Internal migration dynamics in Spain before and after the 2008 economic downturn: continuities and changes

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Abstract: In this paper we try to ascertain whether significant changes in the internal migration distribution in Spain have occurred as a response to the 2008 economic crisis. For it, we first compare the evolution of the entire distribution for provincial net internal migration rates as well as changes within the distribution (intra-distribution dynamics) over the periods 2002-2007 and 2008-2013. Next, we analyse changes in migrants’ location preferences by means of a non-stationary Markov chain approach based on causative matrix estimation. Our results reveal that relevant changes in both the distribution and migrants' location preferences have taken place after the economic downturn in Spain.

Key words: internal migration, economic downturn, distribution dynamics, causative matrices, provinces, Spain.

JEL codes: C13, O18, R23

1. Introduction

If one had to single out one driving force behind the new direction of population dynamics in Spain since the late 1990s, undoubtedly that would be massive international migration. Without neglecting its importance so far, this issue has overshadowed however other migration streams that may be equally challenging at this time. That is the case of internal population movements (i.e. internal migration flows from one municipality to another) and its impact on income and population redistribution. Papers in this strand of the literature are very scant (see, for instance, Hierro, 2009; Paluzie et al. 2009; Hierro and Maza, 2010; Larramona and Sanso, 2012; Minondo et al., 2013). All in all, the severe economic crisis since the early 2008 has rekindled, albeit still not broadly enough, the interest for examining the response of internal migration movements to economic crisis. This issue seems particularly appealing in the Spanish case as the unemployment rate has risen at a faster pace than in its European counterparts and this fact might obviously have direct consequences on internal migration patterns. In this regard, leaving aside whether higher overall unemployment rates have produced, by reducing potential employment opportunities, an overall decline in mobility (Bentolila and Dolalo, 1991), it would be interesting to know whether significant changes in migrants' location preferences, stemming for instance from high dispersion in provincial unemployment rates and common labour reallocation over the cycle business, have taken place. One recent paper by Minondo et al. (2013) tries to throw some light to this and other questions by estimating different causal models. One conclusion they draw is that regions that have performed best in Spain over the crisis period in terms of per capita GDP growth have been common destinations for internal migrants.

Given this background, it is the aim of this paper to explore the main shifts occurred in the internal migration dynamics after the 2008 economic downturn by taking on a different methodological perspective: a distribution approach. Specifically, we aim to analyse changes in the external shape of the distribution, in the relative position of provinces within it and in the relative attractiveness of provinces for migrants. In addition, we think the use of provincial instead of regional data may enrich the study and provide more insights into data analysis.

The internal migration data used in this paper are taken from the 'Statistics of Residential Variations' (EVR in its Spanish acronym) published by the Spanish National Statistics Institute (INE). This database collects annual inter-municipal changes of residence declared by migrants in the Municipal Register (*Padrón Municipal*), thus including both residential displacements and labour market mobility. Although some under-registration problem is assumed when working with this database, as well as a certain disjoint between the time when a migration is made and the time when it is declared, this statistics is extensively used as providing the most trustworthy annual data on internal population movements. As indicated above, we opted to use Spanish provinces as units of analysis, instead of Spanish regions (NUTS 2 level following the Eurostat definition), as it allows one to take into account movements across provinces belonging to a region that go unnoticed when using a regional disaggregation. Annual population data are also collected by the INE.

The paper is organised as follows. Section 2 provides a short overview of internal migration trends in Spain over the sample period 2002-2013. In Section 3 we examine whether the recent economic crisis has altered the dynamics of internal migration movements across the Spanish provinces. For it, we analyse: first, changes in the external shape of the distribution using kernel density estimation; second, changes in intra-distribution dynamics by means of a Markov chain analysis and the well-known stochastic kernel approach; and finally, changes in relative attractiveness of provinces for migrants by using the so-called causative matrix approach. Finally, Section 4 presents our main conclusions.

2. Internal migration in Spain: an overview

Internal migration patterns in Spain have changed drastically over the last few decades. All along the 1960s and 1970s, increasing job opportunities in new industrialised provinces and the high urban-rural wage gap propitiated a "migration drain" of people from poor to rich provinces and, consequently, a situation characterised by extremely polarised internal migration balances that contributed significantly to the reduction of income disparities across the Spanish provinces

(Raymond and García, 1996). In the early 1980s, the industrial restructuring associated with the economic recession after the first oil price shock was accompanied by a great deal of return migrations to poor provinces in the South and Southwest of Spain. From that time on it opened a new phase characterised by a dramatic drop in overall net migration rates that started to make the contribution of internal migration to income convergence practically negligible. The expansion of the welfare state in the late 1980s, the upward trend in housing costs and the increasing role of amenities in migrants' decisions from 1990s onwards, as well as the sudden prominence of foreign-born internal migration from the early 2000, have also reshaped internal migration patterns in Spain in recent decades. A detailed, but not exhaustive, list of contributions to this strand of literature includes the papers by Bentolila and Dolado (1991), Antolín and Bover (1997), Ródenas and Martí (1997), Bentolila (2001), Bover and Arellano (2002), Devillanova and García Fontes (2004), Maza and Villaverde (2004), Hierro (2009), Paluzie et al. (2009) and Minondo et al. (2013).

Herein, our study is confined to a relatively recent period, that is 2002-2013. Taking a close look to Table 1, several characteristics should be noted regarding recent trends in internal migration movements. First, that after a steady increase in internal migration until 2007, a fall −albeit not very pronounced− in migration figures happened. As also seen from Table 1, the gross internal migration rate[1] increased 25% over the period 2002-2007 (from 31.6‰ to 39.7‰), while it dropped nearly 8% between 2008 and 2013 (from 35.6‰ to 32.9‰). Among the reasons that might be hindering higher levels of mobility across the Spanish provinces since 2008 onwards, Minondo et al. (2013) point to the effect of unemployed benefits by reducing the job-search intensity in other provinces, and the housing market. Other factors might be the high increase of people (both Spaniards and foreigners) leaving the country in search of labour opportunities in other countries,[2] as well as the slowdown in international migration to Spain (which played a significant role in internal migration since the early 2000s) due to the lack of job vacancies for

[1] This rate is measured as gross migration flows divided by total population.

[2] Population leaving Spain over the 2008-2013 period reached 2.1 million people.

immigrants during the crisis. Second, although intra-provincial migration is still more prevalent than inter-provincial migration (it represents nearly 60% of total internal migration in 2013), it has decreased more sharply in importance over the crisis period than inter-provincial migration, probably related to the adjustment in the housing industry that took place steeply since 2008 onwards.

Table 1 Internal migration in Spain, 2002-2013

Year	Total internal migration		Intra-provincial migration		Inter-provincial migration	
	number	rate ‰	number	%	number	%
2002	1,323,927	31.6	903,030	18.9	533,445	12.8
2003	1,467,903	34.4	1,008,014	20.6	586,987	13.7
2004	1,527,446	35.4	1,042,358	21.0	619,461	14.3
2005	1,570,361	35.6	1,082,083	21.4	626,324	14.2
2006	1,732,309	38.7	1,199,181	23.5	682,606	15.3
2007	1,795,353	39.7	1,203,728	23.2	744,716	16.5
2008	1,643,210	35.6	1,097,626	20.8	684,384	14.8
2009	1,653,014	35.4	1,122,232	21.0	669,573	14.3
2010	1,681,395	35.8	1,153,071	21.5	671,198	14.3
2011	1,650,298	35.0	1,119,165	20.7	672,003	14.2
2012	1,586,075	33.6	938,482	19.9	647,593	13.7
2013	1,551,940	32.9	929,436	19.7	622,504	13.2

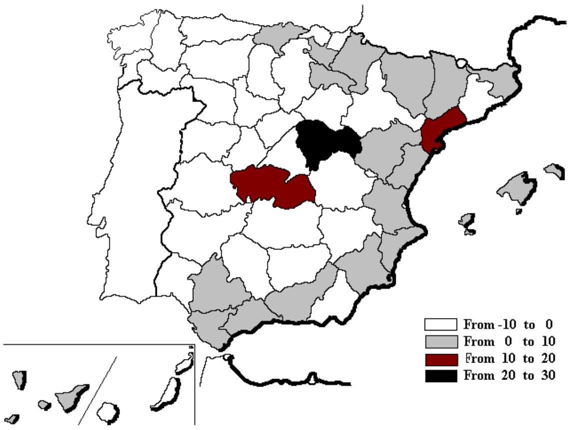
Source: INE and own elaboration.

Now we take a look to the spatial distribution of net migration rates[3] across the Spanish provinces. For convenience of comparison we only display the distribution in years 2003, 2008 and 2013. The value of the net migration rate is reflected in the relative shade used, i.e. the darker the shade the higher the value of the rate. The two main aspects to be highlighted are, on the one hand, the sharp drop of net migration rates over time and, on the other, the increase in the number of provinces with a negative, albeit low, internal migration balance (represented by the unshaded areas) when comparing years 2008 and 2013. Turning to more particular aspects, another

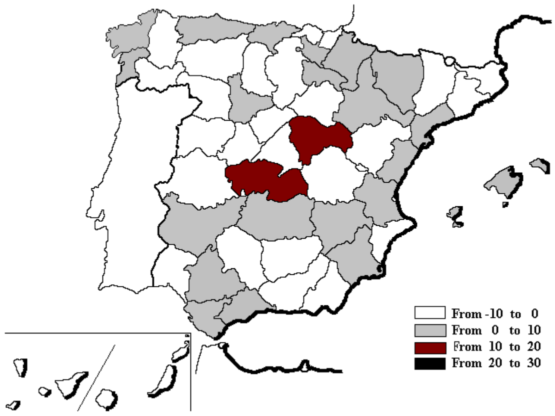
[3] This rate is measured as the difference between migration flows into a province minus the migration out of that province, per 1000 inhabitants.

issue to be looked at is the internal migration turnaround in Guadalajara and Toledo (both well-known as being classical destinations for residential reasons due to its closeness to Madrid) which startlingly transformed them into provinces with a negative internal migration balance (in both cases this exactly happened in 2012).

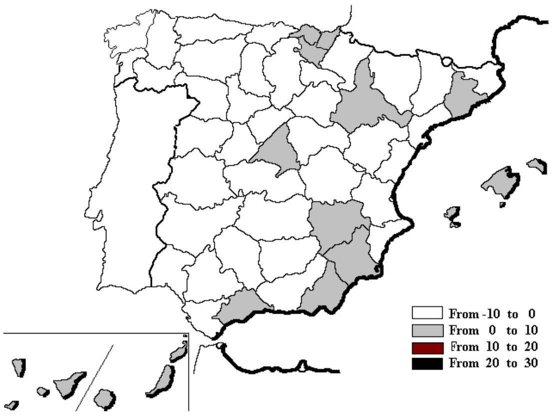
Figure 1. Provincial net internal migration rates in Spain



a) 2002



b) 2008



c) 2013

3. Internal migration dynamics in Spain before and after the 2008 crisis: A distribution dynamics approach

The aim of this section is to ascertain whether, as expected, the 2008 economic downturn has reshaped the dynamics of internal migration flows across the Spanish provinces. In order to gain insights into this issue, we first look into the evolution of the external shape of the distribution of net migration rates for the Spanish provinces as well as changes within it (intra-distribution dynamics). Then, the study directs its attention to changes in provinces' relative attractiveness for migrants.

3.1. Changes in the external shape of the distribution

First we pay attention to the evolution of the external shape of the distribution of net migration rates for the Spanish provinces over the sample period. To do it, we resort to the construction of a density function. This representation, which can be understood as being a smoothed version of a histogram, provides a very intuitive graphical tool for detecting some intrinsic characteristics of the distribution, such as the formation of groups or clusters. In addition, comparison of a density function at different points in time allows us to get some idea of how the distribution evolves over time.

Specifically, in this paper we estimate univariate kernel density functions for the years 2002, 2008 and 2013 using a Gaussian kernel. Because of data sparseness, an adaptive rather than a fixed bandwidth is used. This is a common practice when estimating long-tailed distributions as it allows us to reduce undersmoothing in areas with few observations while oversmoothing in others. In order to do it, we use the standard adaptive two-stage estimator proposed by Abramson (1982) given by:

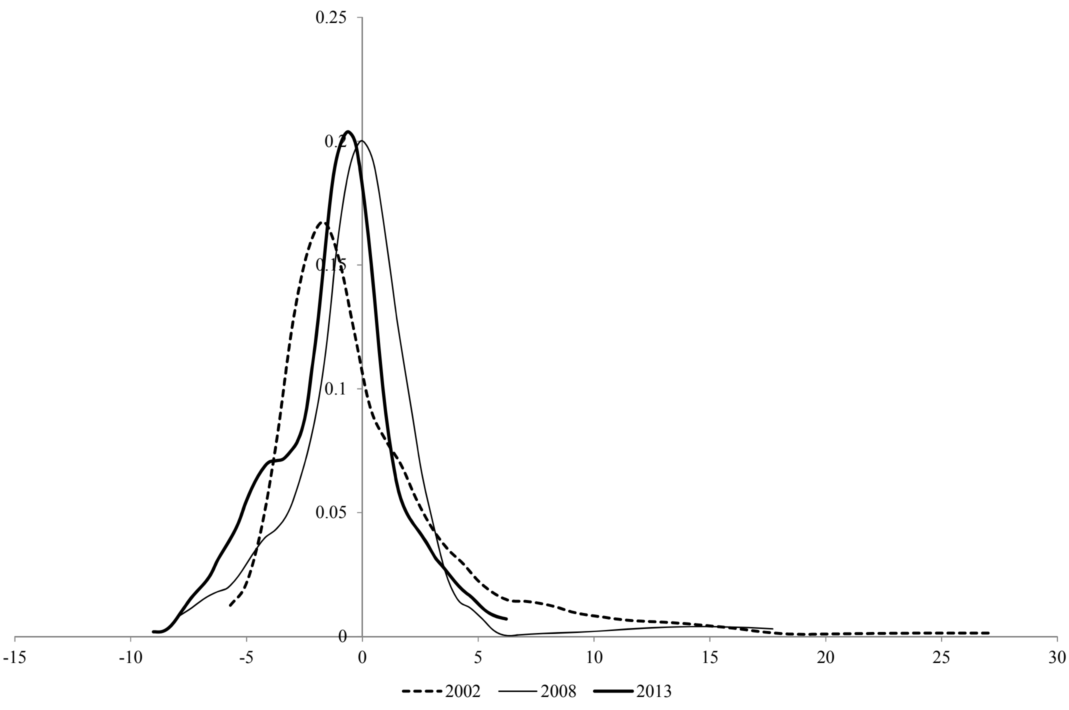
$$\hat{f}(x) = \frac{1}{\sum_{i=1}^{n} w_i} \sum_{i=1}^{n} \frac{w_i}{h_i} K\left(\frac{x - x_i}{h_i}\right) \tag{1}$$

where K is a Gaussian kernel, $h_i = h\lambda_i$ is a varying bandwidth defined as the product of a global fixed bandwidth h and a bandwidth adjustment factor

$\lambda_i = \sqrt{G / \tilde{f}(x_i)}$, and G is the geometric mean over all i of the standard fixed bandwidth kernel density estimate $\tilde{f}(x_i)$.

Figure 2 plots the net migration rate distribution for the years 2002, 2008 and 2013. Among the main conclusions drawn from this plot it can be highlighted the following ones. First, significant shifts have occurred in the external shape of the distribution between 2002 and 2013. The most prominent one is that the long tail of the distribution to the right, extending to very high values at both 2002 and 2008 years and associated to provinces with very high positive net migration rates, nearly vanished at the end of the sample period. On the contrary, the relatively small tail of the distribution to the left of 0 has remained practically size invariant. Second, the main mode of the distribution has remained nearly unchanged around negative values very close to zero, and clearly shows the predominance of a very low but negative net migration balance. Finally, we can note the appearance of a new secondary bump at 2013 around -4 perhaps associated with the emergence of a group of provinces where net migration has declined further as more severe economic conditions and rising unemployment took place.

Figure 2. Adaptive kernel density



3.2. Changes in intradistribution dynamics: A Markov chain approach and a stochastic kernel approach

Although informative, comparison of density functions at different points in time does not allow us to obtain a precise picture on the law of motion of the distribution. Indeed, it might happen that changes in provinces' relative position take place, albeit the external distribution shape not being affected (Quah, 1997). In order to resolve this shortcoming, in this section we first apply a discrete intra-distribution dynamics approach: the so-called Markov chain approach. This is appropriate since it makes a temporal system-wide approach to migration possible by estimating the so-called transition probabilities. The interest for this approach goes further, as it allows us to quantify mobility within the distribution.

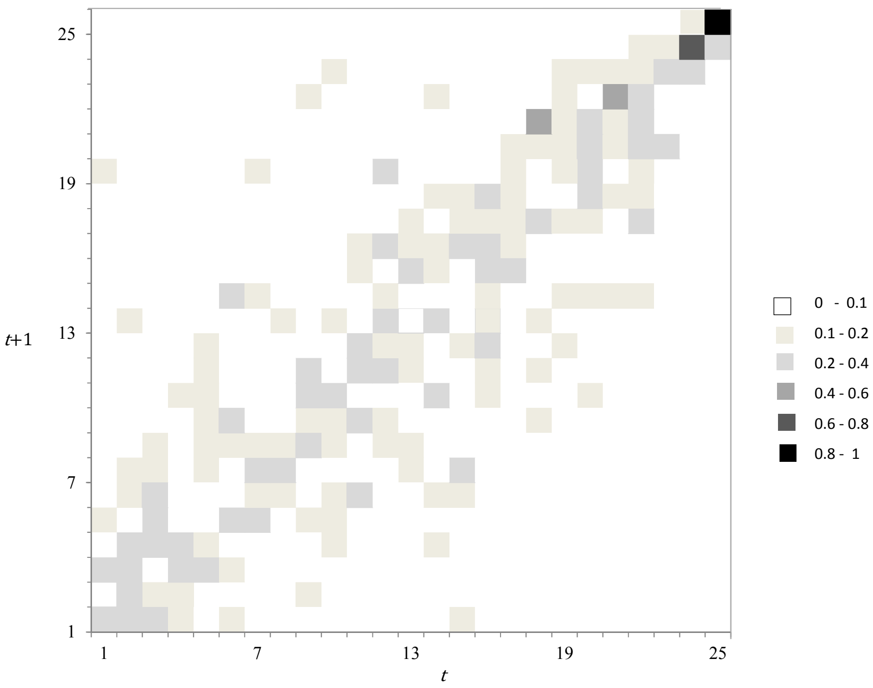
For it, let suppose that provinces are classified into a finite number of exhaustive and mutually exclusive states (in practice, intervals) according to their net migration rate, and that X_t represents the state occupied at time *t*. Then, it is possible to define the distribution for the net migration rate at times *t* and *t*+*s*, denoted by $p(t)$ and $p(t+s)$ and commonly referred as to initial and final distributions respectively. The link between both distributions is given by $p(t+s) = p(t) \cdot P(t, t+s)$, which defines the law of motion of the distribution between *t* and *t*+*s*. The key element in the preceding equation is the operator $P(t, t+s)$, the so-called transition matrix between *t* and *t*+*s* with generic elements $p_{ij}(t, t+s)$ which maps the distribution from period *t* to period *t*+*s*. The interpretation of the transition matrix is particularly intuitive as its elements give the probability of moving from a state *i* to another *j* between *t* and *t*+*s*.

In implementing this approach several types of decisions must be made, one of which involves the transition period length. In this case, and as is common in many applications on internal migration intra-distribution dynamics, we opted for estimating a one-year transition matrix (that is, $s=1$). This is because as migration flows are usually exposed to relatively high variability, a longer transition period might lead, in the case of discrete-time estimation, to a noteworthy loss of information. Another decision, probably the most important in this case, concerns the partition of the state space into a finite number of intervals. In order to avoid, at least partially, some of the issues associated with discretisation of the distribution, in this paper the number of states reaches a total of 25 using as criteria percentiles defined from 0% on increments of 0.04%.[4] The upper bounds of the states are in Table 3. Being conscious that an excessive number of states can expand in excess the dimension of the transition matrix and create a practical difficulty in terms of space and visualization of the results, in this paper we resort to an informal representation of the estimated one-year transition matrix by plotting probability "bands".[5] Figure 3 plots the one-year transition matrix for periods 2002-2007 and 2008-2013, respectively. On the axes we represent the net migration states (location of provinces at t and $t+1$ is shown on the horizontal and vertical axis, respectively), the shaded black areas being, from darkest to lightest, 0-10%, 10-20%, 20-40%, 40-60%, 60-80% and 80-100% probability bands. If we compare both periods, several general observations can be made about the pattern of transition probabilities. First, as revealed by cells in the main diagonal, persistence does not seem to be an inherent characteristic of the distribution. Second, both forward and backward movements exist in both periods, although mobility seems to be more accentuated in the period of crisis than in the preceding one.

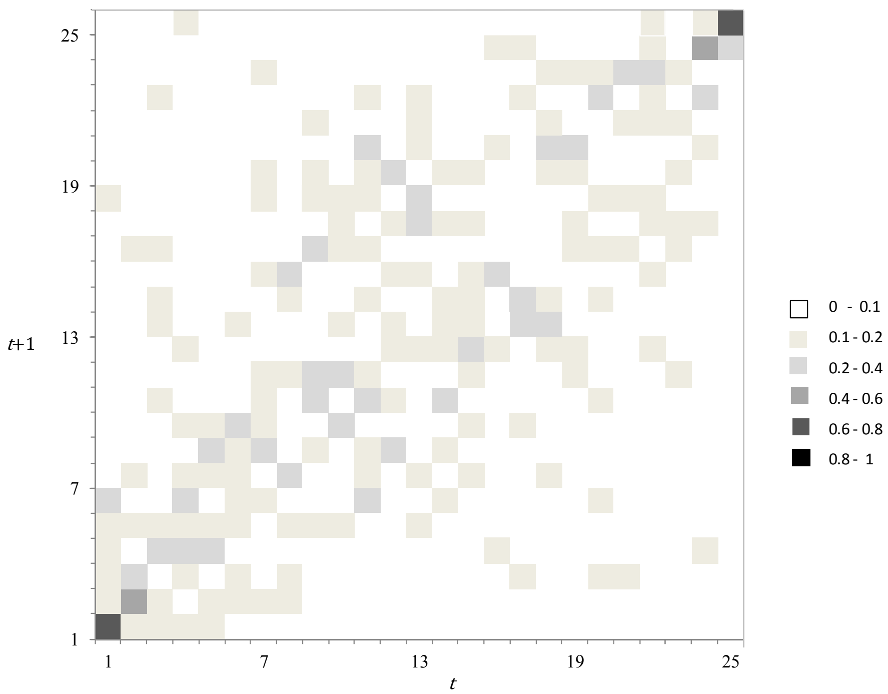
[4] Alternative criteria based on the calculation of an optimal bin width (see, for instance, Scott, 1979) were discarded as generating, in our opinion, not high enough number of states. The recent proposal by Rey (2014) based on examining movements within rank distributions was also disregarded as requiring a number of year periods much higher than the number of regional units.

[5] Before proceeding with the estimation, we first tested for the existence of Markovian dependence using the χ^2−test proposed by Anderson and Goodman (1957). The results lead us to reject the null hypothesis of non-Markovian dependence at the 5% significant level (p−value=0.000), this implying we can properly compute a transition matrix.

Figure 3. Estimated one-year transition matrix



a) 2002-2007



b) 2008-2013

However, the lack of a better knowledge about the relative significance of mobility among states prevents us from drawing more precise conclusions about this question. Such is the case that very close upper bounds obtained in the negative side of the distribution (see Table 3) is likely to detract from the significance of movements at that side. Thus further examination of the significance of movements seems to be essential to arrive at more precise conclusions. In order to do it, we resort to a novel mobility measure formulated by Maza et al. (2010). This consists basically on an extension of Bartholomew's (1996) family of mobility measures that allows to account for both the size of the states and the relative distance between states, this latter being a crucial point for the measurement of intra-distribution mobility in this case. The expression of the referred mobility measure is as follows:

$$d(P(t,t+1)) = \sum_i \sum_j \frac{1}{k_i(t)} \cdot p_i(t) \cdot p_{ij}(t,t+1) \cdot d_{ij}(t) \qquad (2)$$

where p_i are the size of states at *t* (in this case the size is the same as states contain equal number of provinces); p_{ij} denote transition probabilities between *t* and *t*+1; $d_{ij} = \left| \overline{tmn}_j - \overline{tmn}_i \right|$ are absolute differences between the average net migration rate between states at *t*; and, finally, k_i denotes the largest value of each row in matrix *D* (distances matrix with generic elements d_{ij}). This mobility measure is perfectly bounded between 0 and 1, and its interpretation is straightforward: the closer its value to 1, the higher the mobility degree within the distribution. Specifically, $d(P) = 1$ if all provinces change their relative position within the distribution moving either upward or downward towards the more distanced net migration state.

To get insights into the performance of each state separately, the aggregate mobility measure in equation (2) can be decomposed into the so-called state-by-state measures, denoted by $d(P_i)$ that is:

$$d(P(t,t+1)) = \sum_i p_i(t) \cdot d(P_i(t,t+1)) \qquad (3)$$

where

$$d(P_i(t,t+1)) = \sum_j \frac{1}{k_i(t)} \cdot p_{ij}(t,t+1) \cdot d_{ij}(t) \tag{4}$$

Table 3 presents state-by-state and aggregate mobility indexes for the two periods under consideration. If we first take a look into the aggregate index, we find that intra-distribution mobility in period 2008-2013 $\left(d(P) = 0.127\right)$ is distinctly higher (more than two-fold higher) than that found in period 2002-2007 $\left(d(P) = 0.059\right)$.

Table 3. Mobility measures

State	2002-2007			2008-2013		
	Upper bound	$d(P_i)$	$d(P)$	Upper bound	$d(P_i)$	$d(P)$
1	-4.15	0.050		-5.03	0.088	
2	-3.59	0.031		-3.82	0.104	
3	-2.92	0.035		-2.97	0.187	
4	-2.70	0.033		-2.52	0.134	
5	-2.41	0.036		-2.00	0.127	
6	-2.12	0.057		-1.74	0.170	
7	-1.84	0.046		-1.48	0.105	
8	-1.44	0.032		-1.21	0.065	
9	-1.07	0.036		-0.95	0.086	
10	-0.85	0.048		-0.84	0.075	
11	-0.66	0.033		-0.64	0.074	
12	-0.35	0.038		-0.45	0.087	
13	0.02	0.045	0.059	-0.31	0.091	0.127
14	0.23	0.089		-0.18	0.088	
15	0.48	0.026		-0.03	0.089	
16	0.85	0.028		0.12	0.185	
17	1.21	0.050		0.25	0.114	
18	1.69	0.049		0.46	0.196	
19	2.04	0.094		0.63	0.125	
20	2.51	0.060		0.94	0.108	
21	3.29	0.049		1.21	0.127	
22	4.55	0.079		1.59	0.197	
23	6.06	0.161		2.23	0.133	
24	13.18	0.230		3.65	0.258	
25	30.57	0.043		16.85	0.161	

From values of state-by-state indexes it is also apparent that provinces acting as a magnet attracting more internal migration flows in Spain have played a major role in the aggregate mobility. However, pursuing the evolution of the indexes over time seems to reveal that major contribution to aggregate mobility also comes from provinces with a worse migration balance.

Even so it is worth noting that the discrete approach suffers of an important drawback: the results obtained critically depend on the arbitrary number and width of the intervals considered. Although we think that this drawback is partially (or almost fully) resolved in our case, we also think that more sound results would be obtain by resorting to a continuous version of the Markov-chain approach, namely the stochastic kernel approach (see Quah 1997; and Durlauf and Quah 1999). For it, this approach generates a conditional density function defined, for any discrete variables *Y* and *X*, as:

$$\hat{f}_{\tau}(y|x) = \frac{1}{b}\sum_{i=1}^{n}\omega_i(x)K\left(\frac{\|y - Y_i\|_y}{b}\right), \tag{5}$$

where

$$\omega_i(x) = K\left(\frac{\|x - X_i\|_x}{a}\right) \Bigg/ \sum_{j=1}^{n}K\left(\frac{\|x - X_j\|_x}{a}\right). \tag{6}$$

The norms $\|.\|_x$ and $\|.\|_y$ represent Euclidean distances on the spaces of *X* and *Y*, while *a* and *b* are smoothing or bandwidth parameters on the two spaces respectively. $K(.)$ is the kernel function. Equations (5) and (6) show how a conditional density function in the continuous variables *x* and *y* can be obtained as the sum of *n* kernel functions in *Y* space weighted by the $\omega_i(x)$ in *X* space.

Hyndman *et al.* (1996) developed a technique which presents at least two main advantages over the traditional conditional density estimator just described. First, this new estimator has better statistical properties; and second, it provides some powerful visualisation tools (the so-called *stacked conditional density* and the *highest conditional density region*). Plotting these measures permits an easier and more direct interpretation of the results. The estimator proposed by Hyndman *et al.* (1996) is:

$$\hat{f}_{\tau}^{*}(y|x) = \frac{1}{b}\sum_{i=1}^{n}\omega_i(x)K\left(\frac{\left\|y - Y_i^{*}(x)\right\|_y}{b}\right), \tag{7}$$

where $Y_i^{*}(x) = e_i + \hat{r}(x) - \hat{l}(x)$. In this definition, $\hat{r}(x)$ is the estimator of the conditional mean function $r(x) = E[Y|X = x]$, $e_i = y_i - \hat{r}(x_i)$, and $\hat{l}(x)$ is the mean of the estimated conditional density of $e|X = x$.

A fundamental factor in the estimation of stochastic kernels - both the traditional estimator (equation (5)) and the one employed in this paper (equation (7)) - is the choice of the bandwidths; the function of these bandwidths is to put less weight on observations that are further from the point being evaluated. Specifically, we use optimal bandwidths in the two directions x and y following the Bashtannyk and Hyndman's (2001) rules. As regards the kernel function, once again we use a Gaussian kernel.

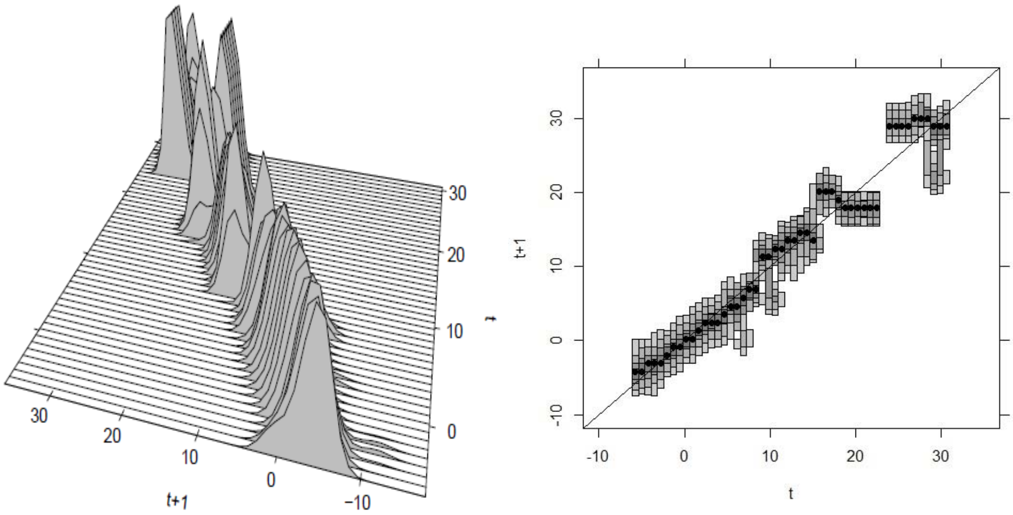
The results of this new approach, applied to the periods 2002-2007 and 2008-2013 and for a time-span of one period for the sake of consistency, are displayed in Figure 4: the stacked density plot on its left side and the highest conditional density region plot on its right side. With reference to the stacked density plots, they allow us to see the changes in the shape of the migratory rates for a given migratory rate in the previous year. According to the results, a striking difference between the two sample periods is evident: while, in the first one, the probability mass and most of the peaks tend to be clustered along the main diagonal, in the second period there are some apparent deviations from this diagonal, mainly at high rates of migration. This simply means that the mobility degree in the Spanish migration rate distribution was much higher during the second than during the first period.

However, a more informative way to represent the changes occurring in a distribution is based on the highest conditional density region plot. Each vertical strip on the right hand side of Figure 4 represents the conditional density for a migration rate in the previous year. In particular, this figure shows the highest density regions for a probability of 25, 50, 75 and 90% (as it passes from a darker to a less dark area). In addition, it illustrates, as a bullet ($\bullet$), the mode (value of migration rate in the year $t+1$ where the density function takes on its maximum value) for each value in the year t.

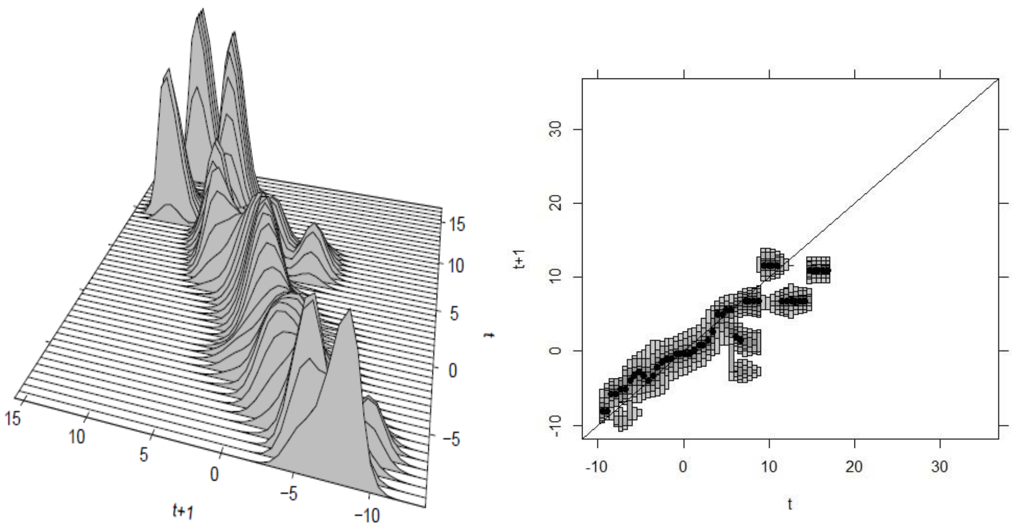
With respect to the first period (right side of Figure 4a), the position of the modes indicates that, generally speaking, changes have not been very significant. This result is confirmed if we observe the mass of probability (dark areas), as we see that, in general terms, the area representing a probability of 25% crosses the diagonal; this reveals again the existence of a high degree of persistence. The only exception occurs in the upper tail of the distribution.

Relative to the second period under study, it is important to notice that (right-hand side of Figure 4b) the mobility degree was higher than in the first one; as can be seen, the modes are now further to the diagonal and the dark areas representing a probability of 25% do not cross it in more cases than in the first period, especially in the tails of the distribution. In particular, the figure shows that mobility has been mainly confined to regions in the migration rate range of -10−-5 and 5−20; in the latter case our results reveal that provinces with relatively high migration rates have seen reduced them over time.

Figure 4. Stacked conditional density and highest conditional density region plots



(a) 2002-2007



(b) 2008-2013

Notes: From dark to light, the shadings represent 25%, 50%, 75% and 90% of the total probability. Bullets indicate the mode. Both the stacked conditional density plot and the high conditional density region plot were estimated at 50 points.

3.3 Changes in relative provincial attractiveness for migrants

At this point, we pause to take a fresh look at changes in relative provincial attractiveness for migrants; put the same in another way, in migrants' preferences for particular destinations. To do it, we resort to a Markov chain approach called the *Causative Matrix (CM) model* suggested by Lipstein (1965), and after extended by Plane and Rogerson (1986) and Hierro (2009). The most attractive features of this approach are as follows: (1) opposite to classical Markov chain analysis based on time-invariant transition probabilities (see, for instance, Magrini, 1999; Hammond, 2004; Ezcurra et al., 2005), it assumes a non-stationary specification; this is appealing in this context as it allows to understand the way provinces move up or down in the internal migration hierarchy (intra-distribution dynamics) under the premise that the probability of a province to move from one state to another can change over time, (2) the consideration of *inter-provincial dependency effects* through a constant causative operator; in doing so, as Hierro and Maza (2009) indicates this model goes beyond a simple comparison of transition matrices, and (3) this approach allows to deal with changes in relative provincial attractiveness. Thus, as long as considering that transition probabilities change over time, this approach assumes that apart from conditions influencing the transition probabilities between provinces, those influencing all other "competing" provinces are also taken into account (Plane and Rogerson, 1986).

Following this approach, changes between transition probabilities can be modelled as follows:

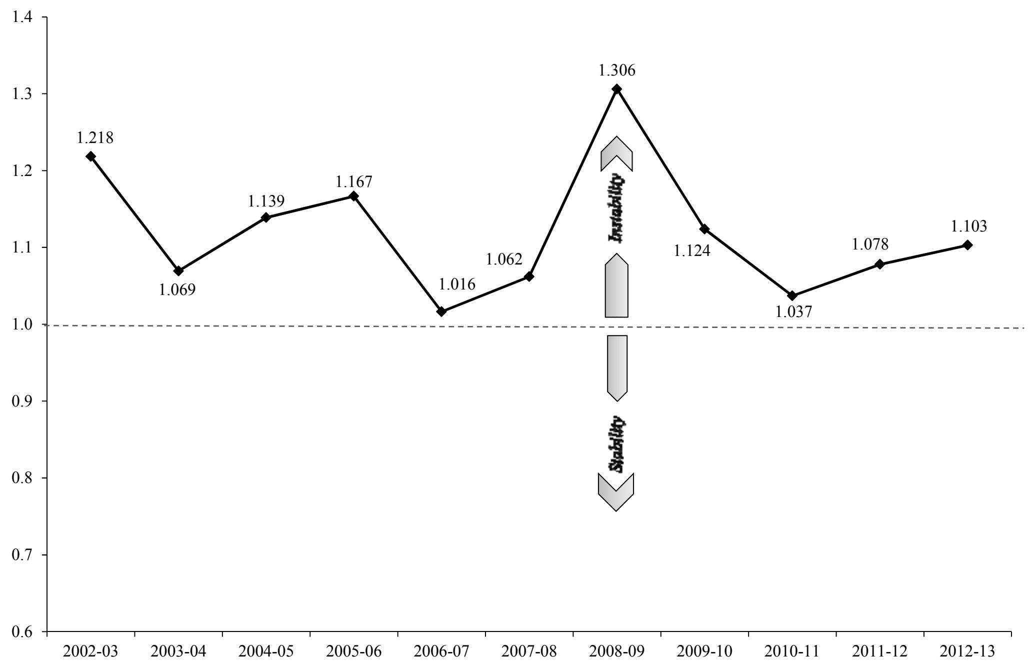
$$p_{ij}(t,t+1) = \sum_{k} p_{ik}(t-1,t) \cdot c_{kj}^{R} + \varepsilon_{ij}(t,t+1) \qquad (8)$$

for all i, j where ε_{ij} are the elements of a matrix of random variables $E(t)$ that captures the combined effect of different kind of factors that exert some influence of migrants' decisions, and c_{kj}^{R} are the elements of the so-called right-causative matrix C^{R} that gauges the rate of change of transition probabilities from a *competing destination perspective* (Plane and Rogerson, 1986). Accordingly, as indicated by equation (8), a transition probability tomorrow $p_{ij}(t,t+1)$ is not only influenced by its value today $p_{ij}(t-1,t)$ (that is to say, when $k = j$ in equation (8)), but also by the current transition likelihood from province i to all the other "competing" provinces outside of destination j (where $k \neq j$). In such manner, the model captures not only the direct effect of $p_{ij}(t-1,t)$ on $p_{ij}(t,t+1)$, but also the induced effect (in some cases offsetting and in other enhancing effect) of probabilities $p_{ik}(t-1,t)$. Put it differently, as the competitive position of other provinces vis-à-vis *j* can change, *j*'s attractiveness for potential migrants from province *i* may also see altered. As explained by Plane and Rogerson (1986), a column sum $\sum_k c_{kj}^R$ may be interpreted as the total change in attractiveness of destination *j*. Specifically, if a column sum is greater (lower) than 1 an increase (decrease) in the relative attractiveness of destination province *j* has happened. [6]

But before going any further with the analysis, it is mandatory to justify the length of the periods to be compared. Specifically, we must address if a disruptive dynamic change has effectively taken placed since 2008. To do it, the causative matrix approach provides an easy tool for structural change determination within the distribution based on the largest eigenvalue of the causative matrix λ^* (Lipstein, 1965). In particular, if $\lambda^* > 1$ some disturbances are pushing the system toward disequilibrium. On the contrary, if $\lambda^* < 1$, the system is tending toward equilibrium. In this last case, it is expected that eventually all provinces will achieve stable migration shares from the system; indeed, the closer the second highest eigenvalue λ^{**} is to unity, then the higher the rate of convergence (Plane and Rogerson, 1986). In a case of perfect stability, i.e. stationary transition probabilities, we would have all eigenvalues equal to unity. Figure 5 displays the evolution of λ^* values obtained from causative matrices derived from comparing year by year transition matrices (see Hierro, 2009).

[6] In case of stationarity, all provinces would undergo no change in relative attractiveness and, subsequently, C^R would be equal to the identity matrix.

Figure 5. Intra-distribution stability: maximum eigenvalue of causative matrices



A quick glance to the figure allows us to draw two main conclusions. First, that instability within the distribution was the rule for the whole period. Second, that a break or disruptive dynamic change occurred at 2008. This last conclusion seems to support our suspicion that the beginning of the crisis provoked a break in the internal migration dynamics in Spain. This piece of evidence justifies the choice of the year 2008 as a cut-off time representing two key periods in the internal migration dynamics since 2002: 2002-2007 (pre-crisis period) and 2008-2013 (crisis period).

Then, we have proceeded by estimating equation (8) for each period separately in order to measure total change in provinces' attractiveness in periods 2002-2007 and 2008-2013, respectively. For it, we have used OLS estimation. As indicated by Dent (1973), this kind of estimation is fruitful in this case as it allows us to obtain rows of C^R adding up to unity. It is worth pointing out that this is a desired property as it makes the aforementioned measurement of relative attractiveness of destinations suggested by Plane and Rogerson (1986) possible.[7] It must also be worth noting that

[7] In addition, rows of a causative matrix should add up to unity as a causative matrix is by definition the product matrix of two matrices with unitary row sums.

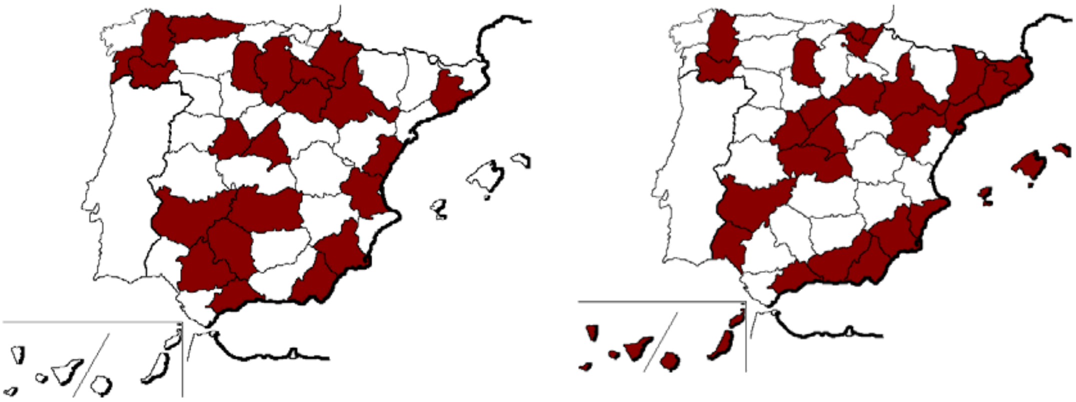
as dependent and independent variables in equation (8) are matrices instead of column vectors, each column of the causative matrix must be estimated separately as follows (Rao, 1965):

$$C^R(.,j) = inv\left(\sum_t P^T(t-1,t)\cdot P(t-1,t)\right)\cdot \sum_t P^T(t-1,t)\cdot P(t,t+1:.,j) \tag{9}$$

where $C^R(.,j)$ and $P(t,t+1:.,j)$ denote the *j-th* column of C^R and $P(t,t+1)$, respectively.

With our estimations in hand, then we analyse changes in the relative attractiveness of provinces. Dark-shaded areas in Figures 6a and 6b represent Spanish provinces with a positive column sum of the estimated right-causative matrices over the periods 2002-2007 (Figure 6a) and 2008-2013 (Figure 6b), respectively.

Figure 6. Spanish provinces with increased relative attractiveness



(a) 2002-2007 (b) 2008-2013

Note: Dark-shaded areas correspond to provinces with increased relative attractiveness.

If we compare both figures we see that the spatial configuration regarding changes in relative attractiveness has changed. The fact that 16 out of every one of the 25

Spanish provinces that have dealt with the economic crisis best[8] (Almería, Segovia, Las Palmas, Granada, Málaga, Huelva, Santa Cruz de Tenerife, Alicante, Gerona, Badajoz, Baleares, Palencia, Vizcaya, Ávila, Madrid, Murcia) appear in the group of Spanish provinces with gains in relative attractiveness over the 2008-2013 period is probably the most remarkable feature to point out. In addition, with the exception of Toledo the remaining provinces (with gains of relative attractiveness but resisting the crisis worse than the national average) are characterised by reaching unemployment rates below the national average over the crisis period. This evidence seems to support results reached by the study of Minondo et al. (2013) who conclude that the Spanish provinces which resist the impact of the economic crisis best arise like main destination for internal migrants over the crisis period.

4 Concluding remarks

This paper examined how the recent economic crisis has affected the internal migration dynamics across the Spanish provinces. With the aim to gain an initial insight into this issue, we first examined the evolution of internal migration movements over the 2002-2013 period. Then, the study applied a distribution approach to obtain a more precise picture of changes occurred in this dynamics: after examining changes in the external shape of the distribution and in mobility within it, we paid attention to changes in relative attractiveness of provinces for migrants. The paper argues as main conclusions: 1) a significant fall in the intensity of internal migration across the Spanish provinces, 2) a more sharply fall of intra-provincial movements over the crisis period than inter-provincial migration, 3) a severe drop of net migration rates over time, while a significant increase in the number of provinces with a negative, albeit low, internal migration balance, 4) a prominent fall in distribution sparseness, 5) an outstanding rise in intra-distribution mobility over the crisis period (mobility is more than two-fold higher than in the previous period), and, finally, 6) that most provinces which resist the impact of the economic crisis

[8] Due to data availability for the whole period 2008-2013 and the severe impact of falling economic activity in Spain on unemployment, we use the growth in the unemployment rate between 2008 and 2013 to address the ability of provinces to deal with the economic crisis.

best arise like provinces with gains in relative attractiveness for migrants. In short, our findings provide further empirical support for the intuitive belief that migratory movements continue to be influenced, apart from residential or quality of life factors, by the provinces' economic environment.

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