Forecasting compositional risk allocations

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Abstract

We analyse models for panel data that arise in risk allocation problems, when a given set of sources are the cause of an aggregate risk value. We focus on the modeling and forecasting of proportional contributions to risk. Compositional data methods are proposed and the regression is flexible to incorporate external information from other variables. We guarantee that projected proportional contributions add up to 100%, and we introduce a method to generate confidence regions with the same restriction. An illustration using data from the stock exchange is provided.

JEL classification: C02, G22, D81.

Keywords: Simplex, capital allocation, dynamic management.

1 Introduction

What managers have in mind when they analyse the aggregation of risks is the "diversification effect". The different sources of risk compensate each other and the overall risk is lower than what would result from considering the sources separately.

The calculus of the contribution of each source of risk to the aggregate result corresponds exactly to the concept of *risk allocation*. An allocation assigns a percent value to each component, in such a way that the contribution of that source to the aggregated value is reflected by the percentage. Naturally, risk evolves over time and so does aggregate risk as well as proportional contributions to risk. For instance, an investment in three stocks can have the largest risk in one of them at the beginning of the year, but the main source of risk may be caused

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by another stock at the end of the year. A major problem when forecasting allocations is that the sum of contributions to the aggregate risk has to be equal to 100%. This is a constraint that prediction must fulfill.

On one-period horizon, risk capital allocations have been studied intensively in the literature, with for instance game-theoretic and optimization-based approaches (see, among others, Tasche, 1999; Denault, 2001; Buch et al., 2011; Dhaene et al., 2012; Zaks and Tsanakas, 2014; Boonen et al., 2017). Risk capital allocations problems emerge when a financial firm holds an aggregate capital determined by the regulator, and needs to allocate this capital to the business units. In this paper, our focus differs from the conventional literature in two fundamental ways. First, we are interested in the dynamics of allocations over time. Second, we do not focus on the absolute value of capital allocations, but we are interested in the proportional contributions (*compositions*), i.e. amounts expressed in percentages with the restriction that they have to add up to 100. Capital allocations fluctuate over time due to economic developments, but we abstract from this and model explicitly how the relative allocations to sources of risk evolve over time. We refer to the relative allocations as risk allocations.

We claim that the methods presented here are useful to analyse the past behaviour of risk allocations and to predict future values together with confidence intervals. This aspect has not been considered before by previous authors. The analysis is then flexible enough to analyse a panel of n different sources of risk and their corresponding contributions to the overall result along time, that is over T periods. We also consider the possibility of introducing external information from other variables, such as market uncertainty, the external evolution of the economy or, other indicators. Our results facilitate decision-making processes in enterprise management that need to consider the dynamics of risk allocations. This is simply done, by means of the compositional panel data regression that incorporates other external variables.

The dynamics of allocations are modeled using compositional time series regressions. Modeling multivariate time series of compositional data have gained some attention in last years (Kynclová et al., 2015; Snyder et al., 2017; Perez-Foguet et al., 2017). These previous studies focused on the expected projections. However, a major issue in risk management is the analysis of possible deviations to expectations. We propose a methodology to compute compositional confidence intervals of projections over time. Projected compositions over these confidence intervals provide valuable information to risk managers of allocations under less likely scenarios.

Applications of compositional techniques in risk-management are scarcely found in the literature. Belles-Sampera et al. (2016) show the connection between capital allocation principles and compositional data for a one-period horizon. The authors apply compositional techniques to compute the distance between a set of risk allocation solutions and to compute the distance between them, but they do not address forecasting problems. Verbelen et al. (2016) study compositional re-

gression models in motor insurance pricing, but they do not study the time series and forecasting.

The article is organized as follows. The vector space structure of compositional data is introduced in the next section. In Section 3, risk allocation models are described. The regression model and the confidence regions for compositional data are presented in Section 4. Section 5 shows the numerical illustration. Concluding remarks can be found in Section 6.

2 The clr and ilr operations

In this section, we review notation and operators for compositional data. Unless mentioned otherwise, we assume that vectors have length $n > 1$. The set of compositions is denoted by the following simplex:

$$\mathbb{S}^n = \Big\{ \boldsymbol{x} = (x_1, \ldots, x_n) : x_i \geq 0, \sum_{i=1}^n x_i = 1 \Big\}. \qquad (2.1)$$

For econometric modelling, it is not convenient that the range of compositional data is constrained. Therefore, we transform compositional data to allow for the range $\mathbb{R}$ per coordinate. In this way, we can construct a regular econometric model to forecast the compositions. The first isometric transformation we use is the centered log-ratio transformation (clr)

$$clr(\boldsymbol{x}) = \Big(\ln \frac{x_i}{g(x)} \Big)_{i=1,\ldots,n}, \text{ with } g(x) = \Big(\prod_{i=1}^n x_i \Big)^{1/n}. \qquad (2.2)$$

where the log ratio of the vector is applied component-wise. The inverse clr transformation is straightforward: if $x^* = clr(x)$, then $x = C[\exp(x^*)]$, where the exponential is applied by components and where C is the closing function:

$$C(\boldsymbol{x}) = \frac{\boldsymbol{x}}{\sum_{i=1}^n x_i}.$$

By definition, the clr transformed components sum up to zero. This may be a source of problems when doing statistical analysis, as, e.g., the variance matrix of a clr-transformed composition is singular (van den Boogaart and Tolosana-Delgado, 2013).

Note that $clr(\boldsymbol{x}) \in \mathbb{R}^n$. We define the isometric logratio transformation, denoted by $\mathrm{ilr}(\boldsymbol{x}) \in \mathbb{R}^{n-1}$, as follows. There exist an isometric linear mapping between the simplex $\mathbb{S}^n$ and the $(n-1)$-dimensional Euclidean space $\mathbb{R}^{n-1}$. This mapping is called the isometric log-ratio transformation (ilr).

The isometry is constructed by representing the result in a basis of the $(n-1)$-dimensional image space of the clr transformation. This is constructed by taking an orthonormal basis of $\mathbb{R}^n$ including the vector $v_n = (1, \ldots, 1)$, i.e., some $(n-1)$ linearly independent vectors $\{v_1^*, \ldots, v_{n-1}^*\}$. In computational terms, one can arrange the vectors

$\{v_j^*\}$ by columns in a $n \times (n-1)$-element matrix, denoted by V (a *contrast* matrix), with the following properties:

- V is a quasi-orthonormal matrix, as

$$V^T \cdot V = 1_{n-1}, \text{ and } V \cdot V^T = 1_n - \tfrac{1}{n}1_{n \times n}, \qquad (2.3)$$

  where 1_n is the $n \times n$ identity matrix, and $1_{n \times n}$ is a $n \times n$ matrix full of ones.

- The columns of V sum up to zero,

$$1 \cdot V = 0, \qquad (2.4)$$

  where 1 indicates a row vector of ones.

Thanks to these properties, we can find simple expressions to have an equivalence between coordinates ζ and compositions x:

$$clr(x) \cdot V = \ln(x) \cdot V := \text{ilr}(x),$$
$$\text{ilr}^{-1}(\zeta) = C[\exp(\zeta \cdot V^T)].$$

Through these expressions, we also defined explicitly the so-called isometric logratio transformation (ilr): this is nothing else than a transformation that provides the coordinates of any composition with respect to a given orthonormal basis. There are as many ilr as orthonormal basis can be defined, as well as matrices V satisfying (2.3) and (2.4).

Note that the operation ilr depends on V, and in the sequel of this paper we will fix V as follows. The sequential binary partition proposed by Egozcue et al. (2003) has been broadly used because it allows the interpretation in terms of grouped parts of the composition. To create the first coordinate, the parts of the composition are divided into two groups. Parts in the first group are coded $+1$ and placed in the numerator. Parts in the second group are coded -1 and placed in the denominator. The first coordinate describes the balance between these two groups. In the second and following steps, each group is in turn divided into two groups. The splitting process of a group stops when it only has one component, which is coded 0 in the subsequent steps.

Each coordinate describes the balance between two groups: $+1$ parts and -1 parts. For this reason, coordinates are also called balances. In step k, coordinate $\text{ilr}(x)_k$ (balance z_k) is created as follows. The r_k parts of the first group and s_k parts of the second group have elements $(x_{n_1}, x_{n_2}, ..., x_{n_{r_k}})$ and $(x_{d_1}, x_{d_2}, ..., x_{d_{s_k}})$, respectively, where r_k indicates the number of parts that takes value $+1$ and s_k the number of parts that takes value -1 in step k. The isometric log-ratio transformation $\text{ilr}(\boldsymbol{x})_k$ is given by

$$z_k = \text{ilr}(\boldsymbol{x})_k = \sqrt{\frac{r_k \cdot s_k}{r_k + s_k}} \ln \frac{(x_{i_1} \cdots x_{i_{r_k}})^{1/r_k}}{(x_{j_1} \cdots x_{j_{s_k}})^{1/s_k}}, \quad k = 1, \ldots, n-1.$$
$$(2.5)$$

The number of steps to obtain a single component in each group is exactly $n-1$.

3 Allocation models

An allocation problem arises when an amount associated to the whole has to be distributed among its parts. Many previous studies deal with the analysis of capital allocation problems in risk management where an economic amount is fixed to cover the risk (see, among others, Tsanakas, 2004; Kalkbrener, 2005; Dhaene et al., 2012; Tsanakas, 2009; Urbina and Guillen, 2014; Zaks and Tsanakas, 2014). In this section, we briefly discuss two well-known allocation methods: the proportional rule and the Euler rule.

We assume that at time t the amount to share among parts is associated with the positive random variable $Y_{t+1}|\mathcal{F}_t$, where Y_{t+1} indicates the (unknown) value of the aggregate investment at time $t+1$ and $\mathcal{F}_t$ is the information set at time t. Positive random variable in risk management can also indicate loss severuty, for example. Let us assume that the aggregate investment consists of n different assets, where $Y_{i,t+1}$ is the positive random variable indicating the investment value of the i-th asset at time $t+1$, i.e. $Y_{t+1} = \sum_{i=1}^n Y_{i,t+1}$.

The value to allocate among the n parts is computed by means of a risk measure, ρ. A risk measure ρ maps random variables into real numbers. We want to allocate a strictly positive amount to the n assets that are $\mathcal{F}_t$-measurable. The proportional risk allocation criterion is frequently followed to allocate the risk value among the parts. At time t the risk proportion allocated to asset i is given by

$$\frac{\rho(Y_{i,t+1}|\mathcal{F}_t)}{\sum_{j=1}^n \rho(Y_{j,t+1}|\mathcal{F}_t)}. \tag{3.1}$$

Here, $\rho(Y_{i,t+1}|\mathcal{F}_t)$ is the value that the risk measure assigns at time t to the investment value of the i-th asset in next period. Obviously, the sum of risk proportions allocated to all assets is equal to one. The main drawback of the simple risk allocation principle defined in (3.1) is that it does not take into account dependencies between parts. To take into account dependency, the Euler risk allocation method can be used. Following the Euler allocation (Tasche, 1999; Denault, 2001), at time t the risk proportion allocated to asset i is given by

$$\frac{\frac{\partial}{\partial \lambda_i}\rho\big(\sum_{j=1}^n \lambda_j Y_{j,t+1}|\mathcal{F}_t\big)\big|_{\lambda=(1,\ldots,1)}}{\rho(Y_{t+1}|\mathcal{F}_t)}, \tag{3.2}$$

where λ_i indicates the number of units of the i-th asset, $i = 1, \ldots, n$. The expression $\frac{\partial}{\partial \lambda_i}\rho\big(\sum_{j=1}^n \lambda_j Y_{j,t+1}|\mathcal{F}_t\big)\big|_{\lambda=(1,\ldots,1)}$ can be interpreted as the partial contribution of $Y_{i,t+1}$ to the value that the risk measure assigns to the aggregate random variable Y_{t+1}. If the risk measure is positively homogeneous, it holds that (Tasche, 1999):

$$\rho(Y_{t+1}|\mathcal{F}_t) = \sum_{i=1}^n \frac{\partial}{\partial \lambda_i}\rho\big(\sum_{i=1}^n \lambda_i Y_{i,t+1}|\mathcal{F}_t\big)\big|_{\lambda=(1,\ldots,1)}.$$

Therefore, Euler allocation involving positively homogeneous risk measures guarantees full risk allocation, i.e. the sum of risk proportions

obtained by expression (3.2) is equal to one.

Some popular examples of positively homogeneous risk measures are the Value-at-Risk (VaR) and the conditional Value-at-Risk (Buch et al., 2011; Cossette et al., 2012; Asimit et al., 2011). Let us consider $\alpha \in (0,1)$. The Value-at-Risk measure at confidence level α is defined as

$$\rho(Y_{i,t+1}|\mathcal{F}_t) = \text{VaR}_\alpha(Y_{i,t+1}|\mathcal{F}_t) = \sup\left\{y \mid F_{Y_{i,t+1}|\mathcal{F}_t}(y) \leq 1 - \alpha\right\}.$$

If the cumulative distribution function $F_{Y_{i,t+1}|\mathcal{F}_t}$ is continuous and strictly increasing then $\text{VaR}_\alpha(Y_{i,t+1}|\mathcal{F}_t) = F^{-1}_{Y_{i,t+1}|\mathcal{F}_t}(1 - \alpha)$, where F^{-1} is the quantile distribution function. The conditional Value-at-Risk can be understood as the expectation that conditional on the event that the risk exceeds the VaR. In this article, the Value-at-Risk risk measure is considered for the application. The chosen allocation method is the one based on the Euler rule.

4 Method

We show our regression model in Section 4.1, and our method to construct confidence intervals of our predictions in Section 4.2.

4.1 Regression model

Let $\boldsymbol{X}_t$ be a composition for all t. We assume that there are no zeros in the compositional data vector $\boldsymbol{X}_t$. We build a regression model with compositional response[1]:

$$\text{ilr}(\boldsymbol{X}_t) = \boldsymbol{b} + B \cdot \text{ilr}(\boldsymbol{X}_{t-1}) + \varepsilon_t, \tag{4.1}$$

with ilr the ilr transformation, $\boldsymbol{b}$ is a vector of parameters in $\mathbb{R}^{n-1}$, B is an $(n-1) \times (n-1)$ matrix of parameters in the real space, and $\varepsilon_t \sim N_{n-1}(\boldsymbol{0}, \Sigma_\varepsilon)$. Point predictions of compositions are invariant on the choice of the orthonormal basis for the isometric logratio transformation (Snyder et al., 2017).

Secondly, we propose the following model:

$$\text{ilr}(\boldsymbol{X}_t) = \boldsymbol{b} + B \cdot \text{ilr}(\boldsymbol{X}_{t-1}) + \gamma \cdot \boldsymbol{W}_t + \varepsilon_t, \tag{4.2}$$

where $\boldsymbol{W}_t$ is a set of control variables (vector of dimension r) and γ is the $(n-1) \times r$ matrix of coefficients defined in the real space. This is a vector auto-regression model with exogenous variables. We include only one control variable. We include the total amount to allocate among the parts $W_t = \rho(Y_{t+1}|\mathcal{F}_t)$, i.e. the total risk measure value of the investment at time $t + 1$ evaluated at time t.

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[1]For a geometric interpretation on the simplex $\mathbb{S}^n$, we refer to Kynclová et al. (2015). Such a geometric interpretation requires the introduction of special perturbation and powering operators for compositions (Aitchison and Egozcue, 2005).

For forecasting $\boldsymbol{X}_t$, we need to forecast W_t as well. We propose to forecast W_t with an AR(1) model, i.e.,

$$W_t = \alpha + \beta W_{t-1} + \zeta_t,$$

with $\zeta_t \sim N(0, \sigma_\zeta^2)$.

Finally, we compare the regression model with compositional data with a naive model. The naive model is an AR(1) model:

$$\boldsymbol{X}_t = \boldsymbol{c} + D \cdot \boldsymbol{X}_{t-1} + \varepsilon_t, \tag{4.3}$$

where c is an $(n \times 1)$ vector and D an $(n \times n)$ diagonal matrix, both defined in the real space, and ε_t has an n-dimensional multivariate normal distribution with expectation $\boldsymbol{0}$ and covariance matrix Σ_ε. Note that in this model, the forecasts do not need to be in $\mathbb{S}^n$, i.e., the forecast may not be compositional.

4.2 Confidence interval of predictions

We use bootstrapping methods to compute confidence intervals of predictions. We first simulate the predictive distribution for the next r periods. At time t, the algorithm design to the next period (the first unobserved period) is as follows:

1. We simulate S matrices of estimated coefficients $\hat{B}^s$, based on $N(\hat{B}, var(\hat{B}))$.

2. We estimate $\mathrm{ilr}(\hat{\boldsymbol{X}}_{t+1})^s = \mathrm{ilr}(\boldsymbol{X}_t) \cdot \hat{B}^s$ for $s = 1, ..., S$.

3. We simulate S error matrices ε^s based on $N_{n-1}(\boldsymbol{0}, \Sigma_\varepsilon)$.

4. The prediction error is incorporated $\mathrm{ilr}(\tilde{\boldsymbol{X}}_{t+1})^s = \mathrm{ilr}(\hat{\boldsymbol{X}}_{t+1})^s + \varepsilon^s$ for $s = 1, ..., S$.

With this algorithm, S simulated forecasts of $\mathrm{ilr}(\tilde{\boldsymbol{X}}_{t+1})$ are obtained. Note that parameter uncertainty is included in step 2. Steps 3 and 4 include uncertainty associated to the process (prediction). This is a common way to simulate the predictive distribution by bootstrapping. The predictive distribution could be easily estimated, for instance, by means of non-parametric methods or fitting a parametric distribution. To obtain the predicted α-confidence region, we identify the α percent of the simulated forecasts with the highest Euclidean distance to the $E[\mathrm{ilr}(\tilde{\boldsymbol{X}}_{t+1})]$, where $E[\mathrm{ilr}(\tilde{\boldsymbol{X}}_{t+1})] = \mathrm{ilr}(\boldsymbol{X}_t) \cdot \hat{B}$. Finally, the inverse of the ilr transformation is considered to expressed $\mathrm{ilr}(\tilde{\boldsymbol{X}}_{t+1})$ as projected compositions, $\tilde{\boldsymbol{X}}_{t+1}$.

Similarly, the algorithm design to the remaining periods is as follows:

1. We simulate S matrices of estimated coefficients $\hat{B}^s$, based on $N(\hat{B}, var(\hat{B}))$.

2. We estimate $\mathrm{ilr}(\hat{\boldsymbol{X}}_{t+r})^s = \mathrm{ilr}(\tilde{\boldsymbol{X}}_{t+r-1})^s \cdot \hat{B}^s$ for $s = 1, ..., S$ and $r > 1$.

3. We simulate S error matrices ε^s based on $N_{n-1}(\mathbf{0}, \Sigma_\varepsilon)$.

4. The prediction error is incorporated $\mathrm{ilr}(\tilde{\boldsymbol{X}}_{t+r})^s = \mathrm{ilr}(\hat{\boldsymbol{X}}_{t+r})^s + \varepsilon^s$ for $s = 1, ..., S$.

At any projected period, $t+r$ with $r > 1$, the predicted α-confidence regions are again obtained identifying the α percent of the simulated forecasts with the highest Euclidean distance to the $E[\mathrm{ilr}(\tilde{\boldsymbol{X}}_{t+r})]$, where $E[\mathrm{ilr}(\tilde{\boldsymbol{X}}_{t+r})] = E[\mathrm{ilr}(\tilde{\boldsymbol{X}}_{t+r-1})] \cdot \hat{B}$.

5 Numerical illustration

In this section, we provide the main econometric analysis of this paper based on the data as described in Section 5.1. The two main models that we study are (4.1) and (4.2), which we will compare with the naive model in (4.3). Results were obtained in the programming language R, with packages *vars* (Pfaff and Stigler, 2013), *robCompositions* (Templ et al., 2017), and *compositions* (van den Boogaart et al., 2015).

5.1 Data

We assume that $n = 3$, and that an investor selects the S&P 500 index, the DAX index, and the CAC 40 index, respectively. We use these indexes from January 1st, 2000, to December 31st, 2016, and monitor for every working day their daily value. We consider a buy-and-hold portfolio with an initial equal investment to each index on January 1st, 2000. The original aggregate investment in the portfolio was 100,000 dollars. The daily value of the investment in each index is considered a source of risk, and the aggregated daily value of the portfolio has an overall risk, which is displayed in Figure 1.

The time series is divided in monthly periods. We calculate every month (20 workings days) the 95%-Value-at-Risk of the empirical distribution of the aggregate daily portfolio value. In this manner, we construct the compositional data, which consists of 210 compositions (x_1, x_2, x_3).

First, we assume that the risk allocation is determined by the Euler rule, as defined in (3.2). In this way, we construct the compositional risk allocation. We display the compositional data in Figure 2. The dynamics of risk allocations over time are shown. We observe the decreasing evolution of the third index risk contribution, which is the CAC 40 index, so this index is the least risky one. The classical form to display three-part compositions is by means of ternary diagrams. We provide a ternary diagram of the compositions in the simplex $\mathbb{S}^3$ in Figure 3.

We now express the compositional data with respect to an orthonormal basis of the underlying vector space with Euclidean structure. We use the sequential binary partition shown in (2.5). Table 1 shows the choice of the sequential binary partition for this example and Table 2
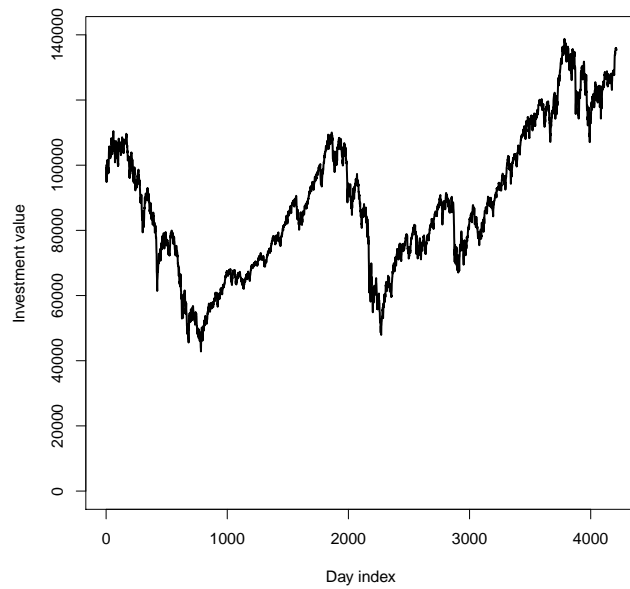
Figure 1: Evolution of the aggregate daily value of a portfolio that invests 100,000 units in the S&P 500 index, the DAX index, and the CAC 40. The investment is equally split in the three components on January 1st, 2000 and the investment goes on until December 31st, 2016.
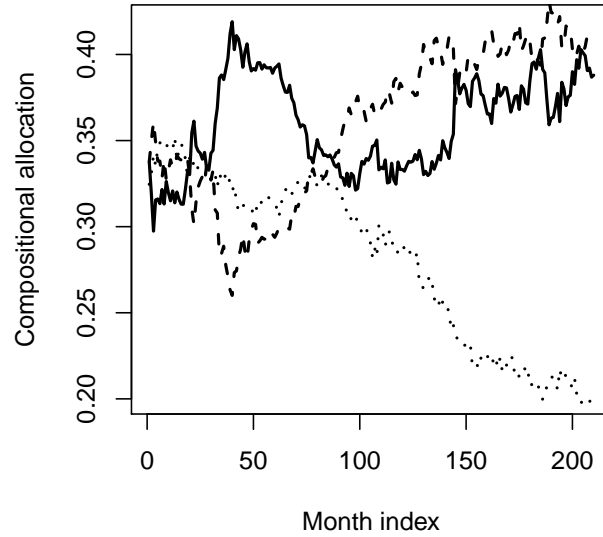
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Figure 2: Evolution of the compositional data that capture the relative contribution of each index to the overall risk. Here, the solid line is the relative risk due to the S&P 500 index, the dashed line is the relative risk due to the DAX index, and the dotted line is the relative risk caused by the CAC 40.
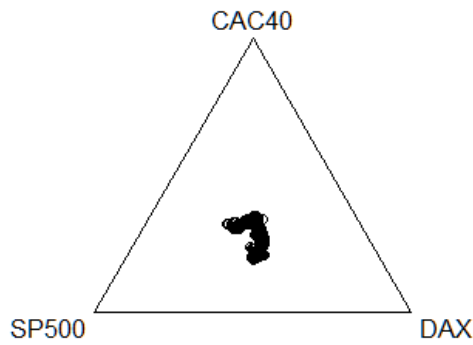

Figure 3: A ternary diagram of the compositional data in our numerical illustration.

shows the associated contrast matrix of this sequential binary partition.

Order	x_1	x_2	x_3	r	s
z_1	+1	-1	-1	1	2
z_2	0	+1	-1	1	1

Table 1: A 3-dimensional sequential binary partition.

x_1	x_2	x_3
$\sqrt{\frac{2}{3}}$	$-\frac{1}{2}\sqrt{\frac{2}{3}}$	$-\frac{1}{2}\sqrt{\frac{2}{3}}$
0	$\sqrt{\frac{1}{2}}$	$-\sqrt{\frac{1}{2}}$

Table 2: The contrast matrix V^T associated with the sequential binary partition in Table 1.

This particular choice of sequential binary partition leads to the following balances,

$$z_1 = \sqrt{\frac{2}{3}} \ln \frac{x_1}{\sqrt{x_2 \cdot x_3}}, \qquad z_2 = \frac{1}{\sqrt{2}} \ln \frac{x_2}{x_3}. \qquad (5.1)$$

With this transformation, balance z_1 contains relative information (ratio) of the part x_1 to the parts x_2 and x_3, and balance z_2 only relative information of the part x_2 to the part x_3. Note that x_1, x_2 and x_3 indicate the proportion of the 95% VaR risk portfolio value associated to the investment in the S&P 500 index, the DAX index, and in the CAC 40 index, respectively.

5.2 Coefficient estimates

Table 3 shows the parameter estimates for the models in (4.1), (4.2), and (4.3). The coefficient associated with the lag is statistically significant at the 0.05 significance level in all three model specifications (original scale or ilr transformation). So, information of the value of the balance (or the original scaled part) at time t has explanatory capacity on the estimation of its value at time $t + 1$.

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Parameter	Model (4.1)	Model (4.2)	Model (4.3)
b_1 or c_1	0.0034 (0.248)	0.0251 (0.009)*	0.0146 (0.041)*
b_2 or c_2	0.0017 (0.404)	-0.0066 (0.320)	0.0036 (0.356)
c_3			-0.0007 (0.713)
$B_{1,1}$	0.9590 (0.000)*	0.9503 (0.000)*	
$B_{2,1}$	0.0119 (0.228)	0.0324 (0.013)*	
$B_{1,2}$	0.0072 (0.595)	0.0105 (0.443)	
$B_{2,2}$	0.9994 (0.000)*	0.9915 (0.000)*	
$D_{1,1}$			0.9600 (0.000)*
$D_{2,2}$			0.9911 (0.000)*
$D_{3,3}$			1.0003 (0.000)*
γ_1		$-2.874 \cdot 10^{-7}$ (0.018)*	
γ_2		$1.097 \cdot 10^{-7}$ (0.191)	

Table 3: OLS parameter estimates of Models (4.1), (4.2), and (4.3) and their p-values (in brackets). * refers to significance at the 0.05 significance level.

An interesting result is observed with the output of model (4.2) is analysed. Coefficient estimate of the balance z_1 is statistically significant at the 0.05 significance level when the dependent variable is the balance z_2. According to this result, the relative information (ratio) of the part x_1 to the parts x_2 and x_3, at time t would be relevant to estimate the relative information of the part x_2 to the part x_3 at time $t + 1$. That is, the proportion of the 95% VaR risk portfolio value associated to the investment in the S&P 500 index in relation to the investment in the DAX index and in the CAC 40 index would be informative to estimate in the next month the proportion of the 95% VaR risk portfolio value associated to the investment in the DAX index in relation to the investment in the CAC 40 index. Note that this coefficient estimate is not significant at the 0.05 significance level in model (4.1).

Additionally, we find that the exogenous variable W_t in Model (4.2) has a statistically significant parameter estimate at the 0.05 significance level. This result would suggest that information of the size of the 95% VaR risk portfolio value at time t has explanatory capacity to estimate the value of the balance z_1 at time $t + 1$, i.e. the VaR risk value of the portfolio would be informative to estimate in next period the proportion of the 95% VaR risk portfolio value associated to the investment in the S&P 500 index in relation to the investment in the DAX and CAC 40 indices.

5.3   Model selection

First, we compare the models in (4.1) and (4.2). Both models are Vector Autoregressive (VAR) models, with or without an exogenous variable. For such models, we can include more lags. For model selection, we consider the following two well-known information criteria,

Akaike's information criterion (AIC, Akaike 1973) and the Bayesian information criterion (BIC, Schwarz 1978):

$$\text{AIC} = -2\ell + 2K,$$
$$\text{BIC} = -2\ell + \ln(N)K,$$

where ℓ is the log-likelihood with the estimated parameters, K the number of parameters, and N the number of observations. The model with the smallest AIC/BIC is selected. Yang (2005) explains the differences between the AIC and BIC and van den Boogaart and Tolosana-Delgado (2013) argue that the use of the AIC and BIC levels for model selection with compositional data is suitable. The AIC and BIC are invariant to the choice of the contrast matrix V for the ilr transformation (Kynclová et al., 2015).

For Model (4.1) and Model (4.2), we find that adding more lags leads to a higher AIC and BIC, and so the formulation in (4.1) and (4.2) with only one lag is preferred. We display the AIC and BIC levels for the models in (4.1) and (4.2) with one lag in Table 4, and find that the AIC and BIC levels do not provide the same model selection conclusion. Therefore, we will continue with studying both models.

	Model (4.1)	Model (4.2)
Log-likelihood ℓ	989.2	992.4
AIC	-1,966.3	-1,968.8
BIC	-1,946.3	-1,942.1

Table 4: The log-likelihood, AIC and BIC levels for Models (4.1) and (4.2).

We continue with forecasting our models. We display the point forecasts for 10 periods in Table 5. These forecasts confirm that for Model 4.3 the forecasts do not need to be compositional.

In the case of the naive model (Model (4.3)), we observe than the sum of forecasts are not equal to the unity for any of the projected periods. Therefore, the aggregate risk is not allocated among the risk units. Note that the distance between the aggregated forecasts and the unity increases as the projected period does. It is due to the fact the error is accumulated with time. So, forecasts should be considered with extreme caution for far-off projections under this naive approach.

Next, we test our forecast accuracy via backtesting. Recall that our data-set contains 210 periods. For a jump-off period $k \in \{200, \dots, 209\}$, we study the forecast performance using the root mean squared forecast error (RMSFE):

$$\text{RMSFE}(k, m) = \sqrt{\frac{1}{210 - k} \sum_{t=k+1}^{210} \|\boldsymbol{X}_t - \hat{\boldsymbol{X}}_t^{k,m}\|^2}, \qquad (5.2)$$

where $\hat{\boldsymbol{X}}_t^{k,m}$ is the at time k-forecasted composition using model m for the period t (see Table 5). The forecast accuracy is best for the model with the lowest RMSFE.

Period	S&P 500	DAX	CAC 40	Sum
Model (4.1)				
$t+1$	0.3872	0.4130	0.1998	1
$t+2$	0.3864	0.4128	0.2009	1
$t+3$	0.3856	0.4125	0.2019	1
$t+4$	0.3849	0.4122	0.2029	1
$t+5$	0.3842	0.4119	0.2039	1
$t+6$	0.3835	0.4116	0.2049	1
$t+7$	0.3829	0.4113	0.2058	1
$t+8$	0.3823	0.4110	0.2068	1
$t+9$	0.3817	0.4106	0.2077	1
$t+10$	0.3811	0.4103	0.2086	1
Model (4.2)				
$t+1$	0.3856	0.4159	0.1985	1
$t+2$	0.3835	0.4183	0.1981	1
$t+3$	0.3816	0.4207	0.1978	1
$t+4$	0.3798	0.4228	0.1974	1
$t+5$	0.3782	0.4248	0.1969	1
$t+6$	0.3768	0.4267	0.1965	1
$t+7$	0.3755	0.4285	0.1960	1
$t+8$	0.3743	0.4301	0.1955	1
$t+9$	0.3733	0.4317	0.1950	1
$t+10$	0.3724	0.4331	0.1945	1
Model (4.3)				
$t+1$	0.3871	0.4131	0.1981	0.9984
$t+2$	0.3862	0.4130	0.1975	0.9968
$t+3$	0.3854	0.4129	0.1969	0.9952
$t+4$	0.3846	0.4128	0.1963	0.9937
$t+5$	0.3838	0.4127	0.1956	0.9922
$t+6$	0.3831	0.4126	0.1950	0.9907
$t+7$	0.3824	0.4125	0.1944	0.9893
$t+8$	0.3817	0.4123	0.1938	0.9878
$t+9$	0.3811	0.4122	0.1931	0.9865
$t+10$	0.3805	0.4121	0.1925	0.9851

Table 5: Point forecasts of the Model (4.1), (4.2), and (4.3) for 10 periods. Here, $t = 210$ is the last sample period.

We show the RMSFE for all values of $k \in \{200, \ldots, 209\}$ in Table 6. We find that Model (4.3) never has the lowest RMSFE for all jump-off periods. Hence, using a compositional model improves the forecast accuracy. There is however mixed evidence about the model selection for Models (4.1) and (4.2).

k	Model (4.1)	Model (4.2)	Model (4.3)
209	0.0023	0.0040	0.0028
208	0.0051	0.0034	0.0054
207	0.0036	0.0060	0.0038
206	0.0107	0.0062	0.0115
205	0.0110	0.0061	0.0126
204	0.0112	0.0048	0.0133
203	0.0087	0.0136	0.0105
202	0.0072	0.0095	0.0107
201	0.0185	0.0248	0.0199
200	0.0108	0.0144	0.0132

Table 6: The RMSFE(k, m) for various jump-off periods k and models m, that are (4.1), (4.2), or (4.3).

A frequent task in the context of multivariate time series analysis is testing for causality. For that reason, Granger causality was introduced, which represents a statistical concept that is based on prediction. This is a pairwise time series analysis that indicates if past values of a time series contain information to predict future values of another time series. In other words, we are interested in testing whether one variable could help to improve predictions of the variable of interest. This is done by including the variable with several lags, and test whether it improves statistically the predictions of the variable of interest. We check Granger causality of our ilr-transformed data and the aggregate risk level W_t in Table 7. We only find statistical evidence that ilr$(\boldsymbol{X}_t)_1$ Granger causes Z_t. So, effects of W_t are partially caused by changes in the composition ilr$(\boldsymbol{X}_t)$, but not the other way around. Therefore, we have no evidence that W_t has an explanatory effect on ilr$(\boldsymbol{X}_t)$. That is, information of the proportion of the 95% VaR risk portfolio value associated to the investment in the S&P 500 index in relation to the investment in the DAX and CAC 40 indices is informative to predict the size of the 95% VaR risk portfolio. The inverse does not hold.

To sum up, we find no strong indicators to prefer Model (4.2) above Model (4.1). Moreover, Model (4.3) provides no compositional forecasts.

Null hypothesis	p-value
$\mathrm{ilr}(\boldsymbol{X}_t)_1$ does not Granger cause $\mathrm{ilr}(\boldsymbol{X}_t)_2$	0.595
$\mathrm{ilr}(\boldsymbol{X}_t)_2$ does not Granger cause $\mathrm{ilr}(\boldsymbol{X}_t)_1$	0.228
$\mathrm{ilr}(\boldsymbol{X}_t)_1$ does not Granger cause W_t	0.007*
$\mathrm{ilr}(\boldsymbol{X}_t)_2$ does not Granger cause W_t	0.098
W_t does not Granger cause $\mathrm{ilr}(\boldsymbol{X}_t)_1$	0.943
W_t does not Granger cause $\mathrm{ilr}(\boldsymbol{X}_t)_2$	0.942

Table 7: Tests for Granger causality of the compositional data, and the variable W_t.

5.4 Confidence interval of predictions

In this section, we study the confidence intervals of the forecasted compositional risk allocations. We apply the simulation approach of Section 4.2. We proceed our analysis by with focusing only on Model 4.1.

In Figure 4, we display a three-dimensional scatter-plot of compositional simulations for the ten projected periods[2]. Note that compositional simulations seems to be in the shape of ellipsoids with axes increasing with time projected periods. The center of the ellipsoids are the point compositional forecasts. We see that all simulations are compositional, as they are elements of $\mathbb{S}^3$. Moreover, the 5% simulated forecasts with the highest Euclidean distance to the point forecast in the simplex are shown in Figure 5.

Some descriptive statistics can be computed for simulated forecasts. In Table 8 we compare at each projection period the correlation between simulated compositional forecasts expressed in ilr coordinates (all simulated balances) and the correlation between the 5%-of simulated balances with the associated highest Euclidean distance (simulated balances in the 95% confidence region).

The estimated correlation is weaker as the forecasting period is farther in the future. This result is expected. Note that balance z_1 is based on the ratio of x_1 with x_2 and x_3 and balance z_2 on the ratio of x_2 with x_3. The projected value of z_1 provides less information on the projected value of z_2 as the projection is further in the future, and inverse.

In fact, a weak estimated correlation is observed in all periods when all simulations are considered. The proportion of the 95% VaR risk portfolio value associated to the investment in the S&P 500 index in relation to the investment in the DAX index and in the CAC 40 index would be weakly correlated with the simulated prediction of the proportion of the 95% VaR risk portfolio value associated to the investment in the DAX index in relation to the investment in the CAC 40 index. However, this conclusion varies when simulations in the confidence re-

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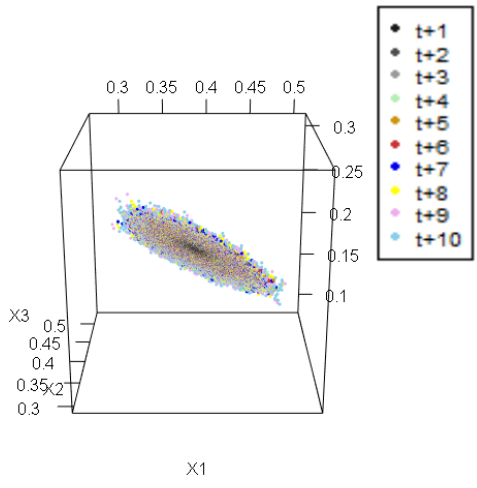[2]Alternatively, simulations could be represented in a two-dimensional contour ternary diagram.

Figure 4: A scatter-plot of the simulated compositional forecasts for ten projected periods. Here, $t = 210$ is the last sample period.
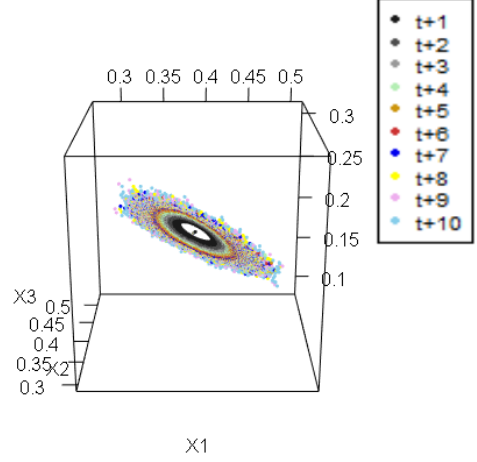

Figure 5: A scatter-plot of the 5% simulated compositional forecasts with the highest Euclidean distance for 10-projected periods. Here, $t = 210$ is the last sample period.

Forecasting period	Correlation (all simulations)	Correlation (5% simulations)
$t+1$	-0.200	-0.423
$t+2$	-0.199	-0.410
$t+3$	-0.187	-0.390
$t+4$	-0.180	-0.387
$t+5$	-0.170	-0.374
$t+6$	-0.162	-0.357
$t+7$	-0.153	-0.325
$t+8$	-0.144	-0.296
$t+9$	-0.136	-0.277
$t+10$	-0.125	-0.250

Table 8: Correlation between simulated balances for all simulations and for the 5% simulations with the highest Euclidean distance. Here, $t = 210$ is the last sample period.

gion are considered. In this case, a stronger correlation is observed in all periods (more than the double value). That is, the simulated projections of balances are more correlated in extreme scenarios.

6 Conclusion

We have presented a method that is useful to analyse panels of data that represent proportions of risk allocated to a give number of risk sources. There are many examples where this is applicable. Consider, for instance, a firm that has three plants located in different parts of the world and each of them producing losses or profits. The interest is in predicting the proportion of risk for each plant in the next period and to analyse which factors influence the compositional vector of proportions. The analysis can be done in certain periods of time, such as every month or every week.

A suggestion for future work is to include the possibility of a null contribution to risk. This would then also solve the problem of having a source of risk that suddenly disappears from the scene due to a default.

Aknowledgements

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Pitt, D., **Guillén, M.** (RFA-IREA)
"An introduction to parametric and non-parametric models for bivariate positive insurance claim severity distributions"
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Bermúdez, Ll. (RFA-IREA), **Karlis, D.**
"Modelling dependence in a ratemaking procedure with multivariate Poisson regression models"
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Di Paolo, A. (IEB)
"Parental education and family characteristics: educational opportunities across cohorts in Italy and Spain"
(Maig 2010)

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Simón, H. (IEB), **Ramos, R.** (AQR-IREA), **Sanromá, E.** (IEB)

"Movilidad ocupacional de los inmigrantes en una economía de bajas cualificaciones. El caso de España"
(Juny 2010)

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Di Paolo, A. (GEAP & IEB), **Raymond, J. Ll.** (GEAP & IEB)
"Language knowledge and earnings in Catalonia"
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Bolancé, C. (RFA-IREA), **Alemany, R.** (RFA-IREA), **Guillén, M.** (RFA-IREA)
"Prediction of the economic cost of individual long-term care in the Spanish population"
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Di Paolo, A. (GEAP & IEB)
"Knowledge of catalan, public/private sector choice and earnings: Evidence from a double sample selection model"
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Coad, A., **Segarra, A.** (GRIT), **Teruel, M.** (GRIT)
"Like milk or wine: Does firm performance improve with age?"
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Di Paolo, A. (GEAP & IEB), **Raymond, J. Ll.** (GEAP & IEB), **Calero, J.** (IEB)
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Borrell, A. (GiM-IREA), **Fernández-Villadangos, L.** (GiM-IREA)
"Clustering or scattering: the underlying reason for regulating distance among retail outlets"
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"School composition effects in Spain"
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Fageda, X. (GiM-IREA), **Flores-Fillol, R.**
"Technology, Business Models and Network Structure in the Airline Industry"
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Albalate, D. (GiM-IREA), **Bel, G.** (GiM-IREA), **Fageda, X.** (GiM-IREA)
"Is it Redistribution or Centralization? On the Determinants of Government Investment in Infrastructure"
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Oppedisano, V., **Turati, G.**
"What are the causes of educational inequalities and of their evolution over time in Europe? Evidence from PISA"
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Canova, L., **Vaglio, A.**
"Why do educated mothers matter? A model of parental help"
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Fageda, X. (GiM-IREA), **Perdiguero, J.** (GiM-IREA)
"An empirical analysis of a merger between a network and low-cost airlines"
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Moreno-Torres, I. (ACCO, CRES & GiM-IREA)
"What if there was a stronger pharmaceutical price competition in Spain? When regulation has a similar effect to collusion"
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Miguélez, E. (AQR-IREA); **Gómez-Miguélez, I.**
"Singling out individual inventors from patent data"
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Moreno-Torres, I. (ACCO, CRES & GiM-IREA)
"Generic drugs in Spain: price competition vs. moral hazard"
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Nieto, S. (AQR-IREA)**, Ramos, R.** (AQR-IREA)
"¿Afecta la sobreeducación de los padres al rendimiento académico de sus hijos?"
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"Estimation of Parametric and Nonparametric Models for Univariate Claim Severity Distributions - an approach using R"
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"How much risk is mitigated by LTC Insurance? A case study of the public system in Spain"
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"Loss risk through fraud in car insurance"
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Bermúdez, Ll. (RFA-IREA), **Karlis, D.**
"Mixture of bivariate Poisson regression models with an application to insurance"
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Varela-Irimia, X-L. (GRIT)
"Age effects, unobserved characteristics and hedonic price indexes: The Spanish car market in the 1990s"
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"A correlation sensitivity analysis of non-life underwriting risk in solvency capital requirement estimation"
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Guillén, M. (RFA-IREA), **Pérez-Marín, A.** (RFA-IREA), **Alcañiz, M.** (RFA-IREA)
"A logistic regression approach to estimating customer profit loss due to lapses in insurance"
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Segarra, A. (GRIT)
"R&D cooperation between Spanish firms and scientific partners: what is the role of tertiary education?"
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García-Pérez, J. I.; **Hidalgo-Hidalgo, M.**; **Robles-Zurita, J. A.**
"Does grade retention affect achievement? Some evidence from PISA"
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González-Val, R. (IEB), **Olmo, J.**
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"The Innovation and Imitation Dichotomy in Spanish firms: do absorptive capacity and the technological frontier matter?"
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"Evaluating Antitrust Leniency Programs"
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"Solvency capital estimation and risk measures"
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"How to use the standard model with own data"
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Perdiguero, J. (GiM-IREA), **Borrell, J.R.** (GiM-IREA)
"Driving competition in local gasoline markets"
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"Discrete time Non-homogeneous Semi-Markov Processes applied to Models for Disability Insurance"
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"Quantitative analysis of image factors in a cultural heritage tourist destination"
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"Sunk costs, extensive R&D subsidies and permanent inducement effects"
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Boj, E. (CREB), **Delicado, P.**, **Fortiana, J**., **Esteve, A.**, **Caballé, A.**
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"What about people in European Regional Science?"
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"Nonparametric estimation of Value-at-Risk"
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Herrera-Idárraga, P. (AQR-IREA), **López-Bazo, E.** (AQR-IREA), **Motellón, E.** (AQR-IREA)
"Informality and overeducation in the labor market of a developing country"
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"(Endogenous) occupational choices and job satisfaction among recent PhD recipients: evidence from Catalonia"
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"Social Determinants of Child Health in Colombia: Can Community Education Moderate the Effect of Family Characteristics?"
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Teixidó-Figueras, J. (GRIT), **Duró, J. A.** (GRIT)
"The building blocks of international ecological footprint inequality: a regression-based decomposition"
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Salcedo-Sanz, S., **Carro-Calvo, L.**, **Claramunt, M.** (CREB), **Castañer, A.** (CREB), **Marmol, M.** (CREB)
"An Analysis of Black-box Optimization Problems in Reinsurance: Evolutionary-based Approaches"
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Alcañiz, M. (RFA), **Guillén, M.** (RFA), **Sánchez-Moscona, D.** (RFA), **Santolino, M.** (RFA), **Llatje, O., Ramon, Ll.**
"Prevalence of alcohol-impaired drivers based on random breath tests in a roadside survey"
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"The Impact of Cooperation on R&D, Innovation andProductivity: an Analysis of Spanish Manufacturing and Services Firms"
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"Testing extreme value copulas to estimate the quantile"
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Solé-Auró, A. (RFA), **Alcañiz, M.** (RFA)
"Are we living longer but less healthy? Trends in mortality and morbidity in Catalonia (Spain), 1994-2011"
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Teixidó-Figueres, J. (GRIT), **Duro, J. A.** (GRIT)
"Spatial Polarization of the Ecological Footprint distribution"
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Cristobal-Cebolla, A.; Gil Lafuente, A. M. (RFA), **Merigó Lindhal, J. M.** (RFA)
"La importancia del control de los costes de la no-calidad en la empresa"
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Castañer, A. (CREB); **Claramunt, M.M.** (CREB)
"Optimal stop-loss reinsurance: a dependence analysis"
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"Job accessibility, employment and job-education mismatch in the metropolitan area of Barcelona"
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Di Paolo, A. (AQR-IREA); **Mañé, F.**
"Are we wasting our talent? Overqualification and overskilling among PhD graduates"
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Segarra, A. (GRIT); **Teruel, M.** (GRIT); **Bové, M. A.** (GRIT)
"A territorial approach to R&D subsidies: Empirical evidence for Catalonian firms"
(Setembre 2014)

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Ramos, R. (AQR-IREA); **Sanromá, E.** (IEB); **Simón, H.**
"Public-private sector wage differentials by type of contract: evidence from Spain"
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Bel, G. (GiM-IREA); **Bolancé, C.** (Riskcenter-IREA); **Guillén, M.** (Riskcenter-IREA)**; Rosell, J.** (GiM-IREA)
"The environmental effects of changing speed limits: a quantile regression approach"
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Bolance, C. (Riskcenter-IREA); **Bahraoui, Z.** (Riskcenter-IREA), **Alemany, R.** (Risckcenter-IREA)
"Estimating extreme value cumulative distribution functions using bias-corrected kernel approaches"
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Ramos, R. (AQR-IREA); **Sanromá, E.** (IEB), **Simón, H.**
"An analysis of wage differentials between full- and part-time workers in Spain"
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Cappellari, L.; **Di Paolo, A.** (AQR-IREA)
"Bilingual Schooling and Earnings: Evidence from a Language-in-Education Reform"
(Setembre 2015)

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Álvarez-Albelo, C. D., Manresa, A. (CREB), **Pigem-Vigo, M.** (CREB)
 "Growing through trade: The role of foreign growth and domestic tariffs"
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Caminal, R., Di Paolo, A. (AQR-IREA)
Your language or mine?
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Choi, H. (AQR-IREA)**, Choi, A.** (IEB)
When one door closes: the impact of the hagwon curfew on the consumption of private tutoring in the Republic of Korea

2016

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Castañer, A. (CREB, XREAP); **Claramunt, M M.** (CREB, XREAP), **Tadeo, A.**, **Varea, J.** (CREB, XREAP)
Modelización de la dependencia del número de siniestros. Aplicación a Solvencia II
(Setembre 2016)

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Local labor market effects of public employment
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Vayá, E. (AQR-IREA, XREAP)**, García, J. R.** (AQR-IREA, XREAP)**, Murillo, J.** (AQR-IREA, XREAP)**, Romaní, J.** (AQR-IREA, XREAP)**, Suriñach, J.** (AQR-IREA, XREAP)**,**
Economic impact of cruise activity: the port of Barcelona
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Ayuso, M. (Riskcenter, XREAP)**, Guillen, M.** (Riskcenter, XREAP)**, Nielsen, J. P.**
Improving automobile insurance ratemaking using telematics: incorporating mileage and driver behaviour data
(Desembre 2016)

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Ruíz, A. (GEAP, XREAP), **Matas, A.** (GEAP, XREAP), **Raymond, J. Ll.**
How do road infrastructure investments affect the regional economy? Evidence from Spain
(Desembre 2016)

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Bernardo, V. (GiM-IREA, XREAP); **Fageda, X.** (GiM-IREA, XREAP)
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Di Paolo, A. (AQR-IREA, XREAP); **Tansel, A.**
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Melguizo, C. (AQR-IREA, XREAP); **Royuela, V.** (AQR-IREA, XREAP)
What drives migration moves across urban areas in Spain? Evidence from the great recession
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Boonen, T.J., Guillén, M. (RISKCENTER, XREAP); **Santolino, M.** (RISKCENTER, XREAP)
Forecasting compositional risk allocations
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