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**FUNDACIÓN DE LAS CAJAS DE AHORROS
DOCUMENTO DE TRABAJO
Nº 532/2010**

De conformidad con la base quinta de la convocatoria del Programa de Estímulo a la Investigación, este trabajo ha sido sometido a evaluación externa anónima de especialistas cualificados a fin de contrastar su nivel técnico.

ISSN: 1988-8767

La serie **DOCUMENTOS DE TRABAJO** incluye avances y resultados de investigaciones dentro de los programas de la Fundación de las Cajas de Ahorros.

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Abstract:

This paper generalizes the Dynamic Conditional Correlation (DCC) model of Engle (2002) to incorporate a flexible non-Gaussian distribution based on Gram-Charlier expansions. The resulting semi-nonparametric (SNP)-DCC model admits a separate estimation of, in a first stage, the individual conditional variances under a Gaussian distribution and, in the second stage, the conditional correlations and the rest of the density parameters, thus overcoming the known "dimensionality curse" of the multivariate volatility models. Furthermore the proposed SNP-DCC model solves the negativity problem inherent to truncated SNP densities providing a parametric structure that may accurately approximate a target heavy-tailed distribution. We test the performance of a SNP-DCC model with respect to the (Gaussian)-DCC through an empirical application of density forecasting for portfolio asset returns data. Our results show that the proposed multivariate model provides a better in-sample fit and forecast of the portfolio returns distribution, being thus useful for financial risk forecasting and evaluation.

JEL classification: C16, G1.

Keywords: Density forecasts; Financial markets; GARCH models; Multivariate time series; Semi-nonparametric methods.

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Acknowledgements: Financial support from the Spanish Ministry of Science and Technology under Projects SEJ2006-06104/ECON and SEJ2007-66592-C03-03 is gratefully acknowledged.

1. Introduction

The traditional concern of economic forecasting has been focused on producing and evaluating point forecasts. In the last decades, however, an increasing interest for probability and full density forecasting has emerged, especially for macroeconomic and financial variables that exhibit features such as time-varying dependent moments, skewness or excess kurtosis. The forecasting of those variables in terms of intervals (particularly, one-side interval forecasts or Value-at-Risk) or density forecasts has become a topic of major research guided by both the practitioners demands to improve risk measures accuracy, and the development of powerful models for time-varying conditional variances and densities (see e.g. Gouriéroux, 1997, and Jondeau, Poon & Rockinger, 2007) and methods for their evaluation (Clements, 2005). In particular, Christoffersen (1998) provided appropriate techniques for evaluating interval forecasts and Diebold, Gunther & Tay (1998) for density forecasting, the latter methodology being applied during the last decade for forecasting the distribution of varied variables such as expected future inflation in Diebold, Tay & Wallis (1999), option prices in Aït-Sahalia & Lo (1998) or intraday electricity prices in Panagiotelis & Smith (2008) (see Tay & Wallis (2000) for a comprehensive survey on density forecast applications). On the other hand, extensions of density forecasting to the multivariate case are considered by Diebold, Hahn & Tay (1999), Clements & Smith (2000) and, very recently, a strand of literature has focused on proposing density forecasts evaluation methods based on different criteria, such as: the Kullback-Leibler information criterion (Mitchell and Hall (2005) and Bao, Lee & Saltoglu (2007)), weighted likelihood ratio tests (Amisano & Giacomini, 2007), Bayesian approach (Geweke & Amisano, 2008, 2009) and strictly proper scoring rules (Gneiting & Raftery (2007), Gneiting & Ranjan (2008), Diks, Panchenko & van Dijk (2008)); see Mitchell & Wallis (2009) for a review on density forecast evaluation procedures.

In this paper we apply density forecasting techniques in a multivariate framework for testing the performance of a generalization of Engle's (2002) Dynamic Conditional Correlation (DCC hereafter) model which can incorporate not only volatility clustering and time-varying correlations but also all the non-normal stylized features of high frequency financial variables, i.e. skewness, leptokurtosis, multimodality, etc. This new model that we call Semi-nonparametric (SNP hereafter)-DCC specifies a general and flexible multivariate density based on truncated Edgeworth and Gram-Charlier series.¹

¹ The Cornish and Fisher and the Laguerre expansions are other alternative SNP that might be implemented (see e.g. Cornish and Fisher (1937) and Marumo and Wolff (2007)).

These series were initially defined in the seminal papers of Edgeworth (1907) and Charlier (1905) to approximate a probability distribution in terms of its cumulants. Since then, they have been investigated from many perspectives and disciplines, Sargan (1975) being who brought them into econometrics. More recently, that literature has attracted a renewed interest and so we find contributions, in the theoretical field --see, e.g., Nishiyama & Robinson (2000) and Velasco & Robinson (2001)- and in the applied field, with particular emphasis in finance, to deal with the unsolved problem of fitting the heavy-tailed distribution of high-frequency asset returns ---see, e.g., Corrado & Su (1996), Mauleón & Perote (2000) and Verhoeven & McAleer (2004). The latter articles provide empirical evidence on the good in-sample performance of univariate SNP distributions. Nonetheless, the applications of SNP densities require truncation of the Hermite expansion and the resulting density presents then the well-known definition problem of not being positive for all values of the parameters in the parametric space, as firstly highlighted by Barton & Dennis (1952). This problem can be dealt with in different ways, e.g., through parametric restrictions or density reformulations, see the following papers for further details on those procedures: Gallant & Nychka (1987), Gallant & Tauchen (1989), Jondeau & Rockinger (2001), León, Rubio & Serna (2005) and León, Mencía & Sentana (2009).

Focusing on the multivariate context, different methodologies have been proposed to deal with the modeling of the joint distribution of correlated variables, including: (1) multivariate distributions such as, Multivariate Skewed Normal (Azzalini & Dalla Valle, 1996), Multivariate Student's t (Kotz & Nadarajah, 2004), Multivariate Weibull (Malevergne & Sornette, 2004) or Kotz-type distributions (Olcay, 2005), (2) copula methods (see, e.g., Chen, Fan & Tsyrennikov, 2006) and, (3) multivariate GARCH models (see, e.g., Bauwens, Laurent & Rombouts (2005) for a comprehensive review of those models). In particular, the latter methodology includes the following contributions: the vech model (Bollerslev, Engle & Wooldridge, 1988); the Constant Conditional Correlation (CCC) (Bollerslev, 1990); the Factor-ARCH (Engle, Ng & Rothschild, 1990); the BEKK (Engle & Kroner, 1995); the DCC model (Engle (2002) and Engle & Sheppard (2001)); the Time-Varying Correlation (Tse and Tsui, 2002); the Asymmetric DCC (Cappiello, Engle & Sheppard, 2006); or the Dynamic Equicorrelation (DECO) model (Engle & Kelly, 2007). However, up to the knowledge of the authors, much less has been done regarding the modeling of dynamics of the variance-covariance matrix in a SNP framework. Particularly, Perote (2004) and Del Brio, Ñíguez & Perote (2009) have shown that the multivariate SNP modeling renders multivariate flexible densities that provide accurate fits to financial returns series of

data, but the densities implemented in those papers assume constant correlations, which although eases their implementation for large portfolios, may be too restrictive. Furthermore, the so-called multivariate Edgeworth-Sargan ES (MES hereafter) distribution introduced by Perote (2004) cannot strictly be considered a probability density function (pdf hereafter) since it is not positive for all values of the parameters in the parametric space, and therefore its applications, e.g. for forecasting, are limited.

Regarding the modeling of multivariate distributions conditional moments, the DCC model (Engle, 2002) was introduced to allow for time-varying correlations, and at the same time deal with the known "curse of dimensionality" of the multivariate context, by means of separate quasi-maximum likelihood estimation (QMLE) of conditional variances, and correlation dynamics. Unfortunately, the DCC model two-step estimation procedure has been so far shown to be theoretically valid only under normality (see Engle & Sheppard (2001) for a detailed discussion on the estimation properties of DCC models). Despite this fact, DCC processes have been implemented in empirical works using the Student's t distribution, i.e., using first stage QMLE of conditional variances followed by second stage MLE of correlation processes and Student's t distribution parameters. Bauwens & Laurent (2005) and Jondeau & Rockinger (2005) showed by means of an empirical application that although the decomposition proposed by Engle (2002) is not formally possible for the Student's t distribution and a one-step approach should be adopted (i.e. joint MLE of conditional variances and covariances processes under the Student's t), the estimation results from both one step and two steps approaches did not differ significantly.

In this article, we jointly analyze the aforementioned problems, namely, the negativity problem inherent to truncated SNP distributions, and the multivariate framework "curse of dimensionality" -Specifically when the model accounts for time-varying correlations-, through a new family of multivariate SNP distributions. That family presents the following features: (i) generality: it encompasses as marginals not only the Gaussian but also the different univariate Edgeworth and Gram-Charlier distributions proposed in the literature, (ii) positiveness: it is positive for all its parameter values in the parametric space, (iii) empirical tractability: it theoretically admits the decomposition proposed by Engle (2002), allowing to obtain consistent MLE (under correct specification), which eases the model implementation, and (iv) it yields a reasonable out-of-sample performance for forecasting the density of portfolio returns, as stems from a density forecasting analysis based on the probability integral transform (PIT) paradigm

(Rosenblatt (1952), Diebold, et al. (1999)) and scoring rules for ranking density forecasting methods (Winkler (1967), Amisano & Giacomini (2007)).

Our SNP-DCC model is inspired by the fact that, up to the authors knowledge, none of the aforementioned non-Gaussian multivariate approaches formally admits the estimation in two stages proposed by Engle (2002). The solution provided by our approach in relation to the latter point, together with its other features, (i), (ii) and (iv) above, makes it a useful tool for financial econometrics research and applications.

The remainder of the article is divided into the following four sections. In Section 2, we present the methodology to define multivariate SNP densities, and discuss their properties. Section 3 describes a particular SNP-DCC model of interest for financial applications. Section 4 provides an empirical application to two bivariate portfolios composed by US stock returns, an exchange rate index and a stock returns index, to test the in- and out-of-sample performance of the proposed model. Section 5 summarizes the main conclusions.

2. Multivariate SNP distributions

The modeling of the multivariate distribution of financial variables faces serious obstacles that are still not fully resolved. For instance, the generalization of univariate marginal distributions to a multivariate framework has been successfully achieved by the use of copulas, but those models present drawbacks related with the integration of the joint distribution (e.g., to compute moments), which becomes analytically intractable and requires the implementation of highly computationally demanding numerical algorithms (see Jondeau et al. (2007) p. 196). On the other hand, the use of the existing heavy-tailed and skewed multivariate distributions are interesting alternatives, but in most cases those distributions cannot capture some of the salient features of financial returns due to the lack of a sufficiently flexible parametric structure. Moreover, the multivariate time-varying models can also be implemented for capturing conditional moments, but at the cost of a parameter structure that might give place to very complicated specifications for large-dimensional portfolios. That problem is known in the literature as the "dimensionality curse", and has been tackled with more parsimonious multivariate models and through estimation procedures in two stages, as in the DCC model. Nevertheless, the implementation of the latter process requires the separability of the log-likelihood function, and this property has been shown, so far in

the literature, to be formally possible only under normality. This fact is an important limitation of the DCC model which was proposed to capture financial variables dynamics, but it is not able to capture the known leptokurtosis presented by those variables.

In this paper we present a different approach to the joint distribution of financial asset returns that solves some of the aforementioned shortcomings, and produces reasonably good empirical results. Specifically, we propose a methodology to generalize the univariate SNP-type of distributions to a multivariate framework, which preserves the good performance of the SNP distributions in terms of generality and flexibility, since the marginal distributions of the proposed joint density are univariate SNP densities. The methodology may also incorporate the Gallant & Nychka's (1987) methodology to ensure positivity, but the main stress is put on achieving the separability of the log-likelihood to be able to implement the two-stage estimation procedure established in the DCC model. For this purpose we initially define the multivariate pdf distribution of a random vector $X_t = (x_{1t}, x_{2t}, \dots, x_{nt})' \in \Re^n$ with zero mean and uncorrelated variables as given in equation (1),

$$F^\zeta(X_t, \gamma) = \frac{1}{n} \left[\prod_{i=1}^n g(x_{it}) \right] \left[\sum_{i=1}^n \omega_i^\zeta q_i^\zeta(x_{it}) \right] = \frac{(2\pi)^{-n/2}}{n} \exp \left\{ -\frac{X_t' X_t}{2} \right\} \left[\sum_{i=1}^n \omega_i^\zeta q_i^\zeta(x_{it}) \right], \quad (1)$$

where $g(\cdot)$ denotes a standard Gaussian density, $q_i^\zeta(\cdot)$ (equation (2)) represents the Hermite polynomials expansion for the i -th variable ($i=1,2,\dots,n$), m is the truncation order, which without loss of generality is considered the same for all i , and ω_i^ζ are the constants that make the marginal densities to integrate up to one,

$$q_i^\zeta(x_{it}) = \begin{cases} 1 + \sum_{s=2}^m \gamma_{is} H_s(x_{it}), & \text{if } \zeta = I, \\ 1 + \sum_{s=2}^m \gamma_{is}^2 H_s(x_{it})^2, & \text{if } \zeta = II. \end{cases} \quad (2)$$

This distribution type involves different specifications depending on the assumed Hermite polynomials structure. Specifically, in this article we compare two different multivariate SNP alternatives, denoted with the index $\zeta=I,II$. These distributions, named SNP_I and SNP_{II} , are inspired on the distributions in Perote (2004) and Del Brio et. al.

(2009), respectively, but in contrast with those distributions which consider a combination of a multivariate Gaussian density with a non-diagonal covariance matrix and a term that incorporates the Hermitian expansion for every variable, they are initially defined in terms of uncorrelated variables (i.e. correlation is introduced by means of linear restrictions). This idea of specifying the multivariate SNP density in these terms is the key to achieve the separability of the log-likelihood function as proposed by Engle (2002), and allows to formally implement two-step maximum likelihood estimation, as discussed in the next section.

The SNP_I distribution has a simpler structure than the SNP_{II} but, unlike the latter, it does not guarantee positivity for all the values of density parameters, $\gamma_{is} \forall i=1,2,\dots,n$ and $\forall s=2,\dots,m$. The Hermite polynomials, denoted by $H_s(\cdot)$, $\forall s \in \mathbb{N}$ ($H_0(\cdot)=1$), can be defined in terms of the s^{th} order derivative of the Gaussian pdf as,

$$\frac{d^s g(z)}{dz^s} = (-1)^s g(z) H_s(z). \quad (3)$$

$H_s(\cdot)$ satisfy, among others, the orthogonality properties given in equations (4), (5) and (6) below; see Kendall & Stuart (1977) for further details on Hermite polynomials properties,

$$\int H_s(z) g(z) dz = 0, \quad \forall s > 0 \quad (4)$$

$$\int H_s(z) H_j(z) g(z) dz = \begin{cases} 0, & \forall s \neq j, \\ s!, & \forall s = j, \end{cases} \quad (5)$$

$$\int H_s(z)^2 H_j(z)^2 g(z) dz = s! j!, \quad \forall s \neq j. \quad (6)$$

Under those orthogonality conditions, it can be straightforwardly shown that multivariate SNP_ζ ($\zeta=I, II$) densities satisfy the following properties:

Property 1. The constants, ω_i^ζ , weighting $q_i^\zeta(\cdot)$, are (by direct application of equation (5)) given by,

$$\omega_i^\zeta = \left[\int g(x_{it}) q_i^\zeta(x_{it}) dx_{it} \right]^{-1} = \begin{cases} 1, & \text{if } \zeta = I, \\ \left[1 + \sum_{s=2}^m \gamma_{is}^2 s! \right]^{-1}, & \text{if } \zeta = II. \end{cases} \quad (7)$$

Property 2. The multivariate SNP_I and SNP_{II} functions integrate up to one (see Proof 1 in the Appendix).

Property 3. The marginal densities of the multivariate SNP_I are the standard Gram-Charlier approximations, whilst the marginals of the multivariate SNP_{II} are mixtures of univariate normal and univariate SNP of the type analyzed in Níguez & Perote (2004), as shown in equation (8) (see Proof 2 in the Appendix).

$$f_i^\zeta(x_{it}) = \begin{cases} g(x_{it}) \left[1 + \sum_{s=2}^n \frac{\gamma_{is}}{n} H_s(x_{it}) \right], & \text{if } \zeta = I, \\ g(x_{it}) \left[\frac{n-1}{n} + \frac{1}{n} \omega_i^{II} q_i^{II}(x_{it}) \right], & \text{if } \zeta = II. \end{cases} \quad (8)$$

Property 4. All order moments of SNP_I and SNP_{II} exist and can be obtained in terms of the density parameters and the moments of the standard Normal density (μ_r , $\forall r \in \mathbb{N}$), as displayed in equation (9),

$$E^\zeta[x_{it}^r] = \begin{cases} \mu_r + \sum_{j=0}^r j! c_j \frac{\gamma_{ij}}{n}, & \text{if } \zeta = I, \quad \forall r \in \mathbb{N}, \\ \left[((n-1)\omega_i^{II} + 1)\mu_r + \sum_{s=2}^m s! \gamma_{is}^2 \sum_{j=0}^{r/2} j! d_j \right], & \text{if } \zeta = II, \quad \forall r \in \mathbb{N} \text{ even,} \end{cases} \quad (9)$$

where $\{c_j\}_{j=0}^r$ is a sequence of constants such that $x_{it}^r = \sum_{j=0}^r c_j H_j(x_{it})$, e.g.,

$$x_{it}^r = \begin{cases} H_1(x_{it}), & \text{if } r = 1, \\ H_2(x_{it}) + 1, & \text{if } r = 2, \\ H_3(x_{it}) + 3H_1(x_{it}), & \text{if } r = 3, \\ H_4(x_{it}) + 6H_2(x_{it}) + 3, & \text{if } r = 4, \end{cases} \quad (10)$$

and $\{d_j\}_{j=0}^{r/2}$ is another sequence of constants such that $x_{it}^r = \sum_{j=0}^{r/2} d_j H_j(x_{it})^2$, e.g.,

$$x_{it}^r = \begin{cases} H_1(x_{it})^2, & \text{if } r = 2, \\ H_2(x_{it})^2 + 2H_1(x_{it})^2 - 1, & \text{if } r = 4, \\ H_3(x_{it})^2 + 6H_2(x_{it})^2 + 3H_1(x_{it})^2 - 6, & \text{if } r = 6, \\ H_4(x_{it})^2 + 12H_3(x_{it})^2 + 30H_2(x_{it})^2 + 12H_1(x_{it})^2 - 21, & \text{if } r = 8, \end{cases} \quad (11)$$

Proof 3 in the Appendix demonstrates this result by assuming (without loss of generality) $r < m$.

Property 5. The SNP cumulative distribution functions (cdf hereafter) can be easily computed by means of equations (12) and (13) (see Proof 4 in the Appendix).

$$\Pr[x_1 \leq a_1, \dots, x_n \leq a_n]^\zeta = \frac{1}{n} \sum_{i=1}^n \Psi^\zeta(a_i) \left[\prod_{j=1, j \neq i}^n \Phi(a_j) \right], \quad \forall \zeta = \text{I, II}, \quad (12)$$

where $\Phi(\cdot)$, $\Psi^{\text{I}}(\cdot)$ and $\Psi^{\text{II}}(\cdot)$ stand for the cdfs of the standard Normal, Gram-Charlier expansion and Positive Edgeworth-Sargan (PES) (*Níguez & Perote, 2004*) univariate distributions, respectively.

$$\Psi^\zeta(a_i) = \begin{cases} \Phi(a_i) - g(a_i) \sum_{s=2}^m \gamma_{is} H_{s-1}(a_i), & \text{if } \zeta = \text{I}, \\ \Phi(a_i) - \omega_i g(a_i) \sum_{s=2}^m \gamma_{is}^2 \sum_{l=0}^{s-1} \frac{s! H_{s-1}(a_i) H_{s-l-1}(a_i)}{(s-l)!}, & \text{if } \zeta = \text{II}, \end{cases} \quad (13)$$

Property 6. For a given vector of quantiles $(a_1, a_2, \dots, a_n)'$, the difference in probabilities between the multivariate Gaussian and SNP densities can be obtained from the following closed formula (see Proof 5 in the Appendix).

$$\Pr[x_1 \leq a_1, \dots, x_n \leq a_n]^N - \Pr[x_1 \leq a_1, \dots, x_n \leq a_n]^\zeta = \frac{1}{n} \sum_{i=1}^n B^\zeta(a_i) \left[\prod_{j=1, j \neq i}^n \Phi(a_j) \right], \quad \forall \zeta = \text{I, II}, \quad (14)$$

where $\Pr[x_1 \leq a_1, \dots, x_n \leq a_n]^N$ and $\Pr[x_1 \leq a_1, \dots, x_n \leq a_n]^\zeta$ stand for the multivariate cdfs of the standard Normal and SNP _{ζ} densities ($\zeta=1,2$), respectively, and

$$B^\zeta(a_i) = \begin{cases} g(a_i) \sum_{s=2}^m \gamma_{is} H_{s-1}(a_i), & \text{if } \zeta = \text{I}, \\ \omega_i g(a_i) \sum_{s=2}^m \gamma_{is}^2 \sum_{l=0}^{s-1} \frac{s! H_{s-1}(a_i) H_{s-l-1}(a_i)}{(s-l)!}, & \text{if } \zeta = \text{II}, \end{cases} \quad (15)$$

Properties 1-6 support the multivariate SNP as a very flexible (i.e. it can capture skewness, leptokurtosis, multimodality and most of the high frequency financial features through its general parametric structure) and easy to implement distribution (e.g. moments, probabilities and quantiles can be straightforwardly computed). Nevertheless, in spite of these interesting properties, this distribution would certainly be more useful if correlation among variables were incorporated. To do so, we transform the vector X_t such as the transformed variable, $U_t = \Sigma_t^{1/2} X_t$, has zero mean and variance-covariance matrix Σ_t , with $\Sigma_t = \Sigma_t^{1/2} \Sigma_t^{1/2}$ being the symmetric spectral decomposition, i.e. $\Sigma_t^{1/2} = C_t \Lambda_t^{1/2} C_t'$, where $\Lambda_t = \text{diag}\{\lambda_{1t}, \lambda_{2t}, \dots, \lambda_{2n}\}$ is the diagonal matrix of the eigenvalues of Σ_t , and C_t the corresponding orthogonal matrix of eigenvectors of Σ_t . Note that this decomposition of the matrix Σ_t , which is always possible for symmetric and positive definite matrices, yields the product of two identical symmetric matrices, unlike either the Cholesky decomposition (in terms of triangular matrices) or the "non symmetric" eigenvector decomposition ($\Sigma_t^{1/2} = C_t \Lambda_t^{1/2}$). Furthermore, note that if the variance and covariance matrix of X_t is given by $K_t^2 = \text{diag}\{k_{1t}^2, k_{2t}^2, \dots, k_{nt}^2\}$, then the vector X_t can be standardized by the following transformation: $X_t^* = K_t^{-1} X_t$. If so, the variance and covariance matrix of $U_t = \Sigma_t^{1/2} X_t^*$ can be interpreted in terms of the matrix Σ_t . However, if the focus of the empirical analysis is forecasting, the latter transformation is unnecessary.

Alternatively, we can write $U_t = D_t R_t^{1/2} X_t$, since Σ_t can be decomposed in terms of the diagonal matrix of conditional standard deviations, $D_t = \text{diag}\{\sigma_{1t}, \dots, \sigma_{nt}\}$, and the correlation matrix, R_t , as given in equation (16),

$$\Sigma_t = D_t R_t D_t = D_t R_t^{1/2} R_t^{1/2} D_t. \quad (16)$$

Then, U_t is distributed according to a multivariate SNP distribution, whose pdf is displayed in equation (17).

$$F^\zeta(U_t; \gamma) = (2\pi)^{-\frac{n}{2}} |\Sigma_t|^{-\frac{n}{2}} \exp\left\{-\frac{1}{2} U_t' \Sigma_t^{-1} U_t\right\} \left[\sum_{i=1}^n \omega_i^\zeta q_i^\zeta \left(\Sigma_t^{-\frac{1}{2}} U_t \right) \right] \frac{1}{n}, \quad \forall \zeta = \text{I, II}, \quad (17)$$

Analogously, the SNP density can be re-written in terms of $E_t = D_t \square^{-1} U_t$ as,

$$F^\zeta(E_t; \gamma) = (2\pi)^{-\frac{n}{2}} |R_t|^{-\frac{n}{2}} \exp\left\{-\frac{1}{2} E_t' R_t^{-1} E_t\right\} \left[\sum_{i=1}^n \omega_i^\zeta q_i^\zeta \left(R_t^{-\frac{1}{2}} E_t \right) \right] \frac{1}{n}, \quad \forall \zeta = \text{I, II}, \quad (18)$$

In particular, for the bivariate case ($i=1,2$) the inverse transformation $X_t = R_t \square^{-1} E_t$, can be easily written as a function of the standardized variables, $\varepsilon_{it} = u_{it}/\sigma_{it}$, and time-varying correlation, ρ_t , as displayed in equation (19),

$$x_{it} = \begin{cases} \frac{1}{2} \left(\frac{1}{\sqrt{1+\rho_t}} + \frac{1}{\sqrt{1-\rho_t}} \right) \varepsilon_{1t} + \frac{1}{2} \left(\frac{1}{\sqrt{1+\rho_t}} - \frac{1}{\sqrt{1-\rho_t}} \right) \varepsilon_{2t}, & \text{if } i = 1, \\ \frac{1}{2} \left(\frac{1}{\sqrt{1+\rho_t}} - \frac{1}{\sqrt{1-\rho_t}} \right) \varepsilon_{1t} + \frac{1}{2} \left(\frac{1}{\sqrt{1+\rho_t}} + \frac{1}{\sqrt{1-\rho_t}} \right) \varepsilon_{2t}, & \text{if } i = 2, \end{cases} \quad (19)$$

3. The multivariate SNP-DCC model

Let Y_t be a $n \times 1$ vector of asset returns conditionally (on the information set Ω_{t-1}) distributed according to a multivariate SNP with conditional first and second moments defined by $\mu_t(\phi)$ and $\Sigma_t(\theta) = D_t(\alpha) R_t(\rho) D_t(\alpha)$, respectively, where ϕ , θ , α and ρ are parameter vectors.² Following Engle's (2002) DCC model specification, the matrices $D_t(\alpha)$ and $R_t(\rho)$ are modelled as displayed in equations (23) to (26); the equation (25) being the MARCH family of Ding & Engle (2001) where S is the unconditional correlation matrix, ξ is a vector of ones, A , B and $\xi \xi' - A - B$ are positive definite matrices, and \square is the Hadamard product of two identically sized matrices (computed by element by element multiplication).

² Usually either an AR(1) or MA(1) process is specified for $\mu_{\{t\}}(\phi)$ selected according to information criteria to account for the small structure in the level of series of financial asset returns, see for instance, Sentana & Wadhwani (1992).

$$Y_t = \mu_t(\phi) + U_t, \quad (20)$$

$$U_t | \Omega_{t-1} \approx SNP(0, \Sigma_t(\theta)), \quad (21)$$

$$\Sigma_t(\theta) = D_t(\alpha)R_t(\rho)D_t(\alpha) = D_t(\alpha)R_t(\rho)^{1/2}R_t(\rho)^{1/2}D_t(\alpha), \quad (22)$$

$$D_t^2 = diag\{\alpha_{i0}\} + diag\{\alpha_{il}\} \circ U_{t-1}U_{t-1}' + diag\{\alpha_{i2}\} \circ D_{t-1}^2, \quad (23)$$

$$E_t = D_t^{-1}U_t, \quad (24)$$

$$Q_t = S \circ (\xi\xi' - A - B) + A \circ E_{t-1}E_{t-1}' + B \circ Q_{t-1}, \quad (25)$$

$$R_t = diag\{Q_t\}^{-1/2}Q_tdiag\{Q_t\}^{-1/2}, \quad (26)$$

This model, named SNP-DCC, encompasses the (Gaussian)-DCC, and presents a log-likelihood which (as the one of the (Gaussian)-DCC) can be split into two different components: (i) the mean-volatility part, $L_{MV}(Y_t, \eta)$, where $\eta = (\phi', \alpha')'$, and (ii) the "standardized" SNP component, $L_{SNP}^\zeta(E_t, \eta, \square)$, where $\varphi = (\rho', \gamma')'$ includes the correlation and the shape parameters (see *Proof 6* in the Appendix). Specifically, these two terms can be written (after deleting the unnecessary constants) as given in equations (27) and (28).

$$L_{MV}(Y_t, \eta) = -\frac{1}{2} \sum_{i=1}^n \left[T \log(2\pi) + \sum_{t=1}^T \left(\ln(\sigma_{it}^2) + \frac{(y_{it} - \mu_{it})^2}{\sigma_{it}^2} \right) \right] = -\frac{1}{2} \sum_{i=1}^n [T \log(2\pi) + L_{MV_i}(\eta_i)], \quad (27)$$

$$L_{SNP}^\zeta(E_t, \eta, \varphi) = -\frac{1}{2} \sum_{t=1}^T \left\{ \ln|R_t| + E_t'E_t - 2 \ln \left[\sum_{i=1}^n \omega_i^\zeta q_i^\zeta \left(R_t^{-\frac{1}{2}} E_t \right) \right] \right\}, \quad \forall \zeta = I, II. \quad (28)$$

Thus, the SNP-DCC model can be estimated by ML in two steps as follows: first, η is obtained by maximizing independently $L_{MV_i}(y_{it}, \eta_i)$ for each $i=1, 2, \dots, n$ (note that this step assumes normality), and second, the conditional correlation parameters and the multivariate SNP shaping parameters, φ , are jointly estimated in the log-likelihood function concentrated with respect to $\hat{\eta} = \text{argmax}\{L_{MV}(Y_t, \eta)\}$, i.e. $L_{SNP}^\zeta(E_t, \hat{\eta}, \square)$. As pointed out by Engle (2002), this two-step estimation problem can be interpreted in terms of the joint Generalized Method of Moments estimation framework discussed in Section 6.1 of Newey & McFadden (1994). Particularly, if the conditions

$\frac{1}{T} \sum_{t=1}^T \nabla_\eta L_{MV}(Y_t, \eta) = 0$ and $\frac{1}{T} \sum_{t=1}^T \nabla_\varphi L_{SNP}^\zeta(E_t, \hat{\eta}, \varphi) = 0$ are satisfied with probability

approaching one, $\hat{\eta} \xrightarrow{P} \eta_0$, $\hat{\phi} \xrightarrow{P} \phi_0$, and under the regularity conditions (i)-(v) of Theorem 3.4 in Newey & McFadden (1994), then $\hat{\eta}$ and $\hat{\phi}$ are asymptotically Normal, and

$$\sqrt{T}(\hat{\phi} - \phi_0) \xrightarrow{d} N(0, V), \quad (29)$$

$$V = [E(\nabla_{\phi\phi} L_{SNP}^{\xi})]^{-1} E(HH')^{-1} [E(\nabla_{\phi\phi} L_{SNP}^{\xi})]^{-1}, \quad (30)$$

$$H = \nabla_{\phi} L_{SNP}^{\xi} - E(\nabla_{\phi\eta} L_{SNP}^{\xi}) [E(\nabla_{\eta\eta} L_{MV})]^{-1} \nabla_{\eta} L_{MV}, \quad (31)$$

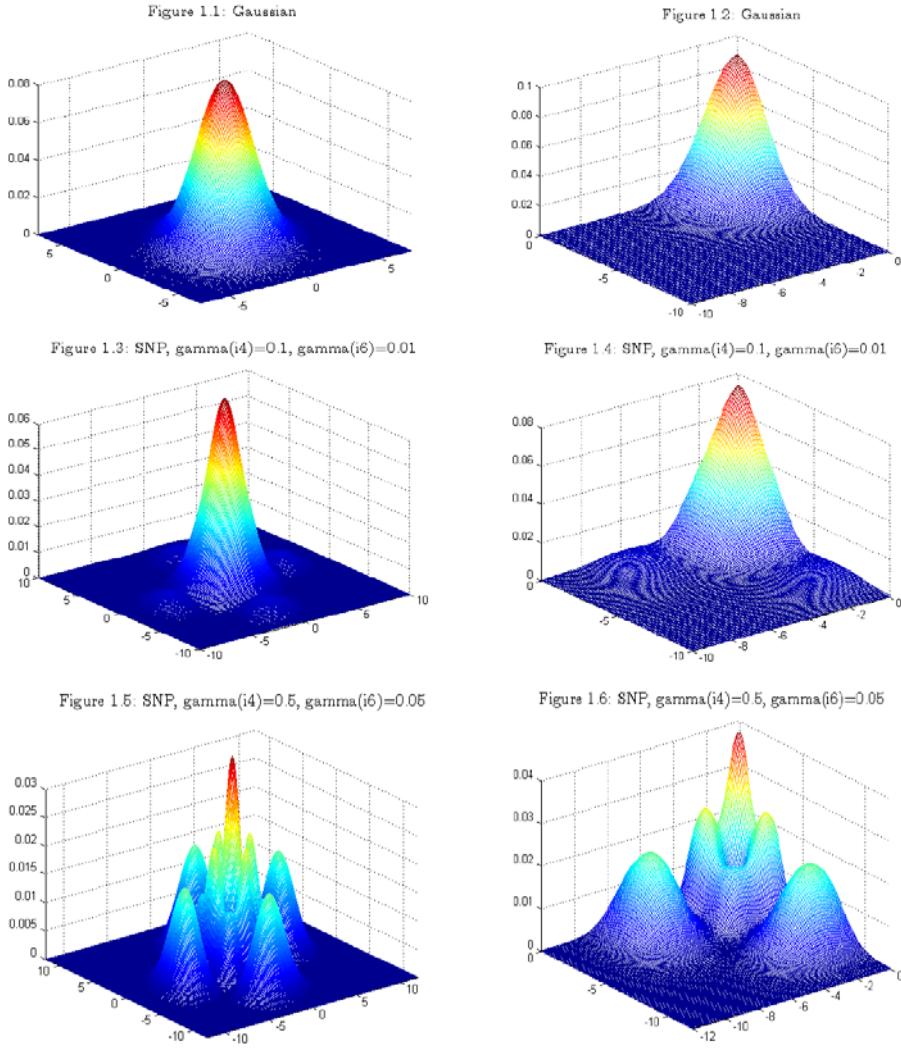
with $\nabla_{\eta}(\square)$ and $\nabla_{\eta\eta}(\square)$ denoting the gradient vector of first derivatives with respect to the variables in η and the Hessian matrix of the second derivatives with respect to η and \square , respectively, (see Theorem 6.1 in Newey & McFadden (1994) and its application to the DCC model in Engle (2002) and Engle & Sheppard (2001) for further details). It must be noted that in the context of the SNP-DCC, the first step is QMLE, but the second step is MLE since the likelihood corresponds to the SNP density. Nevertheless we argue that the second-step MLE is also consistent and asymptotically normal as $n \rightarrow \infty$, since the multivariate (infinite) SNP expansion is almost surely the "true" density (Sargan (1975) and Gallant & Nychka (1987)). Moreover, the asymptotic variances of the GARCH parameters are the robust variance and covariance matrix in Bollerslev & Wooldridge (1992), but the asymptotic variance and covariance matrix of the estimates of the second step parameters involves a more complex structure (see Engle & Sheppard, 2001). Furthermore, the estimates of the SNP-DCC parameters are not fully efficient since they are estimated by Limited Information MLE, although the estimates of the SNP-DCC model are more efficient than those of the DCC models provided that $\gamma \neq 0$. However, following Pagan (1986), an asymptotically efficient estimator can be obtained by performing an additional iteration of the Newton-Raphson algorithm for the joint model starting from the two steps consistent (Q)MLE.

On the other hand, it must be noted though that for the SNP-DCC, consistency of the second-step cannot be guaranteed in practice since the Gram-Charlier expansions need to be truncated. Nonetheless, we argue that the density misspecification error can be minimized by selecting the truncation order of the SNP density according to, for instance, Wald specification tests or Information Criteria. In any case, the two-step SNP-DCC is consistent provided that the SNP distribution is correctly specified, unlike other non-Gaussian distributions widely used (e.g. in financial applications) to model

departures from normality, for which the consistency of the DCC two-step estimation cannot be formally proven, even under correct density specification, since the log-likelihood is not separable (see Newey & Steigerwald (1997), Bauwens & Laurent (2005), Jondeau & Rockinger (2005) and Jondeau et al. (2007)). A further discussion of the estimation procedure for a truncated SNP-DCC model is provided in Section 4.

Finally, Figure 1 plots the bivariate Gaussian density and the allowable shapes of a bivariate symmetric SNP_{II} density for different values of its parameters (as orientating illustrations we display the cases of $(d_{i4}, d_{i6})=(0,0)$, $(d_{i4}, d_{i6})=(0.1,0.01)$ and $(d_{i4}, d_{i6})=(0.5,0.05)$). The figures in the left column highlight the tail shape of the densities. It is striking to see the flexibility presented by the SNP density, which is an advantage to fit any empirical distribution, especially in the tails (see also Del Brio, et al. (2009) for further examples on the flexibility of other SNP related multivariate densities). We investigate further the effect of this feature on the performance of a SNP-DCC model in the forecasting exercise of next section.

FIGURE 1
Plots of Bivariate Gaussian and SNP Distributions



Notes: Plots of bivariate Gaussian, and SNP_{II} densities with different allowable shapes depending on their γ_{is} ($gamma(is)$) parameters ($i = 1, 2$).

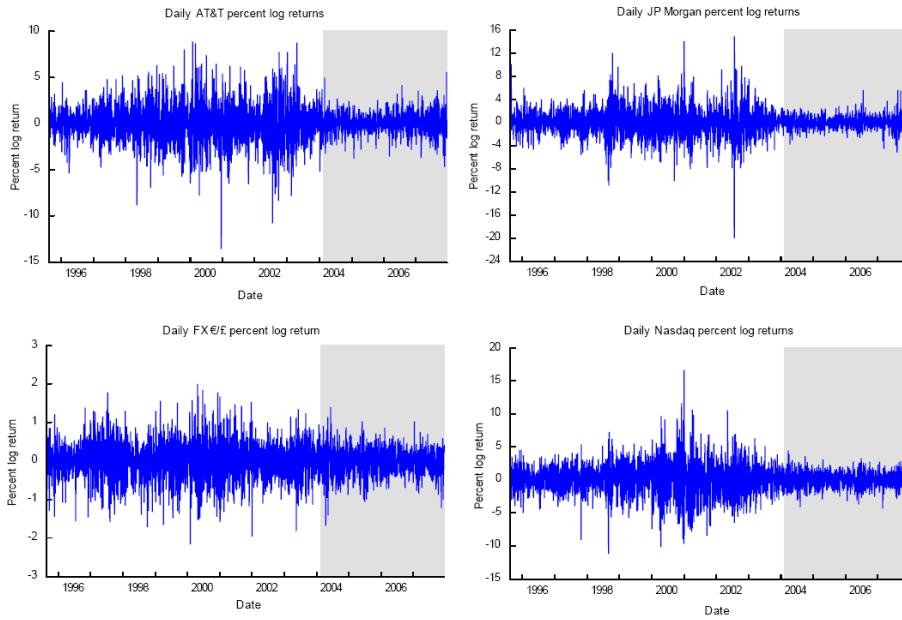
4. An empirical application to portfolio returns

4.1. The data

In this section we investigate the empirical performance of the models discussed above through an in- and out-of-sample comparative analysis. To do so, we consider three types of distributions with either constant or dynamic conditional correlation, which formally admit two-step estimation procedures. The models that we get are then the following: Gaussian (CCC and DCC), non-positive-SNP (SNP_I -CCC and SNP_{II} -DCC) and positive-SNP (SNP_{II} -CCC and SNP_{II} -DCC). The data used are (daily) percent log returns computed from a observed sample of (daily) asset prices, $y_{it} = 100\log(P_{it}/P_{it-1})$, of two (representative) stocks from the Dow Jones index (AT&T and JP Morgan), the Euro(€)/British Pound(£) exchange rate (FX €/£), and the Nasdaq index. We consider two bivariate portfolios: Portfolio AT&T-JP Morgan, and Portfolio

FX €/£-Nasdaq, with returns denoted as $Y_t=(y_{1t}, y_{2t})$. All series are sampled over the period August 17, 1995 to December 17, 2007 for a total of 3,218 observations. The data were obtained from Datastream. Figure 2 displays the plots of the four return series for the total sample. The shaded area in those plots corresponds to the out-of-sample period (February 17, 2004 to December 17, 2007 for a total of $N=1,000$ observations) used for the forecasting analysis in the next section.

FIGURE 2
Plots of daily log returns



Notes: Plots of daily AT&T, JP Morgan, FX €/£ and Nasdaq log returns. In-sample period: 16/08/1995 - 16/02/2004. Out-sample period (graphs shaded area): 17/02/2004 - 17/12/2007. Total observations: 3218.

Table 1 reports some descriptive statistics for the total samples. As expected the stock returns series are much more volatile (as measured by the sample standard deviation) than the FX €/£ series. The unconditional distribution of any of these series shows clearly non-Gaussian features, such as (mild) skewness, and a remarked excess of kurtosis over the Normal distribution. The Jarque-Bera test for normality is easily rejected for the four time series. The analysis of dependence through the Ljung-Box portmanteau test statistic (Q_{LB}) shows some form of weak dependence in the level and a strong correlation in the second-order moment of all series. The asset returns of Portfolio AT&T-JP Morgan present a higher correlation than those of Portfolio FX €/£-Nasdaq. The Likelihood ratio (LR) test for model specification SNP_{II} -DCC versus DCC rejects the null hypothesis $H_0: \gamma_{12}=\gamma_{14}=\gamma_{16}=0$ ($\forall i=1,2$) in favour of the SNP_{II} -DCC model for both portfolios.

TABLE 1
Daily percent log returns descriptive statistics

	Portfolio AT&T - JP Morgan		Portfolio FX €/£ - Nasdaq	
	AT&T	JP Morgan	FX €/£	Nasdaq
Sample	17/08/1995 - 17/12/2007			
Observations	3218			
Mean	0.0160	0.0281	0.0045	0.0294
Median	0.0000	0.0000	0.0211	0.0509
Maximum	8.8337	14.872	1.9909	16.582
Minimum	-13.537	-19.976	-2.1558	-11.139
St. Dev.	1.8610	2.1585	0.4576	2.1451
Skewness	-0.0921	0.0915	-0.1491	0.1385
Kurtosis	6.3862	9.0895	4.2911	6.6315
Jarque-Bera	1542.0*	4976.6*	235.47*	1778.5*
$Q_{LB}(1) - r_t$	1.6591	0.0147	8.3632*	6.1489*
$Q_{LB}(1) - r_t^2$	68.587*	209.91*	64.927*	160.6*
$Q_{LB}(20) - r_t^2$	729.73*	1127.3*	578.09*	2028.7*
$\hat{\rho}$	0.3223		0.0557	
LR _{SNP-DCC_{II} vs DCC}	3186.2*		3130.7*	

Notes: The Jarque-Bera normality test is asymptotically distributed as a $\chi^2(2)$ under the null of normality. $\hat{\rho}$ denotes the sample correlation coefficient for the returns of each bivariate portfolio. $Q_{LB}(\varsigma)$ denotes the Ljung-Box Q -statistic which is asymptotically distributed as $\chi^2(\varsigma)$, ς being the autocorrelation order. The Likelihood Ratio (LR) test is asymptotically distributed as $\chi^2(\iota)$ being ι the number of restrictions under the null —the LR test results reported correspond to the SNP-DCC_{II} model for the null $H_0 : \gamma_{i2} = \gamma_{i4} = \gamma_{i6} = 0$ ($i = 1, 2$). The critical values of $\chi^2(1)$, $\chi^2(2)$, $\chi^2(6)$ and $\chi^2(20)$ are 3.84, 5.99, 12.59 and 31.41, at 5% significance level, respectively. The asterisk (*) denotes rejection of the null hypothesis of the corresponding test at least at 5% significance level.

4.2. Estimation and in-sample analysis

The estimation is carried out in two stages by (Q)MLE techniques, using a moving in-sample window of constant size T=2,217 observations. In the first stage, an AR(1) process for the conditional mean (selected according to the Akaike Information Criteria (AIC hereafter)) and a GARCH(1,1) process for the conditional variance are jointly estimated (under normality) independently for each asset. Then, in the second stage, the standardized residuals from the previous step are used to estimate the conditional correlation equation and the rest of the density parameters. The first stage yields QMLE which is consistent and asymptotically normal, although not efficient. Our second stage MLE is not a priori consistent, since we use a truncated SNP density, but it may be more efficient than the second step QMLE of the Gaussian-DCC. Thus,

we argue that the better our truncated SNP density can approximate the "true" distribution, the more efficient is our second stage MLE. Finally, following Pagan (1996) a further Newton-Raphson iteration without line search for the joint model is performed from the first and second stage (Q)MLE: the estimates do not change but the information matrix is now block diagonal, thus obtaining estimators asymptotically equivalent to joint QMLE. Robust standard errors were computed following Bollerslev & Wooldridge (1992). Taking into account the limitations in the estimation of the truncated SNP-DCC model, the model performance is mainly evaluated in an out-of-sample forecasting framework, in which possible losses in efficiency and consistency are not crucial (see Ruiz & Pascual 2002).

The likelihood functions in each step are maximized using the Berndt, Hall, Hall & Hausman (1974) algorithm. We observe that the estimation of the SNP models is not computationally very demanding providing that starting values are chosen properly. As it is known that SNP densities may present multiple local modes, the optimization is monitored using different starting values to ensure that the (Q)MLE we obtain are global optima.

Table 2 presents the estimation results of the bivariate models. α_{it} , $i=1,2$ and $\tau = 0,1,2$, denote the parameters of the GARCH(1,1) models used for the conditional variances, ϕ_0 and ϕ_1 denote the intercept and the slope, respectively, of the AR(1) process for the conditional mean, and ρ denotes the correlation parameter in CCC models. The rest of parameters displayed in the table follow the same notation used in previous sections. Robust standard errors are in parenthesis next to the parameter estimates. We observed that the presented (Newton-Raphson corrected) standard errors did not change much in relation to those from the separated two-step estimation.

For the specification of the SNP models we started considering densities truncated at the eighth moment, then persistently (across windows) non-significant parameters were removed. In Table 1 we present the final SNP specifications. For both portfolios, the estimated SNP densities are unconditionally symmetric since the odd parameters, γ_{is} $s=3,5,7$, are not significant at any reasonable significance level. For Portfolio AT&T-JP Morgan, the even parameters of the SNP_I , γ_{12} and γ_{16} , were not statistically significant, although they were not removed from the model for the sake of clear comparisons across SNP models. The SNP_{II} models have three significant even parameters for the first asset (γ_{1s} , $s=2,4,6$) and two significant even parameters for the second asset (γ_{2s} , $s=4,6$). For Portfolio FX €/£-Nasdaq, the existing leptokurtosis in the

distribution of the first data window is gathered mainly by y_{14} , y_{24} and y_{26} in SNP_I models, and by y_{14} and y_{26} in SNP_{II} models. It is interesting to note that, the obtained estimates $\hat{\gamma}_{i4}$ and $\hat{\gamma}_{i6}$ for both portfolios, are smaller than those considered in Figures 1.3 and 1.4, and therefore the evidence of multimodality in our data is mild. Nonetheless, the heavy tails of the returns distribution do not necessarily decrease uniformly, as they would do under the commonly assumed parametric distributions in most financial econometrics applications.

According to the AIC, defined as $AIC=2(\square-\ln L)/n$ (\square being the number of the parameters of the model), we observe that SNP models provide a notably better goodness-of-fit than Gaussian models, and that dynamic conditional correlation helps the models to fit the data. Furthermore, among SNP models, SNP_I are preferred to their SNP_{II} counterparts. Regarding the conditional mean and variance estimates, we observe the usual small structure in the conditional mean and high persistence in the conditional variance, that is also observed in the conditional correlation (in line with the results in Engle & Sheppard (2001) and Engle (2002) for US stock returns). Also the obtained estimates of the unconditional correlation coefficient, ρ , given by CCC and SNP-CCC models are close to the sample correlations reported in Table 2.

TABLE 2
Estimation results

$$\begin{aligned} \text{Mean equation: } & y_{it} = \phi_{i0} + \phi_{i1}y_{i,t-1} + u_{it}, \quad u_{it} = \varepsilon_{it}\sigma_{it}, \quad i = 1, 2, \\ \text{Variance equation: } & \sigma_{it}^2 = \alpha_{i0} + \alpha_{i1}u_{i,t-1}^2 + \alpha_{i2}\sigma_{i,t-1}^2, \\ \text{Correlation equation: } & \rho_t = (1 - \delta_1 - \delta_2)\bar{\rho} + \delta_1\varepsilon_{1t-1}\varepsilon_{2t-1} + \delta_2\rho_{t-1}. \end{aligned}$$

	CCC	DCC	$\text{SNP}_I\text{-CCC}$	$\text{SNP}_I\text{-DCC}$	$\text{SNP}_{II}\text{-CCC}$	$\text{SNP}_{II}\text{-DCC}$
Panel 1: Portfolio AT&T - JP Morgan						
Stage 1						
ϕ_{10}			.0343 (1.21)			
ϕ_{11}			-.0494 (-2.23)			
α_{10}			.0362 (1.65)			
α_{11}			.0577 (3.08)			
α_{12}			.9360 (45.1)			
ϕ_{20}			.0990 (2.08)			
ϕ_{21}			.0343 (1.61)			
α_{20}			.0318 (1.66)			
α_{21}			.0551 (4.53)			
α_{22}			.9407 (69.5)			
Stage 2						
γ_{12}		.0037 (0.12)	.0010 (0.07)	.0913 (1.63)	.0888 (1.59)	
γ_{14}		.0912 (6.43)	.0914 (6.55)	.0281 (3.57)	.0267 (3.15)	
γ_{16}		.0012 (0.54)	.0013 (0.55)	.0021 (2.03)	.0020 (1.56)	
γ_{22}		.0001 (0.11)	.0001 (0.24)	.0001 (0.03)	.0001 (0.73)	
γ_{24}		.0902 (6.71)	.0904 (6.71)	-.0293 (-4.95)	-.0283 (-4.79)	
γ_{26}		.0062 (2.43)	.0060 (2.31)	.0031 (3.12)	.0031 (3.13)	
ρ	.2859 (13.5)		.2904 (14.1)		.3240 (14.4)	
δ_1		.0077 (4.19)		.0077 (4.27)		.0085 (4.65)
δ_2		.9877 (329)		.9880 (348)		.9880 (365)
AIC	1.9137	1.9043	.4539	.4446	.4798	.4727

	CCC	DCC	SNP _I -CCC	SNP _I -DCC	SNP _{II} -CCC	SNP _{II} -DCC
Panel 2: Portfolio FX €/£ - Nasdaq						
Stage 1						
ϕ_{10}			.0072 (0.73)			
ϕ_{11}			-.0689 (-3.02)			
α_{10}			.0035 (1.09)			
α_{11}			.0382 (2.21)			
α_{12}			.9474 (33.1)			
ϕ_{20}			.1063 (2.34)			
ϕ_{21}			-.0387 (-1.83)			
α_{20}			.0571 (1.67)			
α_{21}			.0659 (3.37)			
α_{22}			.9246 (39.7)			
Stage 2						
γ_{12}		-.0114 (-0.33)	-.0115 (-0.33)	-.0001 (-0.03)	.0001 (0.13)	
γ_{14}		.0647 (4.43)	.0649 (4.44)	.0270 (4.58)	.0269 (4.55)	
γ_{16}		.0037 (1.37)	.0037 (1.37)	.0011 (0.64)	.0012 (0.71)	
γ_{22}		-.0229 (-0.73)	-.0234 (-0.77)	-.0001 (-0.05)	.0001 (0.10)	
γ_{24}		.0235 (2.10)	.0227 (2.02)	.0059 (0.31)	-.0003 (-0.11)	
γ_{26}		.0039 (2.00)	.0039 (1.99)	.0029 (3.43)	.0031 (4.07)	
ρ	.0642 (2.78)	.0591 (2.66)		.0612 (2.54)		
δ_1		.0090 (1.93)		.0075 (1.12)		.0091 (1.63)
δ_2		.9748 (68.3)		.9802 (38.4)		.9768 (53.1)
AIC	1.9924	1.9901	.5837	.5810	.5858	.5833

Notes: The coefficients presented in this table are (Q)ML estimates of the CCC, DCC, SNP-CCC and SNP-DCC models, for the bivariate portfolios returns. ϕ_{is} ($s = 0, 1$) stand for the AR(1) parameters of the conditional means and α_{is} ($s = 0, 1, 2$) for the GARCH(1,1) parameters of the conditional variances. ρ denotes the unconditional correlation parameter in CCC models and δ_s ($s = 1, 2$) the conditional correlation parameters in DCC models. γ_{is} ($s = 2, 4, 6$) denote the order s polynomial weighting parameter in the SNP models. t statistics calculated from robust standard errors are in parenthesis. Akaike Information Criterion (AIC) is displayed in the last row.

4.3. Density forecasting

In this section we test the performance for full density forecasting of the SNP_{II}-DCC model in comparison with the DCC model. Forecasts are produced by using a rolling window of size N that discards old observations. The recursive optimization is monitored by using the same starting value for all windows, instead of using the usual optimum from the previous data window. This mechanism is used to avoid getting trapped in successive local optima. On the other hand, it is worth noting that although the SNP_I-DCC is in principle a candidate to provide a good forecasting performance, it is not a suitable model in a forecasting experiment that involves data rolling windows over a long time period, since a combination of parameter estimates during the optimization process for a given window may lead to a negative value of the density. In this paper, we have addressed that problem by defining the SNP_{II} distribution for which positivity is guaranteed through reformulation of the density function. Furthermore, we consider models with DCC structures given their better in-sample fit in relation to their CCC counterparts (as shown in Table 1). Thus, SNP_{II}-DCC and (Gaussian)-DCC

models are compared with respect to their density forecasting performance using the following criteria:

1. The PIT paradigm (Pearson (1933), Rosenblatt (1952), Diebold et al. (1998, 1999)) establishes that if $\{y_{it}^*\}_{t=T+1}^{T+N}$ is a series of realizations generated from a series of conditional densities, $\{f(y_{it}|\Omega_{t-1})\}_{t=T+1}^{T+N} = \{f_{i,t-1}(y_{it})\}_{t=T+1}^{T+N}$, and $\{\tilde{f}_{i,t-1}(y_{it})\}_{t=T+1}^{T+N}$ is a series of one-step-ahead density forecasts, then, the series of PIT of $\{y_{it}^*\}_{t=T+1}^{T+N}$ with respect to $\{\tilde{f}_{i,t-1}(y_{it})\}_{t=T+1}^{T+N}$ is i.i.d. $U(0,1)$, i.e.

$$\{p_{it}\}_{t=T+1}^{T+N} = \left\{ \int_{-\infty}^{y_{it}^*} \tilde{f}_{i,t-1}(y_{it}) dy_{it} \right\}_{t=T+1}^{T+N} \stackrel{i.i.d.}{\approx} U(0,1), \quad (32)$$

provided that the forecast densities match the actual densities at each t. Diebold et al. (1999) showed that the application of the PIT paradigm in a multivariate framework can be performed based on the PIT of the realized series with respect to the marginal and conditional one-step-ahead density forecasts as shown in equations (33) and (34), respectively, for the bivariate case ($i,j=1,2$) of the SNP_{II} ,

$$p_{it} = \int_{-\infty}^{y_{it}^*} \tilde{f}_{i,t-1}^{II}(y_{it}) dy_{it}, \quad \forall t = T+1, \dots, T+N, \quad (33)$$

$$p_{i|j,t} = \frac{\int_{-\infty}^{y_{it}^*} \int_{-\infty}^{y_{jt}^*} \tilde{F}_{t-1}^{II}(y_{it}, y_{jt}) dy_{it} dy_{jt}}{\int_{-\infty}^{y_{jt}^*} \tilde{f}_{j,t-1}^{II}(y_{jt}) dy_{jt}}, \quad \forall t = T+1, \dots, T+N. \quad (34)$$

where $\tilde{f}_{i,t-1}^{II}(\cdot)$ and $\tilde{F}_{t-1}^{II}(\cdot, \cdot)$ stand for the one-step ahead marginal and joint density forecasts of the standardized SNP_{II} -DCC distribution, which are given by equations (35) and (36), respectively.³

$$\tilde{f}_{i,t-1}^{II}(y_{it}) = \frac{1}{2\tilde{\sigma}_{it}} g\left(\frac{y_{it} - \tilde{\mu}_{it}}{\tilde{\sigma}_{it}} \right) \left[1 + \hat{\omega}_i^{II} \hat{q}_i^{II} \left(\frac{y_{it} - \tilde{\mu}_{it}}{\tilde{\sigma}_{it}} \right) \right], \quad (35)$$

³ It is worth mentioning that the PIT paradigm does not directly apply to the joint multivariate density (what certainly constitutes an interesting subject of research in the area of density forecasting evaluation), but it does to marginals and conditionals of the multivariate density, as shown by Diebold et al. (1999).

$$F_{t-1}^{II}(Y_t) = \frac{|\tilde{\Sigma}_t|^{-\frac{1}{2}}}{2\sqrt{2\pi}} \exp\left\{-\frac{1}{2}(Y_t - \tilde{\mu}_t)' \tilde{\Sigma}_t^{-1} (Y_t - \tilde{\mu}_t)\right\} \left[\sum_{i=1}^2 \hat{\omega}_i^{II} q_i^{II}\left(\frac{y_{it} - \tilde{\mu}_{it}}{\tilde{\sigma}_{it}}\right) \right]. \quad (36)$$

Note that the forecasted densities are computed for the one-step-ahead forecasts of the conditional means, variances and correlation, denoted by $\tilde{\mu}_{is}$, $\tilde{\sigma}_{it}$ $\forall i=1,2$ and $\tilde{\rho}_t$, respectively, and the estimates of the rest of the density parameters ($\hat{\gamma}_{is}$, $\forall i=1,2$ and $\forall s=2,4,6$) conditioned on the available information set Ω_{t-1} , $t=T+1, \dots, T+N$.⁴ Note also that for the bivariate case $Y_t = (y_{1t}, y_{2t})'$, $\tilde{\mu}_t = (\tilde{\mu}_{1t}, \tilde{\mu}_{2t})$, $\tilde{\Sigma}_t$ is the 2×2 one-step-ahead forecasted conditional variance and covariance matrix, $\hat{\omega}_i^{II} = \left(1 + \sum_{s=2}^m \hat{\gamma}_{is}^2 s!\right)^{-1}$ and $\hat{q}_i^{II}(\cdot) = 1 + \sum_{s=2}^m \hat{\gamma}_{is}^2 H_s(\cdot)$. We consider bivariate portfolios for the sake of simplicity since considering a larger portfolio would make the density forecasting analysis much more complicated, likely to obtain similar conclusions. Since the PIT series $\{p_{it}\}_{t=T+1}^{T+N}$ and $\{p_{i|j,t}\}_{t=T+1}^{T+N}$ are also interpreted as the p-values corresponding to the series of quantiles $\{y_{it}^*\}_{t=T+1}^{T+N}$ of the forecasted marginal and conditional densities, respectively, we use p-value discrepancy plots (i.e. plotting $\hat{P}_{p_{it}}(y_l) - y_l$ against y_l) (Davidson and MacKinnon (1998) and Fiorentini, Sentana and Calzolari (2003)), to test for correct model specification, where $\hat{P}_{p_{it}}(y_l)$ is the cdf of p_{it} calculated as $\hat{P}_{p_{it}}(y_l) - y_l = \frac{1}{N} \sum_{t=T+1}^{T+N} 1(p_{it} \leq y_l)$, with $1\{\cdot\}$ being an indicator function and y_l an arbitrary grid of l points.⁵ Therefore, under correct model specification, the variable $\hat{P}_{p_{it}}(y_l) - y_l$ must converge to zero. The $p_{ij,t}$ sequences can be analyzed analogously.

2. Scoring rules (Winkler (1967), Amisano & Giacomini (2007), Gneiting & Raftery (2007)). A scoring rule is a loss function $\psi(\hat{f}, Y^*)$ whose arguments are the density

⁴ The parameter set might also be forecasted if time-varying conditional models for modeling higher order moments are implemented in the SNP framework (see León et al. (2005) and Níquez, Perote and Rubia (2009) for applications to the univariate case). This is certainly possible and constitutes a clear advantage of the SNP approach, although we leave this topic as an open question for further research.

⁵ Particularly, we use a $\square=215$ points grid that is finer in its extremes to highlight the models forecasting performance of the distribution tails, $y_l=0.001, 0.002, \dots, 0.01, 0.015, \dots, 0.99, 0.991, 0.992, \dots, 0.999$.

forecast \tilde{f} and the realisation Y^* of the future observation of the variable. In this paper we use the logarithmic scoring rule $\psi(\hat{f}, Y^*) = -\log \tilde{f}(Y^*)$ to compare marginal and conditional density forecasts. This is a (strictly proper) scoring rule that rewards a density forecast that assigns high probability to the event that actually occurred.⁶ Density forecast models can be ranked by comparing their average scores, $\bar{\psi}(\tilde{f}, Y^*) = -N^{-1} \sum_{t=T+1}^{T+N} \log \tilde{f}_{t-1}(y_t^*)$. We take the logarithmic scoring rule as a negatively oriented penalty so, we prefer model f if $\bar{\psi}(\tilde{f}, Y^*) < \bar{\psi}(\tilde{g}, Y^*)$, and prefer model g otherwise. The null hypothesis $H_0: E[\psi(\tilde{f}, Y^*) - \psi(\tilde{g}, Y^*)] = 0$ is tested using Amisano & Giacomini (2007) test.

The results in Figure 3 and Table 3 show that the bivariate SNP_{II} distribution provides overall a better performance for forecasting the full density than the Normal. Specifically, the plots in Figure 3 show a remarkable better performable of the SNP-DCC over the DCC at the 5% lower tail, which is highlighted (right column plots in Figure 3) because of its direct relation with widely used financial risk measures such as Value-at-Risk and short-fall probabilities. On the other hand, if we look at higher percentiles, e.g. in the range [0.05, 0.4], the results are mixed but tend to favour the DCC model. Furthermore, Table 3 provides clear evidence in favour of the SNP-DCC for full density forecasting, as the average logarithmic scores from the SNP-DCC model are significantly lower than those obtained from the DCC model. It is worth noting that the observed differences are due to the density specification, since forecasted means and volatilities from both distributions are exactly the same and differences in forecasted conditional correlations are very small. Our empirical results show evidence on that the extra term in the SNP_{II} distribution with respect to the Normal (see equation (15)) provides the SNP_{II} with enough flexibility to adapt to higher and possibly irregular frequencies at the lower percentiles (see right columns of Figure 3), and to more density around its mean.⁷ These results are not surprising since leptokurtic models tend to show significantly better performance the lower the quantile of the forecasted density; besides, they are in line with previous findings in the literature: e.g. Fiorentini et al. (2003) show through the same graphical procedures used in this section, that a Student's t distribution provides a better performance than a Normal for fitting the full

⁶ Examples of strictly proper scoring rules include the logarithmic, quadratic or spherical (Matheson & Winkler (1976), Gneiting & Raftery (2007)).

⁷ The differences between the forecasted PITs under the multivariate (gaussian)-DCC and SNP-DCC models can be directly computed from the equations (14) and (15), provided that the differences between the forecasted correlations between both models are negligible.

density of stock returns, and Níguez & Perote (2004) provide evidence on gains of the PES density with respect to Gaussian and Student's t for full density forecasting, both papers consider empirical applications to asset returns in a univariate framework. On the other hand, in the multivariate context, Perote (2004) showed that multivariate SNP distributions involve more accurate in-sample fits of asset return distributions than Student's t or Gaussian, and Del Brio et al. (2009) show evidence of superior performance of multivariate SNP-CCC models with respect to Gaussian models for asset portfolio data. The results of our analysis show that SNP-DCC models provide a reasonably better performance than (Gaussian) DCC models for forecasting the full and lower percentile of the density of asset portfolio returns.

TABLE 3
Density forecasting performance

	Portfolio AT&T - JP Morgan	Portfolio FX €/£ - Nasdaq
	$\bar{\psi}(\cdot, \cdot)$ (<i>t-stat</i>)	$\bar{\psi}(\cdot, \cdot)$ (<i>t-stat</i>)
SNP marginal 1	.3987	.4194
DCC marginal 1	.4247 (-4.55)	.4256 (-3.95)
SNP marginal 2	.4271	.4318
DCC marginal 2	.4379 (-3.82)	.4382 (-2.38)
SNP conditional 1	.4278	.4319
DCC conditional 1	.4324 (-1.74)	.4391 (-2.30)
SNP conditional 2	.3995	.4195
DCC conditional 2	.4117 (-3.36)	.4258 (-3.55)

Notes: This table presents the results of average logarithmic scores and statistical tests for one-step-ahead density forecast from SNP-DCC and DCC models. $\bar{\psi}(\cdot, \cdot)$ denotes the average logarithmic scoring rule defined as negatively oriented penalty, so the lower the score the better the model. The table also gathers the results of the Amisano and Giacomini (2007) test for the significance of the difference between the average logarithmic scores: the entries are t-statistics (*t-stat*) for pairwise comparisons between $\bar{\psi}(\cdot, \cdot)$ from the SNP-DCC models versus the DCC models; a negative t-statistic means that the SNP model produces a lower $\bar{\psi}(\cdot, \cdot)$ than its DCC counterpart. Predictions 1,000.

FIGURE 3
P-value discrepancy plots

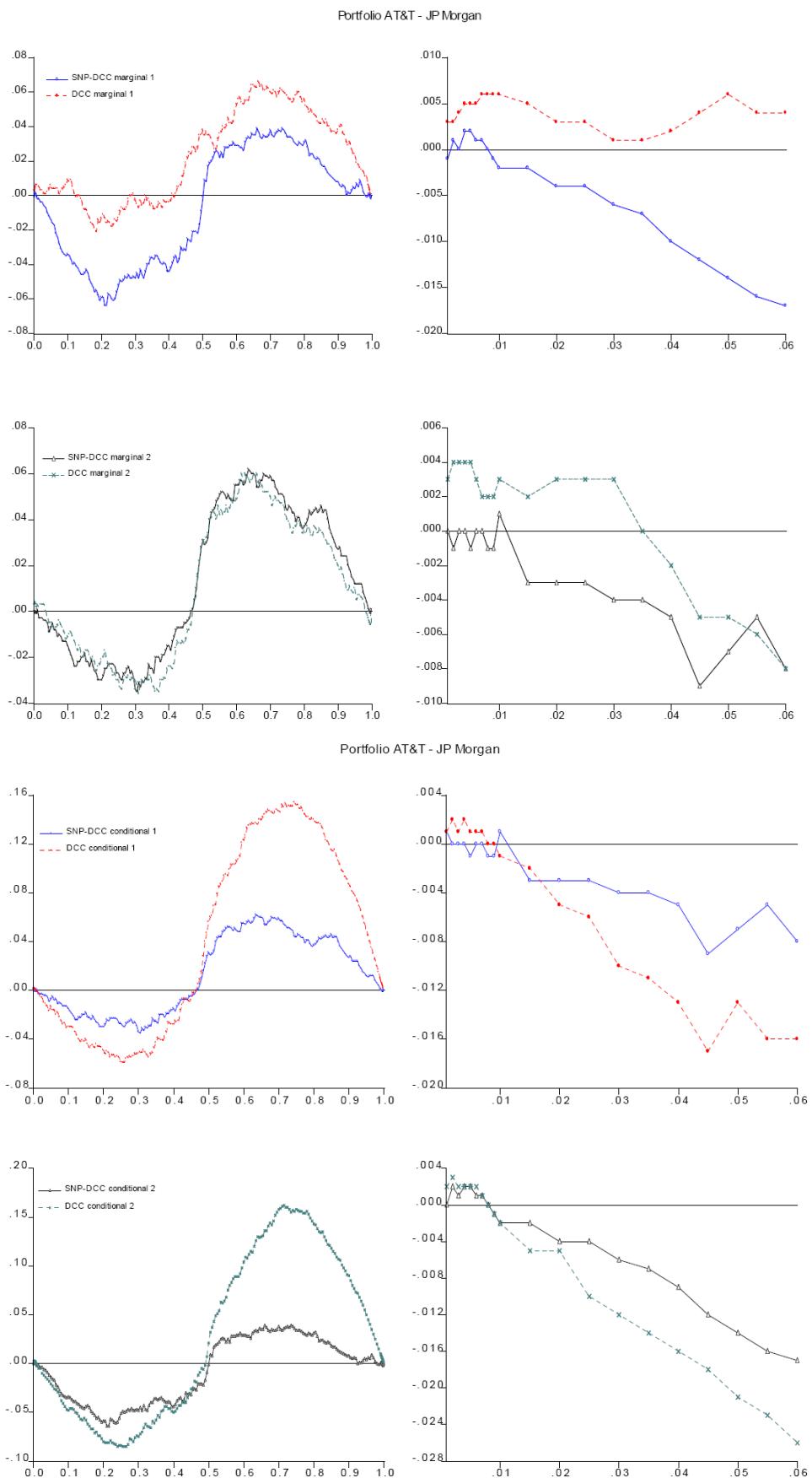
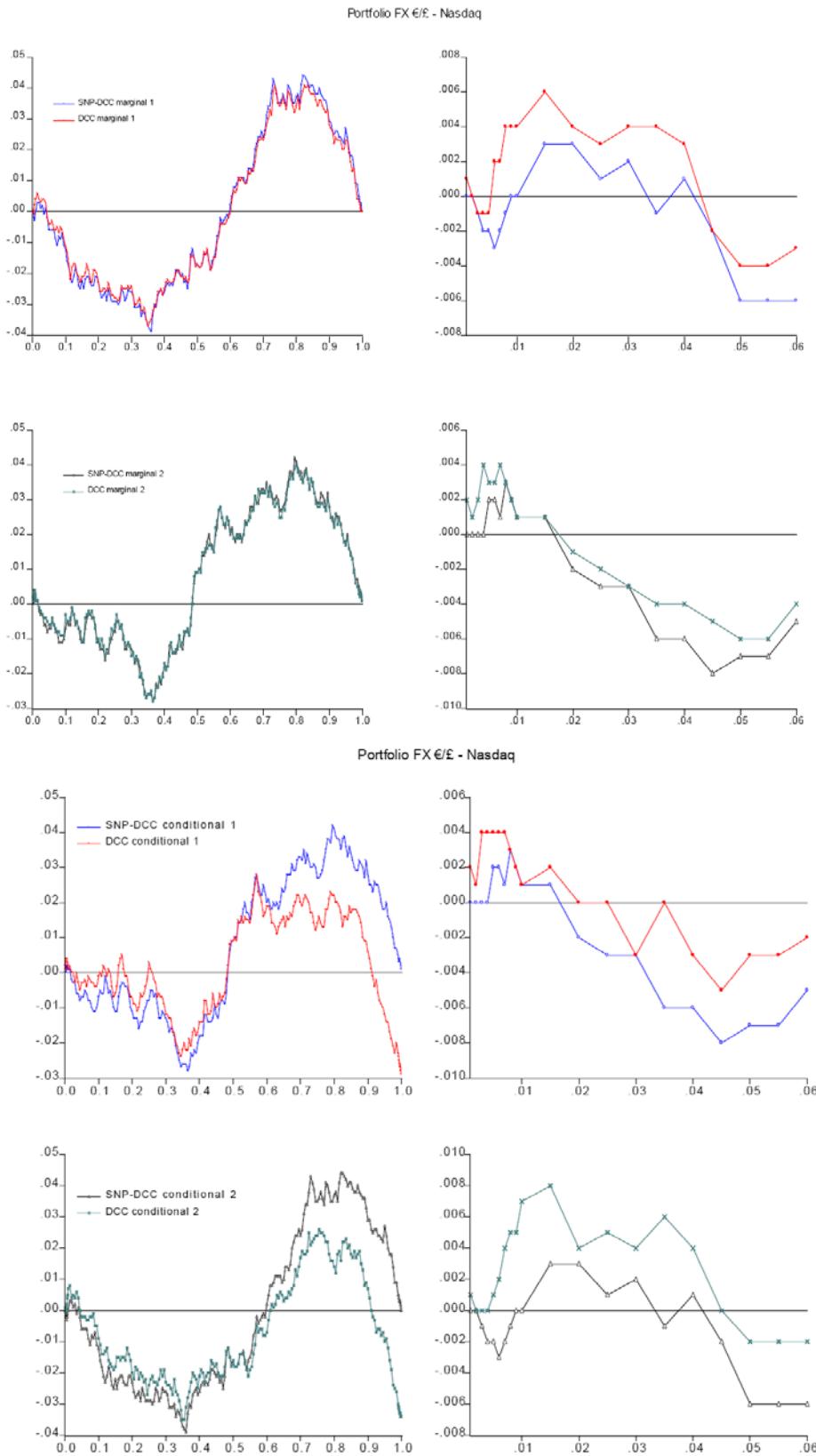


FIGURE 3
(continued)



Notes: P-value discrepancy plots of the PIT empirical cdf of marginal and conditional densities from $\text{SNP}_{II}\text{-DCC}$ and DCC models. The plots of the 2nd column highlight the models forecasting performance on the left tail of the data empirical distribution. Predictions 1000.

5. Concluding remarks

There exists abundant literature on multivariate volatility models for the time-varying first and second conditional moments of the asset returns distribution. However, not many papers have dealt with the modeling of the full distribution of financial variables. In multivariate analysis, it is usually found that the considered distributions are either not flexible enough to incorporate salient empirical regularities of financial returns (such as heavy tails, possible multimodality, skewness, etc.) or analytically intractable if they incorporate those features. The SNP densities, which are based on Edgeworth and Gram-Charlier expansions, allow to address these topics since, not only they can fit any target density through their general and flexible parametric structure, but also they present an analytical specification that is simple and easy to implement due to the orthogonal structure of Hermite polynomials (i.e., marginals, moments or cdfs can be straightforwardly computed). There have been several attempts to generalize SNP densities to a multivariate context, but the proposed specifications present serious drawbacks, such as ensuring positivity for the full density domain, achieving tractability for large portfolios and incorporating DCC models. This article covers this gap in the literature by proposing a new multivariate positive SNP density that, with a very simple structure, is capable of encompassing as marginals, the alternative univariate SNP used in financial literature. Moreover, the proposed multivariate SNP distribution admits the decomposition of the likelihood function proposed in Engle (2002). We note that this is not trivial for non-Gaussian distributions, e.g., up to the knowledge of the authors, those decompositions are neither possible to other multivariate distributions of this type such as the MES, nor to other non-Normal distributions. The implementation of the Engle's (2002) two-step estimation techniques for SNP distributions helps to solve the known "dimensionality curse" and allows to incorporate dynamic conditional correlations, nesting the (Gaussian)-DCC model in the SNP-DCC.

We compare the empirical performance of two types of multivariate SNP specifications (positive and a non-positive) for modeling the distribution of portfolio returns, in relation to the Gaussian taken as benchmark. Those models were estimated either under CCC or DCC hypotheses. Our results show the superiority of the DCC models in relation to their CCC counterparts. We also find that the considered SNP specifications outperform the Normal, and that the non-positive SNP version seems to provide a better in-sample fit. Nevertheless, since a positive SNP is required in an out-of-sample context, we compared the out-of-sample performance of the (Gaussian)-

DCC and the positive SNP-DCC by using graphical procedures and proper scoring rules and find that the latter provide overall a superior performance (particularly at the distribution tails, i.e. for Value-at-Risk forecasting).

In summary, this paper proposes a new approach to modeling the distribution of financial returns that generalizes the univariate SNP distributions to a multivariate framework guaranteeing the positivity of the density for all values of its parameters. Furthermore, the proposed model avoids the known "dimensionality curse" of the multivariate context, by means of the implementation of the DCC methodology. An empirical application to asset portfolio returns shows that the SNP-DCC model provides an overall better in- and out-of sample performance than the (Gaussian) DCC for full density forecasting.

Appendix

This appendix includes the proofs of the properties of the multivariate SNP densities presented in Section 2. *Proof 1* shows that multivariate SNP densities integrate up to one; *Proof 2*, *Proof 3* and *Proof 4* provide closed forms for the marginal distributions, the moments and the cdf, respectively; *Proof 5* shows how to compute differences between the cdfs under multivariate Normal and SNP densities, and *Proof 6* shows the separability of the log-likelihood for the SNP-DCC model.

Proof 1.

$$\begin{aligned} \int \cdots \int F^\zeta(X_{it}) dx_{1t} \cdots dx_{nt} &= \frac{1}{n} \int \cdots \int \left\{ \prod_{i=1}^n g(x_{it}) \right\} \left\{ \sum_{i=1}^n \omega_i^\zeta q_i^\zeta(x_{it}) \right\} dx_{1t} \cdots dx_{nt} \\ &= \frac{1}{n} \sum_{i=1}^n \left[\omega_i^\zeta \int g(x_{it}) q_i^\zeta(x_{it}) dx_{it} \prod_{j=1, j \neq i}^n \int g(x_{jt}) dx_{jt} \right] = \frac{1}{n} n = 1, \quad \forall \zeta = \text{I, II}. \end{aligned}$$

Proof 2.

$$\begin{aligned} f_i^\zeta(x_{it}) &= \int \cdots \int F^\zeta(X_{it}) dx_{1t} \cdots dx_{i-1,t} dx_{i+1,t} \cdots dx_{nt} \\ &= \frac{1}{n} g(x_{it}) \omega_i^\zeta q_i^\zeta(x_{it}) \prod_{j=1, j \neq i}^n \int g(x_{jt}) dx_{jt} + \frac{1}{n} g(x_{it}) \sum_{j=1, j \neq i}^n \left[\prod_{l=1, l \neq i}^n \int \omega_l^\zeta g(x_{lt}) q_l^\zeta(x_{lt}) dx_{lt} \right] \end{aligned}$$

$$= \frac{1}{n} g(x_{it}) \omega_i^\varsigma q_i^\varsigma(x_{it}) + \frac{n-1}{n} g(x_{it}), \quad \forall \varsigma = \text{I, II}.$$

Furthermore, if $\zeta=1$ (i.e. $\omega_i^I = 1$) then

$$f_i^I(x_{it}) = g(x_{it})[(n-1) + q_i^I(x_{it})] \frac{1}{n} = g(x_{it}) \left[1 + \sum_{s=2}^n \frac{\gamma_{is}}{n} H_s(x_{it}) \right].$$

Proof 3.

$$E^I[x_{it}^r] = \int x_{it}^r f_i^I(x_{it}) dx_{it} = \int x_{it}^r g(x_{it}) dx_{it} + \sum_{s=2}^m \frac{\gamma_{is}}{n} \int x_{it}^r H_s(x_{it}) g(x_{it}) dx_{it}$$

$$= \mu_r + \sum_{s=2}^m \frac{\gamma_{is}}{n} \sum_{j=0}^r c_j \int H_j(x_{it}) H_s(x_{it}) g(x_{it}) dx_{it} = \mu_r + \sum_{j=0}^r j! c_j \frac{\gamma_{ij}}{n}.$$

$$E^{II}[x_{it}^r] = \int x_{it}^r f_i^{II}(x_{it}) dx_{it}$$

$$= \frac{n-1}{n} \int x_{it}^r f_i^I(x_{it}) dx_{it} + \frac{1}{n \omega_i^{II}} \int x_{it}^r g(x_{it}) dx_{it} + \sum_{s=2}^m \gamma_{is}^2 \int x_{it}^r H_s(x_{it})^2 g(x_{it}) dx_{it}$$

$$= \left[\frac{n-1}{n} + \frac{1}{n \omega_i^{II}} \right] \mu_r + \frac{1}{n \omega_i^{II}} \sum_{s=2}^m \gamma_{is}^2 \sum_{j=0}^{r/2} d_j \int H_j(x_{it})^2 H_s(x_{it})^2 g(x_{it}) dx_{it}$$

$$= \left[\frac{n-1}{n} + \frac{1}{n \omega_i^{II}} \right] \mu_r + \frac{1}{n \omega_i^{II}} \sum_{s=2}^m s! \gamma_{is}^2 \sum_{j=0}^{r/2} j! d_j.$$

Proof 4.

$$\begin{aligned} \Pr[x_1 \leq a_1, \dots, x_n \leq a_n]^\varsigma &= \frac{1}{n} \int_{-\infty}^{a_1} \cdots \int_{-\infty}^{a_n} \left[\prod_{i=1}^n g(x_{it}) \right] \left\{ \sum_{i=1}^n \omega_i^\varsigma q_i^\varsigma(x_{it}) \right\} dx_{1t} \cdots dx_{nt} \\ &= \frac{1}{n} \sum_{i=1}^n \int_{-\infty}^{a_i} \omega_i^\varsigma g(x_{it}) q_i^\varsigma(x_{it}) dx_{it} \prod_{j=1, j \neq i}^n \int_{-\infty}^{a_j} g(x_{jt}) dx_{jt} = \frac{1}{n} \sum_{i=1}^n \Psi^\varsigma(a_i) \left[\prod_{j=1, j \neq i}^n \Phi(a_j) \right], \quad \forall \varsigma = \text{I, II}. \end{aligned}$$

Proof 5.

$$\begin{aligned}
& \Pr[x_1 \leq a_1, \dots, x_n \leq a_n]^N - \Pr[x_1 \leq a_1, \dots, x_n \leq a_n]^\varsigma = \prod_{i=1}^n \Phi(a_i) - \frac{1}{n} \sum_{i=1}^n \Psi^\varsigma(a_i) \left[\prod_{j=1, j \neq i}^n \Phi(a_j) \right] \\
&= \prod_{i=1}^n \Phi(a_i) - \frac{1}{n} \sum_{i=1}^n \left[\Phi(a_i) - B^\varsigma(a_i) \left[\prod_{j=1, j \neq i}^n \Phi(a_j) \right] \right] \\
&= \prod_{i=1}^n \Phi(a_i) - \frac{1}{n} \sum_{i=1}^n \left[\prod_{j=1}^n \Phi(a_i) - B^\varsigma(a_i) \prod_{j=1, j \neq i}^n \Phi(a_j) \right] = \frac{1}{n} \sum_{i=1}^n B^\varsigma(a_i) \left[\prod_{j=1, j \neq i}^n \Phi(a_j) \right], \quad \forall \varsigma = \text{I, II}.
\end{aligned}$$

Proof 6.

$$\begin{aligned}
L_{SNP}^\varsigma(\Phi, \rho, \gamma) &= -\frac{1}{2} \sum_{t=1}^T \left[n \log(2\pi) + \ln |\Sigma_t| + (Y_t - \mu_t)' \Sigma_t^{-1} (Y_t - \mu_t) \right] \\
&\quad + \sum_{t=1}^T \ln \left[\sum_{i=1}^n \omega_i^\varsigma q_i^\varsigma (\Sigma_t^{-1/2} (Y_t - \mu_t)) \right] - T \ln(n) \\
&= -\frac{1}{2} \sum_{t=1}^T \left\{ n \log(2\pi) + \ln |D_t R_t D_t| + (Y_t - \mu_t)' D_t^{-1} R_t^{-1} D_t^{-1} (Y_t - \mu_t) \right\} \\
&\quad + \sum_{t=1}^T \ln \left[\sum_{i=1}^n \omega_i^\varsigma q_i^\varsigma (R_t^{-1/2} D_t^{-1} (Y_t - \mu_t)) \right] - T \ln(n) \\
&= -\frac{1}{2} \sum_{t=1}^T \left\{ n \log(2\pi) + 2 \ln |D_t| + \ln |R_t| + E_t' R_t^{-1} E_t \right\} + \sum_{t=1}^T \ln \left[\sum_{i=1}^n \omega_i^\varsigma q_i^\varsigma (R_t^{-1/2} E_t) \right] - T \ln(n) \\
&= -\frac{1}{2} \sum_{t=1}^T \left\{ n \log(2\pi) + 2 \ln |D_t| + (Y_t - \mu_t)' D_t^{-2} (Y_t - \mu_t) - E_t' E_t + \ln |R_t| + E_t' R_t^{-1} E_t \right\} \\
&\quad + \sum_{t=1}^T \ln \left[\sum_{i=1}^n \omega_i^\varsigma q_i^\varsigma (R_t^{-1/2} E_t) \right] - T \ln(n) \\
&= -\frac{1}{2} \sum_{t=1}^T \sum_{i=1}^n \left(\ln(2\pi\sigma_{it}^2) + \frac{(y_{it} - \mu_{it})^2}{\sigma_{it}^2} \right) - \frac{1}{2} \sum_{t=1}^T (\ln |R_t| + E_t' R_t^{-1} E_t) \\
&\quad + \sum_{t=1}^T \ln \left[\sum_{i=1}^n \omega_i^\varsigma q_i^\varsigma (R_t^{-1/2} E_t) \right] + \frac{1}{2} \sum_{t=1}^T E_t' E_t - T \ln(n) = L_{MV}(\Phi) + L_{SNP}^\varsigma(\Phi, \varphi) + \kappa.
\end{aligned}$$

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