available at the neighborhood level of Barcelona.

As regards foreigners (i.e. immigrants) from low income countries, some studies have shown that they may contribute to increased health inequalities as

continue to have a low **socio**-economic status, but only in relative terms, compared, for instance, with immigrants from other areas of Spain with a lower income than Catalonia (the autonomous community to which it belongs Barcelona)^[27].

Data analysis

For each sex, we assumed the observed cases of deaths followed a Poisson distribution,

$$O_{it} \sim \text{Poisson}(\mu_{it} \text{Pop}_{it})$$

where O_{it} denoted the observed cases of death for a particular sex in the neighbourhood i ($i=1, \dots, 73$) in year t ($t=2005, 2008, 2009, 2010$ and 2011). μ_{it} was the relative risk in the neighbourhood i in year t , and Pop_{it} was the population for a particular sex in the neighbourhood i in year t .

In turn, the relative risk could be associated with the explanatory variables by means of spatio-temporal ecological regression. In our case, this regression was formulated as the following mixed model with two levels: neighbourhoods (denoted by i) and year (denoted by t):

$$\begin{aligned} \log(\mu_{it}) &= \alpha_i + \sum_{q=2}^5 \beta_{q,t} \text{HDI}_{q,it} + \theta \text{Foreigners}_{it} + \\ \log(\text{Pop}_{it}) &+ \delta_1 \text{Pop4564}_{it} + \delta_2 \text{Pop65}_{it} + S_i + T_t + \eta_{it} \end{aligned} \quad \{1\}$$

where $\text{HDI}_{q,it}$ denoted the q -th quintile of disposable household income in neighbourhood i in year t (all the quintiles were constructed for each year separately). The first quintile was taken as a reference value; Foreigners_{it} was the percentage of foreigners from low level income countries in the neighbourhood i in year t ; Pop_{it} denoted the population of the neighbourhood i in year t ; Pop4564_{it} denoted the percentage of the population aged between 45 and 64 years (both inclusive), and Pop65_{it} denoted the population over 65. The first age group (≤ 44 years) was not included in the model to avoid any problems of collinearity with the other two age groups. α_i the heterogeneity random effects, S_i the spatial random effects, T_t the temporal random effect (based on the temporal trend, $t=2005, 2008, 2009, 2010$ and 2011), η_{it} the spatio-temporal interaction random effects and β 's, γ 's, θ and δ 's were unknown parameters.

Note that in the specified model we did not use the number of cases expected in the neighbourhood as an offset (i.e., denominator), but rather population (male or female). In fact, we used the crude death rate and not the standardized mortality ratio as an indicator of mortality. The reason is to avoid the so-called 'mutual standardization' problem^[28-30]. In short, the problem when using standardized rates as a response variable in regression models is that it leads to biased results because only the response values and not the predictor values are adjusted by the same confounding factor, usually distribution by age^[28]. This problem was able to be solved by using the crude death rate and by including

the age structure of the population^[29,30] as an additional regressor,(i.e., Pop4664 and Pop65, in our case).

Also note that the coefficients associated with disposable household income had subscripts. This was because we allowed these coefficients to be regarded as, in mixed model terminology, random effects^[31]. That is to say, we allowed them to differ for the various levels we considered. In particular, we allowed the basal relative risk (e^{α_i}) to vary by city:

$$\alpha_i = \alpha + u_i$$

Meanwhile, the relative risks associated with the disposable household income varied by year:

$$\beta_{q,t} = \beta_q + w_t$$

Spatio-temporal adjustment

In all models the disturbance terms u , w and v were Gaussian random variables with constant variance, but not identically and independently distributed. In fact, with spatial data, as in our case, the variability in the observed cases is usually higher than expected (which produces overdispersion) and, more importantly, is not random. Here, two sources of extra variability have to be distinguished^[32,33]. Firstly, the largest source is usually named ‘spatial dependence’ and is a consequence of the correlation between the spatial unit and the neighbouring spatial units; generally the adjacent geographical areas. Variables observed at small area level are often interdependent because of proximity and linkages between neighbouring areas. In this sense, the closer areas are much more similar than the more distant ones. However, part of this dependence is not really a structural dependence, but mainly due to variables that are not included in the analysis.

The second source is independent, spatially uncorrelated extra variability, called uncorrelated or non-spatial heterogeneity, and which is due to the unobserved non-spatial variables that could influence the dependent variable^[32,33]. Spatial heterogeneity could be structural instability (the true values of the coefficients vary in space) or groupwise heteroskedasticity (the variance of the error term varies in space).

In our case, as we have a time dimension in our data, there was also temporal dependency (i.e., serial autocorrelation). Again, maybe not results of a dynamic behaviour of the dependent variable per se, but rather the omission of time-varying explanatory variables.

To take into account this spatio-temporal extra-variability, we needed to introduce some structure to the model, otherwise not only would the estimator’s standard errors be wrong but, because the dependent variables were not Gaussian, the estimators would be biased^[34].

Specifically, heterogeneity was captured using the random effects associated with the intercept (α_i), indexed by neighbourhood. For the spatial dependency, we followed the recent work of Lindgren *et al.*^[35] and specified a Matérn structure^[36] for the spatial random effects denoted by S (indexed by small area). In short, we used a representation of the Gaussian Markov Random Field (GMRF), explicitly constructed through stochastic partial differential equations (SPDE), which has as a solution a Gaussian Field (GF) with a Matérn covariance function^[35]. To sum up, instead of using the Matérn in a regular lattice, which is the usual practice and would imply an estimation with high computational costs and, also, one that is weak in terms of efficiency^[35], we specified the structure of the spatial Matérn covariance in a triangulation (Delaunay triangulation^[37]) of the studied area resulting in low computational costs and, more importantly in our context, far greater efficiency.

Both, temporal dependency and spatio-temporal interaction were approximated through a random walk of order 1^[38].

The large number of random effects in our model and, in particular, the explicit specification of the spatio-temporal dependence, called for the inferences to be made using a Bayesian approach. Specifically, we used the Integrated Nested Laplace approach (INLA)^[39,40].

A natural way to compare models is to use the criterion based on a trade-off between the fit of the data to the model and the corresponding complexity of the model. The Bayesian model comparison criterion based on this principle is called the Deviance Information Criterion (DIC)^[41]:

$$\text{DIC} = \text{'goodness of fit'} + \text{'complexity'} = D(\bar{\theta}) + 2p_D$$

where $D(\bar{\theta})$ is the deviance evaluated at the posterior mean of the parameters and p_D denotes the 'effective number of parameters' which measures the complexity of the model^[41]. DIC may under penalize complex models with many random effects^[42,43]. Hence we also used the conditional predictive ordinate (CPO)^[44,45] which expresses the posterior probability of observing the value (or set of values) of y when the model is fitted to all data except y_i , $\text{CPO}_i = \pi(y_i^{\text{obs}} | y_{-i})$ (y_{-i} denotes the observations y with the i -th component omitted). This facilitates the computation of the cross-validated log-score, cv.ls , ($\text{cv.ls} = -(\text{mean}(\log(\text{cpo})))$)^[46], for model choice. Both, the lower DIC and the lower cv.ls involve the best model

All analyses were performed with the free software R (version 3.0.1)^[47] made available through the INLA library^[39,40].

3.- Results

In Figures 1 and 2 we show the geographical distribution of both the response variable (crude death rates) (Figures 1) and disposable household income (Figures 2) by quintiles, in the neighbourhoods of Barcelona for 2008 (pre-crisis

year) (Figures 1a, 1c and 2a) and 2011 (crisis year) (Figures 1b, 1d and 2b). In both years a clear geographical pattern for the socio-economic variables is observed. The most economically deprived neighbourhoods were located in the north, northeast, south, and southeast of the city. The geographical pattern of mortality was very similar, at least with respect to men (Figures 1a and 1b). Note that the geographic pattern for the women was approximately the opposite of the men, with the highest mortality being in wealthy neighbourhoods in the west of Barcelona.

The temporal distribution of the socioeconomic variables by quintiles of disposable household income is shown in Figure 3. Note that there was a clear decrease in disposable household income for the first four quintiles but an increase in the fifth quintile which corresponded to the most affluent neighbourhoods; albeit with a large dispersion. In particular, the median of the disposable household income was 102.29 in 2005; 86.63 in 2008; 84.09 in 2009; 81.14 in 2010; and 81.00 in 2011. The interquartile range of the disposable household income changed from 11.27% of the median in 2005 to 32.72% of the median in 2011.

There was no clear temporal pattern in the death rates by quintiles of disposable household income (Figures 4). There appeared to be a downward trend, although the large dispersion (especially in the first quintile of disposable household income, as well as in 2005) largely prevents a decisive conclusion.

With respect to the spatio-temporal interaction, although in both the response variable and the socioeconomic variables we observed some changes in the geographical pattern between 2008 and 2011, in particular an increase in the number of economically deprived and with higher mortality neighbourhoods (Figures 1 and 2), the great extra-variability impeded establishing a clear spatio-temporal pattern.

Results of the estimation of the models are shown in Table 1 and Figures 5 to 7. The goodness-of-fit was somewhat better for men than for women. Note that we found evidence of socio-economic inequalities in mortality for both men and women i.e., relative risks for mortality were higher in more economically deprived neighbourhoods. However, the disparity was most evident among men rather than women (lower quintiles corresponded to a lower disposable household income). In most cases, the relative risks, i.e., inequalities were higher among men than women. Note also that in the case of men, the relative risk of mortality increased as the percentage of foreigners from low income countries increased.

Spatial dependence was important in both models (men and women) and in both cases spatial dependence was much greater than heterogeneity or the spatio-temporal interaction. Temporal dependency was practically negligible. That is, there was a clear geographical pattern in relative risks for mortality in Barcelona neighbourhoods that remained very stable between 2005 and 2011, although some heterogeneity existed alongside this.

In Figures 5 we show the temporal trend of the relative risk for mortality. There was an increase in the relative risk for men from 2009 onwards and then quite clearly for women from 2010 on.

The temporal evolution of the absolute risks of total mortality by quintiles of disposable household income is shown in Figures 6. Between 2005 and 2011, there was again a more apparent increase in the absolute risks for men in the first quintile of disposable household income (most economically deprived neighbourhoods) and a decrease in the fifth quintile (most affluent neighbourhoods). The temporal evolution of the absolute risks in the intermediate quintiles did not follow a systematic behaviour.

Finally, in Figures 7 we show the temporal evolution of the relative risks of mortality. When comparing the first quintile of disposable household income with respect to the fifth quintile, note that for both men and women there was a (statistically significant) increase in socio-economic inequalities in mortality from 2009.

4.- Discussion

Our results confirmed our hypotheses. Both absolute and, above all, relative risks for mortality for men and women, have increased from 2009 onwards. In relative terms the risk of dying increased much more in the most economically deprived neighbourhoods than the most affluent. As the economic recession in Spain began in the first quarter of 2009, this result could be (as was stated in our second hypothesis) the economic crisis causing the widening gap of socioeconomic inequalities in mortality at an intra-urban level in Barcelona.

It is very difficult to argue that this increase actually corresponds to one of the pathways (material, neo-material and psychosocial) in the relationship between income and health^[48,49]. In addition, if we consider that the analyzed period, 2005-2011, included only seven years and we have used an ecological design. However, evidence from past recessions suggests that health inequalities widen following them^[50-53]. It is very likely that this is a consequence of unemployment. It is argued that unemployed people are more likely to suffer poor physical and mental health than employed ones, besides that being unemployed is associated with unhealthy behaviors (smoking, alcohol and sedentary life)^[53]. This could make people more susceptible to higher death rates^a. Furthermore, when the unemployment rate of the neighborhood is very high and, as happened after 2009, the unemployment rate has increased dramatically, not only be unemployed but also the fear of being unemployed may contribute to increased health inequalities. In fact, we found that, in relative terms, the risk of dying increased much more in those neighbourhoods with the highest rates of unemployment.

However, we were unable to fully confirm our first hypothesis. In fact, although there was a clear geographical pattern in relative risks for mortality in Barcelona neighbourhoods, this remained very stable between 2005 and 2011.

In this sense, we think we have provided evidence in the debate over the existence of two (seemingly contradictory) trends in Spain. On one hand, health inequalities appear to have increased over time in Spain at a national level. On the other, health inequalities have either remained stable, as we found, or have even decreased over time at the intra-urban level^a.

Dalmau *et al.*^[10], using over four cross-sections (1983-1988, 1989-1994, 1995-1999 and 2000-2004) found that, for 1995-1999 and 2000-2004 inequalities in mortality in the city of Barcelona decreased in absolute and relative terms, especially among men (their spatial unit was the Basic Health Area). Likewise, Borrell *et al.*^[20], analyzing trends in mortality inequalities by educational level for main causes of death in the city of Barcelona between 1993-2003, found that relative inequalities in total mortality by sex did not change, whereas absolute inequalities tended to decrease in men. In fact, along the same lines as these studies, we only found an increase in socioeconomic inequalities in mortality from 2009 onwards. Previously, inequalities remained stable.

Although we do not know the ultimate reason for the different temporal evolution of health inequalities at national and intra-urban level, we believe that part of the discrepancy could be attributed to a methodological problem. When considering the national level, heterogeneity that might exist between different areas is omitted. If this is not adequately controlled, the estimation of the temporal trend in health inequalities could be seriously compromised.

With some exceptions (i.e., the temporal evolution of the relative risks of dying), the results in the case of the women were not as clear as that of the men. On one hand, this result is because most of the deceased were elderly, and as older women generally have healthier habits than men, (particularly practically not smoking^[54-57]) the largest number of women's deaths was concentrated in affluent neighbourhoods (compare Figures 1c-2a; and 1d-2b). This trend increases the extra variability, already very high when working with spatial design. In fact, this is the reason that the goodness-of-fit of the model for women was lower than for men and that the spatial dependency and, above all, the heterogeneity were also higher for the model for women. On the other hand, it is possible that the socio-economic variables that we employed did not fully capture women's material deprivation. In this sense, the disposable household income is an ecological index, unique for both sexes, whose component variables could be associated mainly with male material deprivation. However, it should be noted that since crude mortality rates were used, some differences on mortality could be explained, at least partially, because the effect of different distribution of age groups among neighbourhoods and between men and women^b.

As a main limitation to this work, we have used an ecological design and, therefore, usual caveats can be taken into account. In this sense, the methods applied in ecological regressions are able to describe the spatio-temporal distribution of mortality, but do not explain why the risk is higher in some neighbourhoods. For this, studies with individual data which evaluate the risk

^a This point was pointed out by one of the anonymous reviewers

^b This point was pointed out by other of the anonymous reviewers

factors and specific causes of mortality are needed^[58]. In reality, it is well known that in ecological studies a relationship between the exposure and factor risk studied cannot be obtained as it is not known whether those individuals more exposed to a certain factor in each geographic area are in fact the deceased.

However, the conclusions that can be drawn from these studies will make a significant contribution to the planning of health resources, and the implementation of programs for the prevention and the reduction of health inequalities.

Competing interests

There are no conflicts of interest for any of the authors. All authors freely disclose any actual or potential conflict of interest including any financial, personal or other relationships with other people or organisations within three years of beginning the submitted work that could inappropriately influence, or be perceived to influence, their work.

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Table 1.- Results of the estimation of the model

Fixed effects	Relative risk (95% credibility interval)	
	Men	Women
Disposable household income [Q1]		
Second quintile	1.058 (1.009, 1.108)	1.008 (0.959, 1.059)
Third quintile	1.145 (1.081, 1.213)	1.100 (1.037, 1.168)
Fourth quintile	0.907 (0.854, 0.963)	0.889 (0.835, 0.946)
Fifth quintile	0.985 (0.915, 1.060)	1.022 (0.949, 1.101)
Percentage of foreigners from low income countries	1.011 (1.006, 1.023)	1.007 (0.990, 1.004)

Random effects	Variance (standard deviation of the variance)	
	Men	Women
Heterogeneity	0.3664 (0.0426)	0.4324 (0.0356)
Spatial dependence	1.7209 (1.1859)	2.7676 (2.7526)
Temporal dependence	0.0996 (0.0312)	0.0561 (0.0189)
Spatio-temporal interaction	0.1722 (1.7849)	0.1495 (1.4215)

DIC	5871.20	6114.90
Effective number of parameters	84.08	83.95
-mean(log(cpo))	9.7050	9.8369

Reference category between brackets

Shaded, statistically significant at 95%

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Figure 1a.- Geographical distribution of disposable household income

Figure 1b.- Geographical distribution of unemployment rate

Figure 2a.- Geographical distribution of the crude death rate; men

Figure 2b.- Geographical distribution of the crude death rate; women

Figure 3.- Relationship between disposable household income and crude death rates in Barcelona, 2005 and 2008-2011.

Men

Women

Figure 4a.- Temporal distribution of disposable household income by quintiles of disposable income

Figure 4b.- Temporal distribution of unemployment rate by quintiles of disposable income

Figure 5a.- Temporal distribution of death rate among men, by quintiles of disposable income

Figure 5b.- Temporal distribution of death rate, among women, by quintiles of disposable income

Figure 6.- Temporal trend of the relative risk of mortality

Men

Women

Figure 7a.- Absolute risk of total mortality by quintiles of disposable household income; men

Figure 7b.- Absolute risk of total mortality by quintiles of disposable household income; women

Figure 8.- Relative risks, first quintile of disposable household income with respect to the fifth quintile

Men

Women

Dotted line (.....) is drawn in the upper limit of the 95% credibility interval