Estimating multivariate GARCH models equation by equation

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Abstract. This paper investigates the estimation of a wide class of multivariate volatility models. Instead of estimating a m-multivariate volatility model, a much simpler and numerically efficient method consists in estimating m univariate GARCH-type models Equation by Equation (EbE) in the first step, and a correlation matrix in the second step. Strong consistency and asymptotic normality (CAN) of the EbE estimator are established in a very general framework, including Dynamic Conditional Correlation (DCC) models. The EbE estimator can be used to test the restrictions imposed by a particular MGARCH specification. For general Constant Conditional Correlation (CCC) models, we obtain the CAN of the two-step estimator. Comparisons with the global method, in which the model parameters are estimated in one step, are provided. Monte-Carlo experiments and applications to financial series illustrate the interest of the approach.

Keywords: Constant conditional correlation, Dynamic conditional correlation, Multivariate GARCH specification testing, Quasi maximum likelihood estimation.

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1. Introduction

Generalized Autoregressive Conditional Heteroskedasticity (GARCH) models have featured prominently in the analysis of financial time series. The last twenty years have witnessed significant research devoted to the multivariate extension of the concepts and models initially developed for univariate GARCH. Among the numerous specifications of multivariate GARCH (MGARCH) models, the most popular seem to be the Constant Conditional Correlations (CCC) model introduced by Bollerslev (1990) and extended by Jeantheau (1998), the Baba, Engle, Kraft and Kroner (BEKK) model of Engle and Kroner (1995), and the Dynamic Conditional Correlations (DCC) models proposed by Tse and Tsui (2002) and Engle (2002). Reviews on the rapidly changing literature on MGARCH are Bauwens, Laurent and Rombouts (2006), Silvennoinen and Teräsvirta (2009), Francq and Zakoïan (2010, Chapter 11), Bauwens, Hafner and Laurent (2012), Tsay (2014, Chapter 7).

The complexity of MGARCH models has been a major obstacle to their use in applied works. Indeed, in asset pricing applications or portfolio management, cross-sections of hundreds of stocks are common. However, as the dimension of the cross section increases, the number of parameters can become very large in MGARCH models, making estimation increasingly cumbersome. This "dimensionality curse" is general in multivariate time series, but is particularly problematic in GARCH models. The reason is that the parameters of interest are involved in the conditional variance matrix, which has to be inverted in Gaussian likelihood-based estimation methods. Existing approaches to alleviate the dimensionality curse rely on either constraining the structure of the model in order to reduce the number of parameters (see e.g. Engle, Ng and Rotschild (1990), van der Weide (2002), Lanne and Saikkonen (2007)), or using an alternative estimation criterion (see Engle, Shephard and Sheppard (2008)). An approach combining the two concepts, was recently proposed by Engle and Kelly (2012).

A solution to the high-dimension problem which does not preclude a high-dimensional parameter set relies on two steps. In the first stage, univariate GARCH models are estimated for each individual series, *equation by equation*, and in the second stage, standardized residuals are used to estimate the parameters of the dynamic correlation. This approach, initially proposed by Engle and Sheppard (2001) and Engle (2002) in the context of DCC models, was advocated by Pelletier (2006) for regime-switching dynamic correlation models, by Aielli (2013) for DCC models, and it was used in several empirical studies (see e.g. Hafner and Reznikova (2012), Sucarrat, Grønneberg and Escribano (2013) for recent references).

However, the statistical properties of such two-step estimators have not been established.[1]

The first goal of the present paper is to develop asymptotic results for the Equation-by-Equation (EbE) estimator *of the volatility parameters*, based on the Quasi-Maximum Likelihood (QML). Our framework for the individual volatilities specification is extremely general. First, the conditional variance of component k is a parametric function of the past of *all components* of the vector of returns. This allows to capture serial dependencies between components, that do not appear in the conditional correlation matrix. Second, the volatilities, being specified as any parametric functions of the past returns, are able to accommodate leverage-effects or any other type of "nonlinearity". One issue of interest, that we will also investigate, is whether individual estimation of the conditional variances necessarily entails an efficiency loss (asymptotically and/or in finite sample) with respect to a global QML method which estimates them jointly.

Apart from the numerical simplicity, one advantage of this approach is that the derivation of EbE estimators (EbEE) is independent from the specification of a conditional correlation matrix. It can therefore be employed for CCC as well as for DCC GARCH models, leading to the same estimators of the individual volatilities.

Another aim of this paper is to provide asymptotic results for the second step of the two-stage approach, that is the estimation of a time-varying correlation matrix using the standardized returns obtained in the first step. At this stage, a specification of the conditional correlation dynamic is required. For CCC models, the constant conditional correlation matrix can be estimated by the empirical correlation matrix of the EbEE residuals. In this article, we derive asymptotic results for this estimator, which can be seen as an extension of the two-step estimator proposed by Engle and Sheppard (2001) in the case where the individual volatilities have pure GARCH forms with iid innovations. For DCC models, the structure of the time-varying correlation can also be estimated.

The paper is organized as follows. Section 2 presents the main assumptions and notations. In Section 3, we study the estimation of the volatility parameters without any assumption on $\boldsymbol{R}_t$. Particular parameterizations are discussed in Section 4. Section 5 develops the two-step estimation method when the correlation matrix $\boldsymbol{R}_t$ is parameterized. Numerical illustrations are presented in Section 6. Section 7 concludes. The most technical assumptions and the proofs of the main theorems are collected in the Appendix. Other proofs, along with additional numerical illustrations, are included in a supplementary file.

[1]See the recent survey by Caporin and McAleer (2012) for a discussion of the existence, or the absence, of asymptotic results for multivariate GARCH models.

2. Models and assumptions

Let $\boldsymbol{\epsilon}_t = (\epsilon_{1t}, \cdots, \epsilon_{mt})'$ be a $\mathbb{R}^m$-valued process and let $\mathcal{F}_{t-1}^{\boldsymbol{\epsilon}}$ be the σ-field generated by $\{\boldsymbol{\epsilon}_u, u < t\}$. Assume

$$E(\boldsymbol{\epsilon}_t \mid \mathcal{F}_{t-1}^{\boldsymbol{\epsilon}}) = \mathbf{0}, \qquad \mathrm{Var}(\boldsymbol{\epsilon}_t \mid \mathcal{F}_{t-1}^{\boldsymbol{\epsilon}}) = \boldsymbol{H}_t \quad \text{exists and is positive definite.} \tag{2.1}$$

Denoting by σ_{kt}^2 the diagonal elements of $\boldsymbol{H}_t$, that is the variances of the ϵ_{kt} conditional on $\mathcal{F}_{t-1}^{\boldsymbol{\epsilon}}$, let $\boldsymbol{\eta}_t^* = \boldsymbol{D}_t^{-1}\boldsymbol{\epsilon}_t = (\epsilon_{1t}/\sigma_{1t}, \ldots, \epsilon_{mt}/\sigma_{mt})'$ where $\boldsymbol{D}_t = \mathrm{diag}(\sigma_{1t}, \ldots, \sigma_{mt})$. By (2.1) we have, $E(\boldsymbol{\eta}_t^* \mid \mathcal{F}_{t-1}^{\boldsymbol{\epsilon}}) = \mathbf{0}$ and the conditional correlation matrix of $\boldsymbol{\epsilon}_t$ is given by

$$\boldsymbol{R}_t = \mathrm{Var}(\boldsymbol{\eta}_t^* \mid \mathcal{F}_{t-1}^{\boldsymbol{\epsilon}}) = \boldsymbol{D}_t^{-1}\boldsymbol{H}_t\boldsymbol{D}_t^{-1}. \tag{2.2}$$

It follows that the components η_{kt}^* of $\boldsymbol{\eta}_t^*$ satisfy, for $k = 1, \ldots, m$,

$$E(\eta_{kt}^* \mid \mathcal{F}_{t-1}^{\boldsymbol{\epsilon}}) = 0, \qquad \mathrm{Var}(\eta_{kt}^* \mid \mathcal{F}_{t-1}^{\boldsymbol{\epsilon}}) = 1. \tag{2.3}$$

Introducing the vector $\boldsymbol{\eta}_t$ such that $\boldsymbol{\eta}_t^* = \boldsymbol{R}_t^{1/2}\boldsymbol{\eta}_t$, the previous equations can be summarized as follows. The square root has to be understood in the sense of the Cholesky factorization, that is, $\boldsymbol{R}_t^{1/2}(\boldsymbol{R}_t^{1/2})' = \boldsymbol{R}_t$ and $\boldsymbol{H}_t^{1/2}(\boldsymbol{H}_t^{1/2})' = \boldsymbol{H}_t$.

Assumptions and notations: *The $\mathbb{R}^m$-valued process $(\boldsymbol{\epsilon}_t)$ satisfies*

$$\left\{ \begin{array}{rcl} \boldsymbol{\epsilon}_t & = & \boldsymbol{H}_t^{1/2}\boldsymbol{\eta}_t, \qquad E(\boldsymbol{\eta}_t \mid \mathcal{F}_{t-1}^{\boldsymbol{\epsilon}}) = \mathbf{0}, \quad Var(\boldsymbol{\eta}_t \mid \mathcal{F}_{t-1}^{\boldsymbol{\epsilon}}) = \boldsymbol{I}_m, \\ \\ \boldsymbol{H}_t & = & \boldsymbol{H}(\boldsymbol{\epsilon}_{t-1}, \boldsymbol{\epsilon}_{t-2}, \ldots) = \boldsymbol{D}_t\boldsymbol{R}_t\boldsymbol{D}_t, \end{array} \right. \tag{2.4}$$

where $\boldsymbol{H}_t$ is positive definite, $\boldsymbol{D}_t = \{diag(\boldsymbol{H}_t)\}^{1/2}$ and $\boldsymbol{R}_t = Corr(\boldsymbol{\epsilon}_t, \boldsymbol{\epsilon}_t \mid \mathcal{F}_{t-1}^{\boldsymbol{\epsilon}})$.

We assume that σ_{kt}^2 is parameterized by some parameter $\boldsymbol{\theta}_0^{(k)} \in \mathbb{R}^{d_k}$, so that

$$\left\{ \begin{array}{rcl} \epsilon_{kt} & = & \sigma_{kt}\eta_{kt}^*, \\ \sigma_{kt} & = & \sigma_k(\boldsymbol{\epsilon}_{t-1}, \boldsymbol{\epsilon}_{t-2}, \ldots; \boldsymbol{\theta}_0^{(k)}), \end{array} \right. \tag{2.5}$$

where σ_k is a positive function. In view of (2.3), the process $(\boldsymbol{\eta}_t^*)$ can be called the vector of EbE innovations of $(\boldsymbol{\epsilon}_t)$.

Remark 2.1. In Model (2.4)-(2.5), the volatility of any component of $\boldsymbol{\epsilon}_t$ is allowed to depend on the past values of all components. This assumption represents an extension of the classical set up of univariate GARCH models and, for this reason, Model (2.5) can be referred to as an *augmented GARCH* model in the terminology of Hörmann (2008).

This extension is firstly motivated by the sake of generality: it seems very restrictive to assume that the conditional variance of a component is not influenced by the past of other components. On the other hand, the EbE estimation approach of the paper makes this extension amenable to statistical inference, without causing an explosion in the number of parameters. For instance, if the individual volatilities have GARCH(1,1)-type dynamics,

$$\sigma_{kt}^2 = \omega_k + \sum_{\ell=1}^{m} \alpha_{k,\ell}\epsilon_{\ell,t-1}^2 + \beta_k\sigma_{k,t-1}^2, \quad \omega_k > 0, \alpha_{k,\ell} \geq 0, \beta_k \geq 0, \tag{2.6}$$

increasing by K the number of components entails an additional number of K parameters by equation. Finally, this extension allows to tackle the problem of asynchronous data by allowing each conditional variance to depend on the most recent observations. See Section 6.2.1 for more details on this issue.

Remark 2.2. A variety of parametric forms of function $\boldsymbol{H}$ have been introduced in the literature. In particular, a standard specification of diag$(\boldsymbol{H}_t)$ is, in vector form, given by

$$\boldsymbol{h}_t = \boldsymbol{\omega} + \sum_{i=1}^{q} \boldsymbol{A}_i \underline{\boldsymbol{\epsilon}}_{t-i} + \sum_{j=1}^{p} \boldsymbol{B}_j \boldsymbol{h}_{t-j} \tag{2.7}$$

where $\boldsymbol{h}_t = \left(\sigma_{1t}^2, \cdots, \sigma_{mt}^2\right)'$, $\underline{\boldsymbol{\epsilon}}_t = \left(\epsilon_{1t}^2, \cdots, \epsilon_{mt}^2\right)'$, $\boldsymbol{A}_i$ and $\boldsymbol{B}_j$ are $m \times m$ matrices with positive coefficients and $\boldsymbol{\omega} = (\omega_1, \cdots, \omega_m)'$ is a vector of strictly positive coefficients. Conrad and Karanasos (2010) provide less restrictive assumptions that ensure the positive definiteness of $\boldsymbol{H}_t$, and also show that there exists a representation of the form (2.7) in which the $\boldsymbol{B}_j$'s are diagonal. When $p = q = 1$ and $\boldsymbol{B}_1$ is diagonal, the individual volatilities satisfy a dynamic of the form (2.6).

Remark 2.3. The positivity of the function σ_k generally entails restrictions on the parameter values which cannot be made explicit under the general formulation. For particular models such constraints can be explicited, as in (2.6). Note that the EbE innovations η_{kt}^* are not iid in general, and thus (2.5) is not a Data Generating Process (DGP).

GARCH-type models constitute the most important class of DGP satisfying the previous assumptions. Consider a GARCH process, defined as a non anticipative[2] solution of

$$\boldsymbol{\epsilon}_t = \boldsymbol{D}_t \boldsymbol{R}_t^{1/2} \boldsymbol{\eta}_t, \quad \text{where } (\boldsymbol{\eta}_t) \text{ is an iid sequence.} \tag{2.8}$$

Obviously, $(\boldsymbol{\epsilon}_t)$ thus satisfies (2.4). In this paper, we will distinguish CCC models, for which

$$\boldsymbol{R}_t = \boldsymbol{R} \text{ is a constant correlation matrix,} \tag{2.9}$$

[2]that is $\boldsymbol{\epsilon}_t \in \mathcal{F}_t^{\boldsymbol{\eta}}$, the σ-field generated by $\{\boldsymbol{\eta}_u, u \leq t\}$.

from DCC models where $\boldsymbol{R}_t$ is a non constant function of the past of $\boldsymbol{\epsilon}_t$, that is,

$$\boldsymbol{R}_t = \boldsymbol{R}(\boldsymbol{\epsilon}_{t-1}, \boldsymbol{\epsilon}_{t-2}, \dots) \neq \boldsymbol{R}.$$

Note that in the case of CCC models, the sequence $(\boldsymbol{\eta}_t^*)$ is iid which is generally not the case for DCC models. In the econometric literature, CCC models are generally introduced under the specification (2.7) of the individual conditional variances.[3] To avoid confusion, we will refer to (2.7)-(2.9) as the CCC-GARCH(p, q) model.

3. Equation-by-equation estimation of volatility parameters in MGARCH models

In this section, we are interested in estimating the conditional variance of each component of $\boldsymbol{\epsilon}_t$ satisfying (2.4). In other words, we study the estimation of the parameter $\boldsymbol{\theta}_0^{(k)}$ in the augmented GARCH model (2.5), under (2.3), for $k = 1, \dots, m$. We will use the Gaussian QML, but other methods could be considered as well (for instance the weighted QML studied by Ling (2007), the non Gaussian QML studied by Berkes and Horváth (2004)). In view of Remarks 2.1 and 2.3, the augmented GARCH model (2.5) is not, in general, a univariate GARCH and we cannot directly rely on existing results for its estimation.

Given observations $\boldsymbol{\epsilon}_1, \dots, \boldsymbol{\epsilon}_n$, and arbitrary initial values $\tilde{\boldsymbol{\epsilon}}_i$ for $i \leq 0$, we define $\tilde{\sigma}_{kt}(\boldsymbol{\theta}^{(k)}) = \sigma_k(\boldsymbol{\epsilon}_{t-1}, \boldsymbol{\epsilon}_{t-2}, \dots, \boldsymbol{\epsilon}_1, \tilde{\boldsymbol{\epsilon}}_0, \tilde{\boldsymbol{\epsilon}}_{-1}, \dots; \boldsymbol{\theta}^{(k)})$ for $k = 1, \dots, m$ and $\boldsymbol{\theta}^{(k)} \in \boldsymbol{\Theta}_k$, assuming that $\boldsymbol{\Theta}_k$ is a compact parameter set and $\boldsymbol{\theta}_0^{(k)} \in \boldsymbol{\Theta}_k$. This random variable will be approximated by $\sigma_{kt}(\boldsymbol{\theta}^{(k)}) = \sigma_k(\boldsymbol{\epsilon}_{t-1}, \boldsymbol{\epsilon}_{t-2}, \dots; \boldsymbol{\theta}^{(k)})$.

Let $\hat{\boldsymbol{\theta}}_n^{(k)}$ denote the equation-by-equation estimator of $\boldsymbol{\theta}_0^{(k)}$:

$$\hat{\boldsymbol{\theta}}_n^{(k)} = \arg \min_{\boldsymbol{\theta}^{(k)} \in \boldsymbol{\Theta}^{(k)}} \tilde{Q}_n^{(k)}(\boldsymbol{\theta}^{(k)}), \quad \tilde{Q}_n^{(k)}(\boldsymbol{\theta}^{(k)}) = \frac{1}{n} \sum_{t=1}^{n} \log \tilde{\sigma}_{kt}^2 \left(\boldsymbol{\theta}^{(k)}\right) + \frac{\epsilon_{kt}^2}{\tilde{\sigma}_{kt}^2 \left(\boldsymbol{\theta}^{(k)}\right)}.$$

3.1. *Consistency and asymptotic normality of the EbEE*

We make the following assumption on the process $(\boldsymbol{\epsilon}_t)$.

A1: $(\boldsymbol{\epsilon}_t)$ is a strictly stationary and ergodic process satisfying (2.4), with $E|\epsilon_{kt}|^s < \infty$ for some $s > 0$. Moreover, $E \log \sigma_{kt}^2 < \infty$.

This assumption will be made more explicit for specific models in Section 4 (see also Theorem 2.1 and Corollary 2.2 in Francq and Zakoïan (2012)). Technical assumptions on the

[3]Bollerslev (1990) introduced this model in the case of diagonal matrices $\boldsymbol{A}_i$ and $\boldsymbol{B}_j$. Ling and McAleer (2003) proved the asymptotic properties of a general version of this model (without any diagonality assumption) subsequently called the *Extended* CCC model by He and Teräsvirta (2004).

function σ_k are relegated to Appendix A. We also assume the existence of a lower bound, ensuring that the criterion be well defined for any parameter value.

A4: we have $\sigma_{kt}(\cdot) > \underline{\omega}$ for some $\underline{\omega} > 0$.

To prove the asymptotic normality, we need to assume

A7: $\boldsymbol{\theta}_0^{(k)}$ belongs to the interior of $\boldsymbol{\Theta}^{(k)}$,

A8: $E\left|\eta_{kt}^*\right|^{4(1+\delta)} < \infty$, for some $\delta > 0$,

and some additional technical assumptions **A9-A12**.

THEOREM 3.1. *If* **A1** *and* **A4-A6** *hold, the EbEE of* $\boldsymbol{\theta}_0^{(k)}$ *in the augmented GARCH model* (2.5) *is strongly consistent:* $\hat{\boldsymbol{\theta}}_n^{(k)} \to \boldsymbol{\theta}_0^{(k)}$, *a.s. as* $n \to \infty$.

If, in addition, **A7-A12** *hold, then*

$$\sqrt{n}\left(\hat{\boldsymbol{\theta}}_n^{(k)} - \boldsymbol{\theta}_0^{(k)}\right) \xrightarrow{\mathcal{L}} \mathcal{N}\left\{0, \boldsymbol{J}_{kk}^{-1}\boldsymbol{I}_{kk}\boldsymbol{J}_{kk}^{-1}\right\},$$

where $\boldsymbol{I}_{kk} = E\left(\{\eta_{kt}^{*4} - 1\}\boldsymbol{d}_{kt}\boldsymbol{d}_{kt}'\right)$, $\boldsymbol{J}_{kk} = E\left(\boldsymbol{d}_{kt}\boldsymbol{d}_{kt}'\right)$, $\boldsymbol{d}_{kt} = \frac{1}{\sigma_{kt}^2}\frac{\partial\sigma_{kt}^2(\boldsymbol{\theta}_0^{(k)})}{\partial\boldsymbol{\theta}^{(k)}}$.

REMARK 3.1. The sequence $(\boldsymbol{\eta}_t)$ in (2.4) is only assumed to be a conditionally homoscedastic martingale difference (not necessarily iid), as in Bollerslev and Wooldridge (1992). An analogous result was established, in the case of univariate GARCH(p,q) models, by Escanciano (2009) as an extension of Berkes, Horváth and Kokoszka (2003) and Francq and Zakoïan (2004).

REMARK 3.2. The EbE approach is particularly suited for the specification (2.7) with diagonal matrices $\boldsymbol{B}_j$, for which each component of $\boldsymbol{\theta}_0$ is only involved in one volatility equation. If in addition the matrices $\boldsymbol{A}_i$ are diagonal, more primitive assumptions can be given in Theorem 3.1 (see the supplementary file).

3.2. Comparison with the theoretical QML estimator

A question of interest is whether the EbEE approach necessarily entails an efficiency loss (the price paid for its simplicity) with respect to a QML method in which the volatility parameters are jointly estimated. To be able to write the global quasi-likelihood, it is necessary to specify the conditional correlation matrix. Because we wish to compare the estimators of the volatility parameters, we consider the global QML estimator (QMLE) of $\boldsymbol{\theta}_0$ based on the assumption that the matrix $\boldsymbol{R}_t$ is constant and is known.[4]

[4]We can thus call this estimator theoretical QMLE, or infeasible QMLE.

A theoretical QMLE of $\boldsymbol{\theta}_0$ is defined as any measurable solution $\hat{\boldsymbol{\theta}}_n^{QML}$ of

$$\hat{\boldsymbol{\theta}}_n^{QML} = \arg\min_{\boldsymbol{\theta}\in\Theta} n^{-1}\sum_{t=1}^{n}\tilde{\ell}_t(\boldsymbol{\theta}), \qquad \tilde{\ell}_t(\boldsymbol{\theta}) = \boldsymbol{\epsilon}_t'\tilde{\boldsymbol{H}}_t^{-1}\boldsymbol{\epsilon}_t + \log|\tilde{\boldsymbol{H}}_t|,$$

where $\tilde{\boldsymbol{H}}_t = \tilde{\boldsymbol{D}}_t\boldsymbol{R}\tilde{\boldsymbol{D}}_t$ and $\tilde{\boldsymbol{D}}_t = \mathrm{diag}(\tilde{\sigma}_{1t}(\boldsymbol{\theta}^{(1)}),\dots,\tilde{\sigma}_{mt}(\boldsymbol{\theta}^{(m)}))$. Let $\boldsymbol{R}^{-1} = (r_{k\ell}^*)$. Let the $d\times d$ matrix $\boldsymbol{M} = (\boldsymbol{M}_{k\ell})$ where $\boldsymbol{M}_{k\ell} = \tau_{k\ell}\boldsymbol{J}_{k\ell} - \sum_{i,j=1}^{m}\xi_{ki}\xi_{j\ell}(\kappa_{ij}-1)\boldsymbol{J}_{ki}\boldsymbol{J}_{ii}^{-1}\boldsymbol{J}_{ij}\boldsymbol{J}_{jj}^{-1}\boldsymbol{J}_{j\ell}$, and

$$\kappa_{k\ell} = E\left(\eta_{kt}^{*2}\eta_{\ell t}^{*2}\right), \quad \xi_{k\ell} = \frac{1}{2}\begin{cases} r_{kk}^* + 1 & \text{if} \quad k=\ell \\ r_{k\ell}r_{k\ell}^* & \text{if} \quad k\neq\ell \end{cases}, \quad \tau_{k\ell} = \sum_{i,j=1}^{m} r_{ki}^* r_{\ell j}^* E(\eta_{kt}^*\eta_{it}^*\eta_{jt}^*\eta_{\ell t}^*) - 1.$$

PROPOSITION 3.1. *Under the assumptions of Theorem 3.1, the QMLE of the volatility parameters, assuming that $\boldsymbol{R}_t = \boldsymbol{R}$ is known, is asymptotically more efficient (resp. less efficient) than the EbEE if and only if $\boldsymbol{M}$ is negative definite (resp. positive definite).*

When $\boldsymbol{R}$ is the identity matrix, $\boldsymbol{M} = \boldsymbol{0}$ and the two methods are equivalent, producing the same estimators. In practical implementation of the QML, matrix $\boldsymbol{R}$ has to be estimated, which may lower the accuracy of the volatility parameters estimators. It is interesting to note that the QMLE is not always asymptotically more efficient than the EbEE, even in the favorable situation where $\boldsymbol{R}$ is known (which has no consequence for the EbEE). Calculations reported in the supplementary file show that, in the particular case where the only unknown coefficients are the parameters of the first volatility, $\boldsymbol{M} = \{\tau_{11} - \xi_{11}^2(\kappa_{11}-1)\}J_{11}$. Thus $\boldsymbol{M}$ is positive definite if and only if $\tau_{11} > \xi_{11}^2(\kappa_{11}-1)$, whatever $\boldsymbol{\theta}_0$. It follows that the EbEE may be asymptotically superior to the QMLE when the distribution of $(\boldsymbol{\eta}_t^*)$ is sufficiently far from the Gaussian. In the general case, it does not seem possible to characterize more explicitly the models and errors distribution for which one estimation method asymptotically outperforms the other. However, matrix $\boldsymbol{M}$ can be consistently estimated, making it possible to compare the two methods.

3.3. Asymptotic results for strong augmented GARCH models

The asymptotic distribution of the EbEE can be simplified under the assumption that

$$\eta_{kt}^* \text{ is independent from } \mathcal{F}_{t-1}^{\boldsymbol{\epsilon}}. \tag{3.1}$$

Moreover, **A8** can be replaced by the weaker assumption

A8*: $E|\eta_{kt}^*|^4 < \infty$,

and the technical assumptions **A10** on the volatility function can be slightly weakened (see **A10*** in Appendix A). The asymptotic distribution of the EbEE is modified as follows.

THEOREM 3.2. *Under* (3.1) *and the assumptions of Theorem 3.1, with* **A8** *replaced by* **A8*** *and* **A10** *replaced by* **A10****, we have* $\sqrt{n}\left(\hat{\boldsymbol{\theta}}_n^{(k)} - \boldsymbol{\theta}_0^{(k)}\right) \xrightarrow{\mathcal{L}} \mathcal{N}\left\{0, (E\eta_{kt}^{*4} - 1)\boldsymbol{J}_{kk}^{-1}\right\}$.

It can be noted that (3.1) is always satisfied in the CCC case, that is, under (2.8) and $\boldsymbol{R}_t = \boldsymbol{R}$. The next result shows that (3.1) can be satisfied for other GARCH-type models.

PROPOSITION 3.2. *Assume that the distribution of* $\boldsymbol{\eta}_t$ *is spherical in Model* (2.8)*. Then* (3.1) *is satisfied. Moreover,* (η_{kt}^*) *is an iid (0,1) sequence.*

REMARK 3.3. It is worth noting that, under the assumptions of Proposition 3.2, the process $(\boldsymbol{\eta}_t^*)$ is neither independent nor identically distributed in general (even if its components are iid). To see this, consider for example, for $\lambda_1, \lambda_2 \in \mathbb{R}$ and for $k \neq \ell$,

$$\lambda_1\eta_{kt}^* + \lambda_2\eta_{\ell t}^* \stackrel{d}{=} \|(\lambda_1\boldsymbol{e}_k' + \lambda_2\boldsymbol{e}_\ell')\boldsymbol{R}_t^{1/2}\|\eta_{1t} = \{\lambda_1^2 + \lambda_2^2 + 2\lambda_1\lambda_2\boldsymbol{R}_t(k,\ell)\}^{1/2}\eta_{1t},$$

conditionally on $\mathcal{F}_{t-1}^{\boldsymbol{\epsilon}}$, where $\boldsymbol{e}_k$ denotes the k-th column of $\boldsymbol{I}_m$. The variable in the right-hand side of the latter equality is in general non independent of the past values of $\boldsymbol{\eta}_t^*$.

3.4. Adding an intercept

We consider an extension of Model (2.4) in which an intercept is included. The $\mathbb{R}^m$-valued process $(\boldsymbol{y}_t)$ is supposed to satisfy

$$\begin{cases} \boldsymbol{y}_t &= \boldsymbol{\mu}_0 + \boldsymbol{H}_t^{1/2}\boldsymbol{\eta}_t, \qquad E(\boldsymbol{\eta}_t \mid \mathcal{F}_{t-1}^{\boldsymbol{y}}) = \boldsymbol{0}, \quad \mathrm{Var}(\boldsymbol{\eta}_t \mid \mathcal{F}_{t-1}^{\boldsymbol{y}}) = \boldsymbol{I}_m, \\ \boldsymbol{H}_t &= \boldsymbol{H}(\boldsymbol{y}_{t-1}, \boldsymbol{y}_{t-2}, \dots) = \boldsymbol{D}_t\boldsymbol{R}_t\boldsymbol{D}_t, \end{cases}$$

where $\boldsymbol{\mu}_0 = (\mu_0^{(1)}, \dots, \mu_0^{(m)})' \in \mathbb{R}^m$, $\boldsymbol{D}_t = \{\mathrm{diag}(\boldsymbol{H}_t)\}^{1/2}$ and $\boldsymbol{R}_t = \mathrm{Corr}(\boldsymbol{y}_t, \boldsymbol{y}_t \mid \mathcal{F}_{t-1}^{\boldsymbol{y}})$.

Letting $\boldsymbol{\eta}_t^* = \boldsymbol{D}_t^{-1}(\boldsymbol{y}_t - \boldsymbol{\mu}_0)$, we get

$$\begin{cases} y_{kt} &= \mu_0^{(k)} + \sigma_{kt}\eta_{kt}^*, \\ \sigma_{kt} &= \sigma_k(\boldsymbol{y}_{t-1}, \boldsymbol{y}_{t-2}, \dots; \boldsymbol{\theta}_0^{(k)}), \end{cases} \tag{3.2}$$

and we study the estimation of the parameter $\boldsymbol{\gamma}_0^{(k)} = (\mu_0^{(k)}, \boldsymbol{\theta}_0^{(k)'})'$ in Model (3.2), under

$$E(\eta_{kt}^* \mid \mathcal{F}_{t-1}^{\boldsymbol{y}}) = 0, \qquad \mathrm{Var}(\eta_{kt}^* \mid \mathcal{F}_{t-1}^{\boldsymbol{y}}) = 1.$$

Given observations $\boldsymbol{y}_1, \dots, \boldsymbol{y}_n$, and arbitrary initial values $\tilde{\boldsymbol{y}}_i$ for $i \leq 0$, we define $\tilde{\sigma}_{kt}(\boldsymbol{\theta}^{(k)}) = \sigma_k(\boldsymbol{y}_{t-1}, \boldsymbol{y}_{t-2}, \dots, \boldsymbol{y}_1, \tilde{\boldsymbol{y}}_0, \tilde{\boldsymbol{y}}_{-1}, \dots; \boldsymbol{\theta}^{(k)})$ for $k = 1, \dots, m$. Let also $\sigma_{kt}(\boldsymbol{\theta}^{(k)}) = \sigma_k(\boldsymbol{y}_{t-1}, \boldsymbol{y}_{t-2}, \dots; \boldsymbol{\theta}^{(k)})$. Let $\hat{\boldsymbol{\gamma}}_n^{(k)} = (\hat{\mu}_n^{(k)}, \hat{\boldsymbol{\theta}}_n^{(k)'})'$ denote the EbEE of $\boldsymbol{\gamma}_0^{(k)}$:

$$\hat{\boldsymbol{\gamma}}_n^{(k)} = \arg \min_{\boldsymbol{\gamma}^{(k)} \in M^{(k)} \times \boldsymbol{\Theta}^{(k)}} \tilde{Q}_n^{(k)}(\boldsymbol{\gamma}^{(k)}), \quad \tilde{Q}_n^{(k)}(\boldsymbol{\gamma}^{(k)}) = \frac{1}{n} \sum_{t=1}^{n} \log \tilde{\sigma}_{kt}^2 \left(\boldsymbol{\theta}^{(k)}\right) + \frac{\{\epsilon_{kt}(\mu^{(k)})\}^2}{\tilde{\sigma}_{kt}^2 \left(\boldsymbol{\theta}^{(k)}\right)},$$

where $\epsilon_{kt}(\mu^{(k)}) = y_{kt} - \mu^{(k)}$ and $M^{(k)}$ is a compact subset of $\mathbb{R}$.

THEOREM 3.3. *If* **A1** *and* **A4**-**A6** *hold, then* $\hat{\boldsymbol{\gamma}}_n^{(k)} \to \boldsymbol{\gamma}_0^{(k)}$, *a.s. as* $n \to \infty$. *If, in addition,* **A7**-**A12** *hold and* $\mu_0^{(k)}$ *belongs to the interior of* $M^{(k)}$, *then*

$$\sqrt{n}\left(\hat{\boldsymbol{\gamma}}_n^{(k)} - \boldsymbol{\gamma}_0^{(k)}\right) \stackrel{\mathcal{L}}{\to} \mathcal{N}\left\{0, \boldsymbol{\Upsilon}\right\}, \qquad \textit{where}$$

$$\boldsymbol{\Upsilon} = \begin{pmatrix} \left\{E\left(\frac{1}{\sigma_{kt}^2}\right)\right\}^{-1} & -\left\{E\left(\frac{1}{\sigma_{kt}^2}\right)\right\}^{-1} E\left(\eta_{kt}^{*3} \frac{1}{\sigma_{kt}} \boldsymbol{d}_{kt}'\right) \boldsymbol{J}_{kk}^{-1} \\ -\left\{E\left(\frac{1}{\sigma_{kt}^2}\right)\right\}^{-1} \boldsymbol{J}_{kk}^{-1} E\left(\eta_{kt}^{*3} \frac{1}{\sigma_{kt}} \boldsymbol{d}_{kt}\right) & \boldsymbol{J}_{kk}^{-1} \boldsymbol{I}_{kk} \boldsymbol{J}_{kk}^{-1} \end{pmatrix}.$$

It is interesting to note that, despite the presence of an intercept, the asymptotic variance of $\hat{\boldsymbol{\theta}}_n^{(k)}$ is the same as in Theorem 3.1. Note also that $\hat{\mu}_n^{(k)}$ and $\hat{\boldsymbol{\theta}}_n^{(k)}$ are not asymptotically independent in general. A case where the asymptotic independence holds is when (3.1) holds and $E(\eta_{kt}^{*3}) = 0$.

4. Inference in particular MGARCH models based on the EbE approach

Theorem 3.1 can be used for estimating the individual conditional variances in particular classes of MGARCH models. It can also be used for testing their adequacy, preliminary to their estimation. Indeed, most commonly used MGARCH specifications imply strong restrictions on the volatility of the individual components. We focus on the classes of DCC-GARCH and BEKK models.

4.1. Estimating the conditional variances in DCC-GARCH models

DCC-GARCH models are generally used under the assumption that the diagonal elements of $\boldsymbol{D}_t$ follow univariate GARCH(1,1) models, that is,

$$\sigma_{kt}^2 = \omega_k + \alpha_k \epsilon_{k,t-1}^2 + \beta_k \sigma_{k,t-1}^2, \quad \omega_k > 0, \alpha_k \geq 0, \beta_k \geq 0. \tag{4.1}$$

In the so-called corrected DCC (cDCC) of Aielli (2013),[5] the conditional correlation matrix is modelled as a function of the past standardized returns as

$$\boldsymbol{R}_t = \boldsymbol{Q}_t^{*-1/2}\boldsymbol{Q}_t\boldsymbol{Q}_t^{*-1/2}, \quad \boldsymbol{Q}_t = (1-\alpha-\beta)\boldsymbol{S} + \alpha\boldsymbol{Q}_{t-1}^{*1/2}\boldsymbol{\eta}_{t-1}^{*}\boldsymbol{\eta}_{t-1}^{*'}\boldsymbol{Q}_{t-1}^{*1/2} + \beta\boldsymbol{Q}_{t-1}, \tag{4.2}$$

where $\alpha, \beta \geq 0, \alpha+\beta < 1$, $\boldsymbol{S}$ is a correlation matrix, and $\boldsymbol{Q}_t^*$ is the diagonal matrix with the same diagonal elements as $\boldsymbol{Q}_t$. No formally established asymptotic results exist for the full estimation of the DCC and cDCC models. The strong consistency and asymptotic normality of the EbEE of $\boldsymbol{\theta}_0^{(k)} = (\omega_k, \alpha_k, \beta_k)'$ in (4.1) could be obtained by applying Theorem 3.1. We establish them under more explicit conditions in the following theorem.

THEOREM 4.1. *Assume that $\alpha + \beta < 1$, $\alpha_\ell + \beta_\ell < 1$, and either $\alpha_\ell\beta_\ell > 0$ or $\beta_\ell = 0$, for $\ell = 1, \dots, m$. Let $\boldsymbol{\eta}_1$ admit, with respect to the Lebesgue measure on $\mathbb{R}^m$, a positive density around 0. Suppose that $\boldsymbol{\theta}_0^{(k)} \in \boldsymbol{\Theta}_k$ where $\boldsymbol{\Theta}_k$ is any compact subset of $(0,\infty) \times [0,\infty) \times [0,1)$. Then $\hat{\boldsymbol{\theta}}_n^{(k)} \to \boldsymbol{\theta}_0^{(k)}$, a.s. as $n \to \infty$. If, in addition, $\boldsymbol{\theta}_0^{(k)}$ is an interior point of $\boldsymbol{\Theta}_k$, and $E\|\boldsymbol{\eta}_t\|^{4(1+\delta)} < \infty$, for some $\delta > 0$, then the sequence $\sqrt{n}(\hat{\boldsymbol{\theta}}_n^{(k)} - \boldsymbol{\theta}_0^{(k)})$ is asymptotically normally distributed.*

4.2. Estimating semi-diagonal BEKK models

Consider a BEKK-GARCH(p,q) model given by

$$\boldsymbol{\epsilon}_t = \boldsymbol{H}_t^{1/2}\boldsymbol{\eta}_t, \qquad \boldsymbol{H}_t = \boldsymbol{\Omega}_0 + \sum_{i=1}^q \boldsymbol{A}_{0i}\boldsymbol{\epsilon}_{t-i}\boldsymbol{\epsilon}_{t-i}'\boldsymbol{A}_{0i}' + \sum_{j=1}^p \boldsymbol{B}_{0j}\boldsymbol{H}_{t-j}\boldsymbol{B}_{0j}, \tag{4.3}$$

where $(\boldsymbol{\eta}_t)$ is an iid $\mathbb{R}^m$-valued centered sequence with $E\boldsymbol{\eta}_t\boldsymbol{\eta}_t' = \boldsymbol{I}_m$, $\boldsymbol{A}_{0i} = (a_{ik\ell})_{1\leq k,\ell\leq m}$, $\boldsymbol{B}_{0j} = \mathrm{diag}(b_{j1}, \dots, b_{jm})$, and $\boldsymbol{\Omega}_0 = (\omega_{k\ell})_{1\leq k,\ell\leq m}$ is a positive definite $m \times m$ matrix. In this model, which can be called "semi-diagonal" (as opposed to the diagonal BEKK in which both the $\boldsymbol{B}_{0j}$ and $\boldsymbol{A}_{0i}$ are diagonal matrices), the conditional variance of any return may depend on the past of all returns. We refer to McAleer, Chan, Hoti and Liebermann (2009) for more details on BEKK models. The k-th diagonal entry of $\boldsymbol{H}_t$ satisfies a stochastic

[5]In the original DCC model of Engle (2002), the dynamics of $\boldsymbol{Q}_t$ is given by

$$\boldsymbol{Q}_t = (1-\alpha-\beta)\boldsymbol{S} + \alpha\boldsymbol{\eta}_{t-1}^{*}\boldsymbol{\eta}_{t-1}^{*'} + \beta\boldsymbol{Q}_{t-1}.$$

Aielli (2013) pointed out that the commonly used estimator of $\boldsymbol{S}$ defined as the sample second moment of the standardized returns is not consistent in this formulation. Stationarity conditions for DCC models have been recently established by Fermanian and Malongo (2014).

recurrence equation of the form

$$h_{kk,t} = \omega_{kk} + \sum_{i=1}^{q} \left(\sum_{\ell=1}^{m} a_{ik\ell} \epsilon_{\ell,t-i} \right)^2 + \sum_{j=1}^{p} b_{jk}^2 h_{kk,t-j}. \tag{4.4}$$

Let $\boldsymbol{\theta}_0^{(k)} = (\omega_{kk}, \boldsymbol{a}'_{1k}, \dots, \boldsymbol{a}'_{qk}, \boldsymbol{b}_k)'$ for $k = 1, \dots, m$, where $\boldsymbol{a}'_{ik}$ denotes the k-th row of the matrix $\boldsymbol{A}_{0i}$, and $\boldsymbol{b}_k = (b_{1k}^2, \dots, b_{pk}^2)$. It is clear that an identifiability restriction is needed, $h_{kk,t}$ being invariant to a change of sign of the k-th row of any matrix $\boldsymbol{A}_i$. For simplicity, we therefore assume that $a_{ik1} > 0$ for $i = 1, \dots, q$. Let $\boldsymbol{\theta}^{(k)} = (\boldsymbol{\theta}_i^{(k)}) \in \mathbb{R}^{1+mq+p}$ denote a generic parameter value. The parameter space $\boldsymbol{\Theta}_k$ is any compact subset of

$$\left\{ \boldsymbol{\theta}^{(k)} \mid \theta_1^{(k)} > 0, \theta_2^{(k)}, \theta_{m+2}^{(k)}, \dots, \theta_{(q-1)m+2}^{(k)} > 0, \theta_{1+mq+1}^{(k)}, \dots, \theta_{1+mq+p}^{(k)} \geq 0, \sum_{j=1}^{p} \theta_{1+mq+j}^{(k)} < 1 \right\}.$$

Let

$$\boldsymbol{A}_0 = \sum_{i=1}^{q} \boldsymbol{H}_m (\boldsymbol{A}_{0i} \otimes \boldsymbol{A}_{0i}) \boldsymbol{K}'_m, \quad \boldsymbol{B}_0 = \sum_{j=1}^{p} \boldsymbol{H}_m (\boldsymbol{B}_{0j} \otimes \boldsymbol{B}_{0j}) \boldsymbol{K}'_m$$

where $\otimes$ is the Kronecker product and $\boldsymbol{H}_m$ and $\boldsymbol{K}_m$ are the usual elimination and duplication matrices.[6]

THEOREM 4.2. *Let the spectral radius of $\boldsymbol{A}_0 + \boldsymbol{B}_0$ be less than 1. Let $\boldsymbol{\eta}_1$ admit, with respect to the Lebesgue measure on $\mathbb{R}^m$, a positive density around 0, and suppose that $E|\eta_{kt}|^{4(1+\delta)} < \infty$ for some $\delta > 0$. Suppose that $\boldsymbol{\theta}_0^{(k)} \in \boldsymbol{\Theta}_k$. Then $\hat{\boldsymbol{\theta}}_n^{(k)} \to \boldsymbol{\theta}_0^{(k)}$, a.s. as $n \to \infty$. If, in addition, $\boldsymbol{\theta}_0^{(k)}$ is an interior point of $\boldsymbol{\Theta}_k$, the sequence $\sqrt{n}(\hat{\boldsymbol{\theta}}_n^{(k)} - \boldsymbol{\theta}_0^{(k)})$ is asymptotically normally distributed.*

Full BEKK models are generally considered as unfeasible for large cross-sectional dimensions (see for instance Laurent, Rombouts and Violante (2012)) and practitioners focus on diagonal, or even scalar, models. It follows from Theorem 4.2 that the diagonal elements of $\boldsymbol{\Omega}_0$ and the matrices $\boldsymbol{A}_{0i}$ and $\boldsymbol{B}_{0j}$ can be consistently estimated by successively applying the EbEE to each equation. Indeed, each parameter of the model appears in one, and only one, equation. The next result shows that the semi-diagonal BEKK-GARCH(p,q) model (4.3) can be fully estimated by the EbEE approach.

COROLLARY 4.1. *Let $\hat{\boldsymbol{A}}$ and $\hat{\boldsymbol{B}}$ denote the EbEE estimators of $\boldsymbol{A}_0$ and $\boldsymbol{B}_0$, respectively. Under the assumptions of Theorem 4.2, a consistent estimator of $\boldsymbol{\Omega}_0$ is obtained from*

$$\mathit{vech}(\hat{\boldsymbol{\Omega}}) = \left(\boldsymbol{I}_{m(m+1)/2} - \hat{\boldsymbol{A}} - \hat{\boldsymbol{B}} \right)^{-1} \mathit{vech} \left(\frac{1}{n} \sum_{t=1}^{n} \boldsymbol{\epsilon}_t \boldsymbol{\epsilon}'_t \right).$$

[6] $\boldsymbol{H}_m$ and $\boldsymbol{K}_m$ are $\frac{m(m+1)}{2} \times m^2$ matrices such that $\boldsymbol{H}_m \boldsymbol{K}'_m = \boldsymbol{I}_{m(m+1)/2}$ and $\mathrm{vec}(\boldsymbol{A}) = \boldsymbol{K}'_m \mathrm{vech}(\boldsymbol{A})$, $\quad \mathrm{vech}(\boldsymbol{A}) = \boldsymbol{H}_m \mathrm{vec}(\boldsymbol{A})$ for any symmetric $m \times m$ matrix $\boldsymbol{A}$.

For the general BEKK (without assuming diagonality of the matrices $\boldsymbol{B}_{0j}$), the asymptotic properties of the QML method were derived by Comte and Lieberman (2003), though under some high-level assumptions.[7]

4.3. Testing adequacy of BEKK models

Equation (4.4) can be viewed as a restricted form, implied by the BEKK model, of a more general volatility specification. Testing for such a restriction in this more general framework can thus be used to check the validity of the BEKK specification. For ease of presentation, we focus on the case $m = 2$ and $p = q = 1$. Letting $\boldsymbol{\theta}_0^{(k)} = (\omega_{kk}, a_{1k}^2, 2a_{1k}a_{2k}, a_{2k}^2, b_{1k}^2)'$ for $k = 1, 2$, the validity of Model (4.3) can be studied by estimating Model (2.5) for each component of $\boldsymbol{\epsilon}_t$ with, in view of (4.4),

$$\sigma_{kt}^2 = \theta_{01}^{(k)} + \theta_{02}^{(k)}\epsilon_{1,t-1}^2 + \theta_{03}^{(k)}\epsilon_{1,t-1}\epsilon_{2,t-1} + \theta_{04}^{(k)}\epsilon_{2,t-1}^2 + \theta_{05}^{(k)}\sigma_{k,t-1}^2, \quad k = 1, 2, \tag{4.5}$$

under the positivity constraints $\theta_{01}^{(k)} > 0$, $\theta_{0i}^{(k)} \geq 0, i = 2, 4, 5$. The restrictions implied by the BEEK-GARCH(1,1) model (4.3) are of the form:

$$H_0^{(k)}: \qquad |\theta_{03}^{(k)}| = 2\sqrt{\theta_{02}^{(k)}\theta_{04}^{(k)}}, \quad k = 1, 2.$$

Let

$$\boldsymbol{\Theta}^{(k)} = \boldsymbol{\Theta}_k^* \cap \left\{ \boldsymbol{\theta}^{(k)}; |\theta_3^{(k)}| \in \left[0, 2\sqrt{\theta_2^{(k)}\theta_4^{(k)}}\right]\right\},$$

where $\boldsymbol{\Theta}_k^*$ is a compact subset of $\{\theta_1^{(k)} > 0, \theta_i^{(k)} \geq 0$, for $i = 2, 4$ and $\theta_5^{(k)} \in [0, 1)\}$. Note that, under $H_0^{(k)}$, the true parameter value is at the boundary of the parameter set.

Proposition 4.1. *Let $\boldsymbol{\theta}_0^{(k)}$ belong to the interior of $\boldsymbol{\Theta}_k^*$ for $k = 1, 2$ and let $\hat{\boldsymbol{\theta}}_n^{(k)} = (\hat{\theta}_{n1}^{(k)}, \ldots, \hat{\theta}_{n5}^{(k)})'$ denote the EbEE estimator of $\boldsymbol{\theta}_0^{(k)}$ in (4.5). Let the Wald statistic for the hypothesis $H_0^{(k)}$,*

$$\boldsymbol{W}_n^{(k)} = \frac{n\left\{\hat{\theta}_{n3}^{(k)2} - 4\hat{\theta}_{n2}^{(k)}\hat{\theta}_{n4}^{(k)}\right\}^2}{\boldsymbol{X}_n'\hat{\boldsymbol{J}}_{kk}^{-1}\hat{\boldsymbol{I}}_{kk}\hat{\boldsymbol{J}}_{kk}^{-1}\boldsymbol{X}_n}, \quad \textit{where } \boldsymbol{X}_n = \left(0, 4\hat{\theta}_{n4}^{(k)}, -2\hat{\theta}_{n3}^{(k)}, 4\hat{\theta}_{n2}^{(k)}, 0\right)', \quad \hat{\eta}_{kt}^* = \frac{\epsilon_{kt}}{\tilde{\sigma}_{kt}(\hat{\boldsymbol{\theta}}_n^{(k)})}$$

and

$$\hat{\boldsymbol{J}}_{kk} = \frac{1}{n}\sum_{t=1}^n \hat{\boldsymbol{d}}_{kt}\hat{\boldsymbol{d}}_{kt}', \quad \hat{\boldsymbol{I}}_{kk} = \frac{1}{n}\sum_{t=1}^n \{\hat{\eta}_{kt}^{*4} - 1\}\hat{\boldsymbol{d}}_{kt}\hat{\boldsymbol{d}}_{kt}', \quad \hat{\boldsymbol{d}}_{kt} = \frac{1}{\tilde{\sigma}_{kt}^2(\hat{\boldsymbol{\theta}}_n)}\frac{\partial\tilde{\sigma}_{kt}^2(\hat{\boldsymbol{\theta}}_n^{(k)})}{\partial\boldsymbol{\theta}^{(k)}}.$$

[7]In particular, the model was assumed to be identifiable and the existence of eight-order moments was required for $\boldsymbol{\epsilon}_t$. On the other hand, Avarucci, Beutner and Zaffaroni (2013) showed that for the BEKK, the finiteness of the variance of the scores requires at least the existence of second-order moments of the observable process.

Then, under the assumptions of Theorem 4.2, $\boldsymbol{W}_n^{(k)}$ asymptotically follows a mixture of the χ^2 distribution with one degree of freedom and the Dirac measure at 0:

$$\boldsymbol{W}_n^{(k)} \stackrel{\mathcal{L}}{\to} \frac{1}{2}\chi^2(1) + \frac{1}{2}\delta_0, \qquad \text{as } n \to \infty.$$

In view of this result, testing $H_0^{(k)}$ at the asymptotic level $\alpha \in (0, 1/2)$ can thus be achieved by using the critical region $\{\boldsymbol{W}_n^{(k)} > \chi^2_{1-2\alpha}(1)\}$.

5. Estimating conditional correlation matrices

Having estimated the individual conditional variances of a vector $(\boldsymbol{\epsilon}_t)$ satisfying (2.4) in a first step, it is generally of interest to estimate the complete conditional variance matrix $\boldsymbol{H}_t$, which thus reduces to estimating the conditional correlation $\boldsymbol{R}_t$.

Suppose that matrix $\boldsymbol{R}_t$ is parameterized by some parameter $\boldsymbol{\rho}_0 \in \mathbb{R}^r$, together with the volatility parameter $\boldsymbol{\theta}_0$, as

$$\boldsymbol{R}_t = \boldsymbol{R}(\boldsymbol{\epsilon}_{t-1}, \boldsymbol{\epsilon}_{t-2}, \dots; \boldsymbol{\theta}_0, \boldsymbol{\rho}_0) = \boldsymbol{\mathcal{R}}(\boldsymbol{\eta}^*_{t-1}, \boldsymbol{\eta}^*_{t-2}, \dots; \boldsymbol{\rho}_0).$$

Let $\Lambda \subset \mathbb{R}^r$ denote a parameter set such that $\boldsymbol{\rho}_0 \in \Lambda$. If the $\boldsymbol{\eta}_t^*$ were observed, in view of (2.2) a QMLE of $\boldsymbol{\rho}_0$ would be obtained as any measurable solution of

$$\arg\min_{\boldsymbol{\rho}\in\Lambda} n^{-1} \sum_{t=1}^{n} \boldsymbol{\eta}_t^{*'} \tilde{\boldsymbol{R}}_t^{-1} \boldsymbol{\eta}_t^* + \log|\tilde{\boldsymbol{R}}_t|,$$

where, introducing initial values $\tilde{\boldsymbol{\eta}}_i^*$ for $i \le 0$, $\tilde{\boldsymbol{R}}_t = \boldsymbol{\mathcal{R}}(\boldsymbol{\eta}^*_{t-1}, \boldsymbol{\eta}^*_{t-2}, \dots, \tilde{\boldsymbol{\eta}}^*_0, \tilde{\boldsymbol{\eta}}^*_{-1}, \dots; \boldsymbol{\rho})$.

We therefore consider the two-step estimation method of the parameters of Model (2.4).

(a) **First step**: EbE estimation of the volatility parameters $\boldsymbol{\theta}_0^{(k)}$ and extraction of the vectors of residuals $\hat{\boldsymbol{\eta}}_t^* = (\hat{\eta}^*_{1t}, \dots, \hat{\eta}^*_{mt})'$ where $\hat{\eta}^*_{kt} = \tilde{\sigma}_{kt}^{-1}(\hat{\boldsymbol{\theta}}^{(k)})\epsilon_{kt}$;

(b) **Second step**: QML estimation of the conditional correlation matrix $\boldsymbol{\rho}_0$ by EbE, as a solution of

$$\arg\min_{\boldsymbol{\rho}\in\Lambda} n^{-1} \sum_{t=1}^{n} \hat{\boldsymbol{\eta}}_t^{*'} \hat{\boldsymbol{R}}_t^{-1} \hat{\boldsymbol{\eta}}_t^* + \log|\hat{\boldsymbol{R}}_t|,$$

where $\hat{\boldsymbol{R}}_t = \boldsymbol{\mathcal{R}}(\hat{\boldsymbol{\eta}}^*_{t-1}, \hat{\boldsymbol{\eta}}^*_{t-2}, \dots, \hat{\boldsymbol{\eta}}^*_1, \tilde{\boldsymbol{\eta}}^*_0, \tilde{\boldsymbol{\eta}}^*_{-1}, \dots; \boldsymbol{\rho})$.

We will establish the asymptotic properties of this approach in the case where $\boldsymbol{R}_t$ is constant, that is for Model (2.8)-(2.9). The case of the classical CCC-GARCH(p,q) models will be considered in Section 6.1.1.

5.1. Estimating general CCC models

Let $\boldsymbol{\rho} = (R_{21}, \ldots, R_{m1}, R_{32}, \ldots, R_{m2}, \ldots, R_{m,m-1})' = \mathrm{vech}^0(\boldsymbol{R})$, denoting by vech^0 the operator which stacks the sub-diagonal elements (excluding the diagonal) of a matrix. The global parameter is denoted

$$\boldsymbol{\vartheta} = (\boldsymbol{\theta}^{(1)'}, \ldots, \boldsymbol{\theta}^{(m)'}, \boldsymbol{\rho}')' := (\boldsymbol{\theta}', \boldsymbol{\rho}')' \in \mathbb{R}^d \times [-1,1]^{m(m-1)/2}, \quad d = \sum_{k=1}^m d_k,$$

and it belongs to the compact parameter set $\boldsymbol{\Theta} = \prod_{k=1}^m \boldsymbol{\Theta}_k \times [-1,1]^{m(m-1)/2}$. The second-step estimator of the constant correlation matrix $\boldsymbol{R}_t$ is given by $\hat{\boldsymbol{R}}_n = \frac{1}{n}\sum_{t=1}^n \hat{\boldsymbol{\eta}}_t^* \left(\hat{\boldsymbol{\eta}}_t^*\right)'$. Let $\hat{\boldsymbol{\vartheta}}_n = \left(\hat{\boldsymbol{\theta}}_n' := (\hat{\boldsymbol{\theta}}_n^{(1)'}, \ldots, \hat{\boldsymbol{\theta}}_n^{(m)'}), \hat{\boldsymbol{\rho}}_n'\right)'$, where $\hat{\boldsymbol{\rho}}_n = \mathrm{vech}^0(\hat{\boldsymbol{R}}_n)$.

THEOREM 5.1. *For the CCC model* (2.8)-(2.9)*, if* **A1**-**A6** *hold, then* $\hat{\boldsymbol{\vartheta}}_n \to \boldsymbol{\vartheta}_0$, *a.s. as* $n \to \infty$.

For the asymptotic normality, we introduce the following notations. Let the $d \times d$ matrix $\boldsymbol{J}^* = ((\kappa_{k\ell} - 1)\boldsymbol{J}_{k\ell})$ for $k, \ell = 1, \ldots, m$, and $\boldsymbol{J}_{k\ell} = E\left(\boldsymbol{d}_{kt}\boldsymbol{d}_{\ell t}'\right)$. Let, for $\boldsymbol{J}_0 = \mathrm{diag}(\boldsymbol{J}_{11}, \ldots, \boldsymbol{J}_{mm})$ in bloc-matrix notation,

$$\boldsymbol{\Sigma}_{\boldsymbol{\theta}} = \boldsymbol{J}_0^{-1}\boldsymbol{J}^*\boldsymbol{J}_0^{-1} = \left((\kappa_{k\ell} - 1)\boldsymbol{J}_{kk}^{-1}\boldsymbol{J}_{k\ell}\boldsymbol{J}_{\ell\ell}^{-1}\right).$$

Let also $\boldsymbol{d}_t = (\boldsymbol{d}_{1t}', \ldots, \boldsymbol{d}_{mt}')' \in \mathbb{R}^d$, $\boldsymbol{\Omega}_k = E\boldsymbol{d}_{kt}$ and $\boldsymbol{\Omega} = (\boldsymbol{\Omega}_1', \ldots, \boldsymbol{\Omega}_m')' \in \mathbb{R}^d$. Let $\boldsymbol{\Gamma} = \mathrm{var}\left(\mathrm{vech}^0\left\{\boldsymbol{\eta}_t^*\left(\boldsymbol{\eta}_t^*\right)'\right\}\right)$. For $\boldsymbol{x} \in \mathbb{R}^m$, let the $d \times d$ matrices $\boldsymbol{F}(\boldsymbol{x}) = \mathrm{diag}\{(1-x_1^2)\boldsymbol{j}_1, \ldots, (1-x_m^2)\boldsymbol{j}_m\}$, where $\boldsymbol{j}_k = (1, \ldots, 1) \in \mathbb{R}^{d_k}$, and $\boldsymbol{A}_{k\ell} = E\{\eta_{kt}^*\eta_{\ell t}^*\boldsymbol{F}(\boldsymbol{\eta}_t^*)\}$. Let, for $k, \ell = 2, \ldots, m$, the $d \times d$ matrix $\mathbf{M}_{k,\ell-1} = \mathrm{diag}\left(\mathbf{M}_{k,\ell-1}^{(1)}, \ldots, \mathbf{M}_{k,\ell-1}^{(m)}\right)$ where

$$\mathbf{M}_{k,\ell-1}^{(i)} = \begin{cases} \mathbf{0}_{d_i \times d_i} & \text{if } i \neq k \quad \text{ and } \quad i \neq \ell \\ R_{k,\ell-1}\boldsymbol{I}_{d_i} & \text{otherwise.} \end{cases}$$

Let the $d \times dm(m-1)/2$ matrices $\boldsymbol{A} = (\boldsymbol{A}_{21} \ldots \boldsymbol{A}_{m1} \;\; \boldsymbol{A}_{32} \ldots \boldsymbol{A}_{m,m-1})$ and $\boldsymbol{M} = (\boldsymbol{M}_{21} \ldots \boldsymbol{M}_{m1} \;\; \boldsymbol{M}_{32} \ldots \boldsymbol{M}_{m,m-1})$. Let the $d \times m(m-1)/2$ matrices $\boldsymbol{L} = \boldsymbol{A}(\boldsymbol{I}_{m(m-1)/2} \otimes \boldsymbol{\Omega})$, $\boldsymbol{\Lambda} = \boldsymbol{M}(\boldsymbol{I}_{m(m-1)/2} \otimes \boldsymbol{\Omega})$. Let $\boldsymbol{\Sigma}_{\boldsymbol{\theta\rho}} = -\frac{1}{2}\boldsymbol{\Sigma}_{\boldsymbol{\theta}}\boldsymbol{\Lambda} - \boldsymbol{J}_0^{-1}\boldsymbol{L}$, $\boldsymbol{\Sigma}_{\boldsymbol{\rho}} = \frac{1}{4}\boldsymbol{\Lambda}'\boldsymbol{\Sigma}_{\boldsymbol{\theta}}\boldsymbol{\Lambda} + \frac{1}{2}\left(\boldsymbol{\Lambda}'\boldsymbol{J}_0^{-1}\boldsymbol{L} + \boldsymbol{L}'\boldsymbol{J}_0^{-1}\boldsymbol{\Lambda}\right) + \boldsymbol{\Gamma}$. We need an additional assumption.

A13: The distribution of $\mathrm{vech}(\boldsymbol{\eta}_t\boldsymbol{\eta}_t')$ is not supported on an hyperplane.

THEOREM 5.2. *For the CCC model* (2.8)-(2.9)*, if* **A1**-**A13** *hold, for* $k = 1, \ldots, m$*, and* $\boldsymbol{\rho}_0 \in (-1,1)^{m(m-1)/2}$*, then*

$$\begin{pmatrix} \sqrt{n}\left(\hat{\boldsymbol{\theta}}_n - \boldsymbol{\theta}_0\right) \\ \sqrt{n}(\hat{\boldsymbol{\rho}}_n - \boldsymbol{\rho}_0) \end{pmatrix} \overset{\mathcal{L}}{\to} \mathcal{N}\left\{0, \boldsymbol{\Sigma} := \begin{pmatrix} \boldsymbol{\Sigma}_{\boldsymbol{\theta}} & \boldsymbol{\Sigma}_{\boldsymbol{\theta\rho}} \\ \boldsymbol{\Sigma}_{\boldsymbol{\theta\rho}}' & \boldsymbol{\Sigma}_{\boldsymbol{\rho}} \end{pmatrix}\right\},$$

and $\mathbf{\Sigma}$ *is a non-singular matrix.*

Remark 5.1. Even though the components of $\boldsymbol{\theta}_0$ are estimated independently, the components $\hat{\boldsymbol{\theta}}_n^{(k)}$ of $\hat{\boldsymbol{\theta}}_n$ are asymptotically non independent in general. More precisely, it can be seen that $\mathbf{\Sigma}_{\boldsymbol{\theta}}$ is bloc diagonal if $\mathrm{Cov}(\eta_{kt}^{*2}, \eta_{\ell t}^{*2}) = 0$ for any $k \neq \ell$.

Remark 5.2. In the asymptotic variance $\mathbf{\Sigma}_{\boldsymbol{\rho}}$ of $\hat{\boldsymbol{\rho}}_n$, the first two matrices in the sum reflect the effect of the estimation of $\boldsymbol{\theta}_0$, while the remaining matrix, $\mathbf{\Gamma}$, is independent of $\boldsymbol{\theta}_0$. A limit case is when the components of $\boldsymbol{\eta}_t^*$ are serially independent, that is when $\boldsymbol{\eta}_t^* = \boldsymbol{\eta}_t$ and $\boldsymbol{R}$ is the identity matrix. Then, straightforward computation shows that $\boldsymbol{L} = \mathbf{\Lambda} = \mathbf{0}$ and thus, in bloc-matrix notation,

$$\mathbf{\Sigma} = \begin{pmatrix} \mathbf{\Sigma}_{\boldsymbol{\theta}} & \mathbf{0} \\ \mathbf{0} & \boldsymbol{I}_{m(m-1)/2} \end{pmatrix} \quad \text{and} \quad \mathbf{\Sigma}_{\boldsymbol{\theta}} = \mathrm{diag}((\kappa_{11} - 1)\boldsymbol{J}_{11}^{-1}, \dots, (\kappa_{mm} - 1)\boldsymbol{J}_{mm}^{-1}).$$

Remark 5.3. It is worthnoting that all the matrices involved in the asymptotic covariance matrix $\mathbf{\Sigma}$ take the form of expectations. A simple estimator of $\mathbf{\Sigma}$ is thus obtained by replacing those expectations by their sample counterparts. For instance, it can be shown that a consistent estimator of $\boldsymbol{A}_{k\ell}$ is $\hat{\boldsymbol{A}}_{k\ell} = \frac{1}{n}\sum_{t=1}^n \hat{\eta}_{kt}^* \hat{\eta}_{\ell t}^* F(\hat{\boldsymbol{\eta}}_t^*)$.

Remark 5.4. In financial applications, the different returns are generally not available over the same time horizons. Discarding dates for which at least one return is not available may entail a severe sample size reduction. Instead, the correlations can be estimated by considering the returns pairwise (with different sample lengths for different pairs). Such estimators of the correlations are consistent, even if the estimated global correlation matrix may not be positive definite. This approach will be used in the empirical section.

5.2. Estimating DCC models

The asymptotic properties of the first-step EbEE were established in Theorem 4.1 for diagonal first-order DCC models. The second step can be applied for estimating $\boldsymbol{\rho} = (\alpha, \beta, (\mathrm{vech}^0(\boldsymbol{S}))')'$ in Model (4.2). The matrices $\tilde{\boldsymbol{R}}_t$ involved in the second step are obtained as $\tilde{\boldsymbol{R}}_t = \tilde{\boldsymbol{Q}}_t^{*-1/2}(\boldsymbol{\rho})\tilde{\boldsymbol{Q}}_t(\boldsymbol{\rho})\tilde{\boldsymbol{Q}}_t^{*-1/2}(\boldsymbol{\rho})$, where the $\tilde{\boldsymbol{Q}}_t(\boldsymbol{\rho})$ are computed recursively as

$$\tilde{\boldsymbol{Q}}_t(\boldsymbol{\rho}) = (1 - \alpha - \beta)\boldsymbol{S} + \alpha\tilde{\boldsymbol{Q}}_{t-1}^{*1/2}(\boldsymbol{\rho})\hat{\boldsymbol{\eta}}_{t-1}^*\hat{\boldsymbol{\eta}}_{t-1}^{*'}\tilde{\boldsymbol{Q}}_{t-1}^{*1/2}(\boldsymbol{\rho}) + \beta\tilde{\boldsymbol{Q}}_{t-1}(\boldsymbol{\rho}), \quad t \geq 1,$$

with initial value $\tilde{\boldsymbol{Q}}_0(\boldsymbol{\rho}) = \boldsymbol{S}$. The asymptotic properties of the second-step EbEE are an open issue.

6. Numerical Illustrations

The first part of the section presents a selection of Monte-Carlo experiments aiming at studying the performance of the EbE approach in finite sample. Real data examples will be presented in the second part.[8]

6.1. Monte-Carlo study

We will first illustrate the gains in computation time brought by the two-step EbE approach, by comparison with the usual Full QML (FQML) in which all the parameters are estimated in one step. We will also investigate, for CCC and DCC models, whether the gains in numerical complexity have a price in terms of finite-sample accuracy.

6.1.1. *Time complexity and accuracy comparison of the EbEE and the full QMLE*

Let us compare the computation cost of the EbEE with that of the FQMLE in the case of a diagonal CCC-GARCH$(1,1)$ model of dimension m, that is, under the specification (2.7) with $p=q=1$ and diagonal matrices $\boldsymbol{A}_1$ and $\boldsymbol{B}_1$. EbEE of all the model parameters requires m estimations of univariate GARCH-type models with 3 parameters, plus the computation of the empirical correlation of the EbE residuals. The full QMLE requires the optimization of a function of $3m+m(m-1)/2$ parameters. Because the time complexity of an optimization generally grows rapidly with the dimension of the objective function, the full QMLE should be much more costly than the EbEE in terms of computation time. The two estimators were fitted on simulations of length $n=2000$ of the CCC-GARCH$(1,1)$ model (2.7) with $\boldsymbol{A}_1=0.05\boldsymbol{I}_m$ and $\boldsymbol{B}_1=0.9\boldsymbol{I}_m$ (such values are close to those generally fitted to real series). The correlation matrix used for the simulations is $\boldsymbol{R}=\boldsymbol{I}_m$, but the $m(m-1)/2$ subdiagonal terms of $\boldsymbol{R}$ were estimated, together with the $3m$ other parameters of the model. The distribution of $\boldsymbol{\eta}_t$ is Gaussian, which has little impact on the computation times, but should give an advantage to the FQMLE (which is then the MLE) in terms of accuracy. Table 1 compares the effective computation times required by the two estimators as a function of the dimension m. As expected, the comparison of the CPU's is clearly in favor of the EbEE. Note that these computation times have been obtained using a single processor. Since the EbEE is clearly easily parallelizable (using one processor for each of the m optimizations), the advantage of the EbEE should be even more pronounced with a

[8] The code and data used in the paper are available on the web site `http://perso.univ-lille3.fr/~cfrancq/Christian-Francq/EbEE.html`

Table 1. Computation time of the two estimators (CPU time in seconds) and Relative Efficiency (RE) of the EbEE with respect to the FQMLE (NA means "Not Available" because of the impossibility to compute the FQMLE) for m-dimensional CCC-GARCH(1,1) models.

Dim. m	2	3	4	5	6	7	8	9
Nb. of param.	7	12	18	25	33	42	52	63
CPU for EbEE	0.57	0.88	1.18	1.31	1.52	1.85	2.04	2.37
CPU for FQMLE	32.49	100.78	123.33	215.38	317.85	617.33	876.52	1113.68
ratio of CPU	57.00	114.52	104.52	164.41	209.11	333.69	429.67	469.91
RE	0.96	1.00	0.99	0.97	0.99	0.99	0.97	1.00
Dim. m	10	11	12	50	100	200	400	800
Nb. of param.	75	88	102	1375	5250	20500	81000	322000
CPU for EbEE	2.82	2.98	3.49	13.67	27.89	56.58	110.00	226.32
CPU for FQMLE	1292.34	1520.60	1986.38	NA	NA	NA	NA	NA
ratio of CPU	458.28	510.27	569.16	NA	NA	NA	NA	NA
RE	102.42	304.36	14.22	NA	NA	NA	NA	NA

multiprocessing implementation. Table 1 also compares the relative efficiencies of the EbEE with respect to the FQMLE. To this aim, we first computed the approximated information matrix $\boldsymbol{J}_n = -\frac{1}{2n}\frac{\partial^2}{\partial\boldsymbol{\vartheta}\partial\boldsymbol{\vartheta}'}\sum_{t=1}^n \boldsymbol{\epsilon}_t'\boldsymbol{H}_t^{-1}\boldsymbol{\epsilon}_t + \log|\boldsymbol{H}_t|$. Note that when $(\boldsymbol{\eta}_t)$ is Gaussian and when $\hat{\boldsymbol{\vartheta}}_{ML}$ is the (Q)MLE, then $n(\hat{\boldsymbol{\vartheta}}_{ML} - \boldsymbol{\vartheta}_0)'\boldsymbol{J}_n(\hat{\boldsymbol{\vartheta}}_{ML} - \boldsymbol{\vartheta}_0)$ follows asymptotically a χ^2 distribution. More generally, the quadratic form $n(\hat{\boldsymbol{\vartheta}}_n - \boldsymbol{\vartheta}_0)'\boldsymbol{J}_n(\hat{\boldsymbol{\vartheta}}_n - \boldsymbol{\vartheta}_0)$ can serve as a measure of accuracy of an estimator $\hat{\boldsymbol{\vartheta}}_n$ (the Euclidean distance, obtained by replacing $\boldsymbol{J}_n$ by the identity matrix, has the drawback of being scale dependent). The relative efficiency (RE) displayed in Table 1 is equal to

$$RE = \frac{(\hat{\boldsymbol{\vartheta}}_{QMLE} - \boldsymbol{\vartheta}_0)'\boldsymbol{J}_n(\hat{\boldsymbol{\vartheta}}_{QMLE} - \boldsymbol{\vartheta}_0)}{(\hat{\boldsymbol{\vartheta}}_{EbEE} - \boldsymbol{\vartheta}_0)'\boldsymbol{J}_n(\hat{\boldsymbol{\vartheta}}_{EbEE} - \boldsymbol{\vartheta}_0)}$$

where $\hat{\boldsymbol{\vartheta}}_{EbEE}$ and $\hat{\boldsymbol{\vartheta}}_{QMLE}$ denote respectively the EbEE and FQMLE. Because the computation time of the FQMLE is enormous when m is large, the RE and CPU times are only computed on 1 simulation, but they are typical of what is generally observed. When $m \leq 9$, the accuracies are very similar, with a slight advantage to the FQMLE (which corresponds here to the MLE). When the number of parameters becomes too large ($m > 9$) the computation time of the FQMLE becomes prohibitive, and more importantly the optimization fails to give a reasonable value of $\hat{\boldsymbol{\vartheta}}_{QMLE}$ (see the RE for $m \geq 10$).

6.1.2. *Estimating a DCC model by two-step EbE and by FQML*

We now compare the standard one-step FQMLE with the two step method described in Section 5 in the case of a bivariate cDCC-GARCH(1,1) model defined by (2.7) and (4.2),

Table 2. Comparison of the EbEE and QMLE over 100 replications for a bivariate DCC-GARCH(1,1) of length $n = 1000$ and $n = 5000$.

			$n = 1000$			$n = 5000$		
	true val.	estim.	Median	bias	RMSE	Median	bias	RMSE
$\boldsymbol{\omega}$	0.01	EbEE	0.014	0.037	0.134	0.012	0.002	0.005
		QMLE	0.017	0.116	0.239	0.013	0.096	0.212
	0.01	EbEE	0.014	0.040	0.159	0.011	0.002	0.005
		QMLE	0.019	0.104	0.229	0.013	0.076	0.183
$\boldsymbol{A}$	0.025	EbEE	0.020	-0.001	0.017	0.024	-0.001	0.006
		QMLE	0.025	0.037	0.105	0.027	0.043	0.104
	0.025	EbEE	0.026	0.005	0.028	0.027	0.002	0.007
		QMLE	0.032	0.046	0.109	0.028	0.043	0.105
	0.025	EbEE	0.031	0.011	0.023	0.026	0.001	0.006
		QMLE	0.034	0.048	0.108	0.029	0.044	0.104
	0.025	EbEE	0.024	0.000	0.019	0.024	-0.001	0.007
		QMLE	0.027	0.040	0.107	0.026	0.040	0.105
diag($\boldsymbol{B}$)	0.94	EbEE	0.932	-0.058	0.194	0.937	-0.003	0.012
		QMLE	0.926	-0.157	0.319	0.932	-0.131	0.279
	0.94	EbEE	0.934	-0.049	0.193	0.937	-0.003	0.011
		QMLE	0.925	-0.147	0.309	0.935	-0.118	0.263
$\boldsymbol{S}[1, 2]$	0.3	EbEE	0.308	-0.001	0.137	0.305	-0.001	0.065
		QMLE	0.336	0.024	0.222	0.314	0.063	0.205
α	0.04	EbEE	0.043	0.002	0.015	0.041	0.000	0.006
		QMLE	0.046	0.017	0.048	0.042	0.008	0.027
β	0.95	EbEE	0.936	-0.013	0.028	0.948	-0.002	0.008
		QMLE	0.931	-0.055	0.172	0.946	-0.010	0.034

with a full matrix $\boldsymbol{A} = \boldsymbol{A}_1$ and a diagonal matrix $\boldsymbol{B} = \boldsymbol{B}_1$. The value of $\boldsymbol{\vartheta}_0$ is given in the first column of Table 2, and $(\boldsymbol{\eta}_t)$ is an iid sequence distributed as a Student distribution with $\nu = 9$ degrees of freedom, standardized in such a way that $\mathrm{Var}(\boldsymbol{\eta}_t) = \boldsymbol{I}_2$. Note that our Monte Carlo experiment is restricted to a bivariate model because the computation time of the FQMLE is too demanding when $m > 2$.[9] Table 2 summarizes the distribution of the two estimators over 100 independent simulations of the model, for the lengths $n = 1000$ and $n = 5000$. The EbEE is remarkably more accurate than the FQMLE, whatever the parameter. A detailed examination of the estimations reveals that the FQMLE produces more outliers (such as for example $\hat{b} = 0$) than the EbEE, and the Root Mean Square Errors (RMSE) of estimation are much smaller for the EbEE than for the FQMLE. One difficulty encountered in the implementation of the two estimators is that the constraints $\rho(\boldsymbol{B}) < 1$ and $\beta < 1$ are not sufficient to ensure the non explosiveness of $\boldsymbol{Q}_t(\boldsymbol{\vartheta})$ as $t \to \infty$. The problem seems to be more severe for the FQMLE than for the EbEE, which may explain the surprisingly poor performance of the FQMLE compared to the EbEE.

[9]Results reported in the supplementary file illustrate the ability of the EbEE to estimate the individual volatilities of a cDCC for larger dimensions ($m > 2$).

6.2. Empirical examples

6.2.1. Dealing with missing or asynchronous data

One problem encountered in modelling multivariate financial series is that the different return components may not be available over the same time horizon. Another issue with financial returns is the lack of synchronicity. For daily returns, the time of measurement is typically the closing time, which can be very different for series across different markets entering in the construction of portfolios. Different techniques of synchronization have been proposed. For instance Audrino and Bühlmann (2004) developed a procedure for the CCC-GARCH(1,1) model. However, the need to choose an auxiliary model for predicting the missing observations may be found unsatisfactory.

The EbE procedure has interest for both issues, missing data and asynchronicity. First, the estimation of a given equation generally does not require observability of the whole returns over the entire period. This is in particular the case for diagonal models. Moreover, the estimation of the correlation matrix in CCC models can be achieved by considering the returns pairwise (see Remark 5.4). The missing data issue is illustrated in the supplementary file. Concerning asynchronicity, we propose the following illustration based on world stock market indices.

At the opening of the New York stock exchange, investors have knowledge of the closing price at the Tokyo stock exchange. It is thus possible to use e.g. the squared return of the Nikkei 225 of day t (say Nik_t) to predict the squared return of the SP500 at the same date (say SP_t). Since Nik_t conveys more recent information than SP_{t-1}, it is reasonable to think that it may appear significantly in the volatility of the SP500 at time t. Modeling the individual volatilities by augmented GARCH models is a convenient way to tackle the problem. For simplicity, we considered only four indices: the SP500 (closing price at around 21 GMT), the CAC and FTSE (closing price at 16:30) and the Nikkei (closing price at 6). As a function of the most recent available returns and a feedback mechanism, the fitted individual volatilities using the period from 1990-01-01 to 2013-04-22 can be written, with obvious notations, as

$$\sigma^2_{\mathrm{SP}_t} = \underset{(0.008)}{0.039} + \underset{(0.013)}{0.064}\ \mathrm{SP}_{t-1} + \underset{(0.009)}{0.038}\ \mathrm{CAC}_t + \underset{(0.020)}{0.187}\ \mathrm{FTSE}_t + \underset{(0.003)}{0.000}\ \mathrm{Nik}_t + \underset{(0.024)}{0.660}\sigma^2_{\mathrm{SP}_{t-1}}$$

$$\sigma^2_{\mathrm{CAC}_t} = \underset{(0.010)}{0.042} + \underset{(0.014)}{0.050}\ \mathrm{SP}_{t-1} + \underset{(0.012)}{0.064}\ \mathrm{CAC}_{t-1} + \underset{(0.018)}{0.036}\ \mathrm{FTSE}_{t-1} + \underset{(0.004)}{0.015}\ \mathrm{Nik}_t + \underset{(0.018)}{0.844}\sigma^2_{\mathrm{CAC}_{t-1}}$$

$$\sigma^2_{\mathrm{FTSE}_t} = \underset{(0.004)}{0.013} + \underset{(0.007)}{0.039}\ \mathrm{SP}_{t-1} + \underset{(0.004)}{0.000}\ \mathrm{CAC}_{t-1} + \underset{(0.0010)}{0.071}\ \mathrm{FTSE}_{t-1} + \underset{(0.002)}{0.006}\ \mathrm{Nik}_t + \underset{(0.013)}{0.869}\sigma^2_{\mathrm{CAC}_{t-1}}$$

$$\sigma^2_{\mathrm{Nik}_t} = \underset{(0.015)}{0.068} + \underset{(0.016)}{0.055}\ \mathrm{SP}_{t-1} + \underset{(0.011)}{0.006}\ \mathrm{CAC}_{t-1} + \underset{(0.019)}{0.010}\ \mathrm{FTSE}_{t-1} + \underset{(0.014)}{0.108}\ \mathrm{Nik}_{t-1} + \underset{(0.019)}{0.826}\sigma^2_{\mathrm{CAC}_{t-1}}$$

where the estimated standard deviations, obtained from Theorem 3.1, are given into brack-

Table 3. For each pair of exchange rates: p-values of the tests of the null hypotheses $H_0^{(1)}$ and $H_0^{(2)}$ implied by the bivariate BEKK-GARCH(1,1) model. Gray cells contain p-values less than 2.5%.

	CAD		CHF		CNY		GBP		JPY	
	$H_0^{(1)}$	$H_0^{(2)}$	$H_0^{(1)}$	$H_0^{(2)}$	$H_0^{(1)}$	$H_0^{(2)}$	$H_0^{(1)}$	$H_0^{(2)}$	$H_0^{(1)}$	$H_0^{(2)}$
CHF	0.000	0.163								
CNY	0.120	0.015	0.122	0.500						
GBP	0.012	0.023	0.128	0.000	0.005	0.100				
JPY	0.007	0.006	0.500	0.500	0.500	0.087	0.050	0.000		
USD	0.500	0.021	0.114	0.000	0.500	0.381	0.068	0.000	0.102	0.000

ets. It is seen that, for instance, the FTSE at time t has strong influence on the volatility of the SP500 at the same date (but a few hours later). Thus, by taking into account the availability of the most recent observations the model reveals spillover effects between series.

6.2.2. Bivariate BEKK for exchange rates?

We considered returns series of six daily exchange rates with respect to the Euro, from January 14, 2000 to May 16, 2013. We tested the adequacy of bivariate BEKK models, using Proposition 4.1. For each pair of exchange rates, we estimated Model (4.5) and we tested the restrictions $H_0^{(1)}$ and $H_0^{(2)}$ that are satisfied when the DGP is the BEKK-GARCH(1,1) model (4.3). Table 3 shows that, for 12 bivariate series over 15, either $H_0^{(1)}$ or $H_0^{(2)}$ is clearly rejected, which invalidates the adequacy of the bivariate BEKK model for the 12 pairs. Using the Bonferroni correction, one can indeed reject the model at significant level less than α if one of the two hypothesis $H_0^{(k)}$ is rejected at the level $\alpha/2$. This does not mean that a global BEKK model would be rejected for the vector of 6 series. An extension of Proposition 4.1 for larger m would allow to perform a test but such an extension is left for future research.

7. Conclusion

EbE estimation of MGARCH models is a standard method used in applied works to alleviate the computational burden implied by large cross-sectional dimensions. In this study, we established asymptotic properties of the EbEE of the individual conditional variances, under general assumptions on their parameterization. Unexpectedly, we found that such EbE estimators may be superior to the QMLE in terms of asymptotic accuracy. Our framework covers the most widely used MGARCH models in financial applications. For semi-diagonal BEKK models and DCC models, the asymptotic results were shown to hold under explicit conditions. In the former case, we explained how to test the constraints implied by the

BEKK specification. For CCC models (including the standard CCC-GARCH(p, q) model) we proved the consistency and the joint asymptotic normality of the EbE volatility and correlation matrix estimators.

The main motivation for using an EbE approach in applications is the important gains in computation time, and our simulation experiments confirmed that such gains can be huge. For moderate dimensions the global QML estimator can even be unfeasible, while we did not encounter such difficulties with the EbE approach. Our experiments revealed that the EbE estimator may be superior to the QMLE in terms of accuracy, not only for the volatility parameters but also for the parameters of a DCC specification of the conditional correlation. For real series, the separate estimation of the volatilities allows to handle, without discarding too many data, series that are not available at the same date, or at the same hour for daily returns.

Appendix

A. Technical assumptions

A2: for any real sequence $(e_i)_{i\geq 1}$, the function $\boldsymbol{\theta}^{(k)} \mapsto \sigma_k(e_1, e_2, \dots; \boldsymbol{\theta}^{(k)})$ is continuous and there exists a measurable function $K : \mathbb{R}^\infty \mapsto (0, \infty)$ such that

$$|\sigma_k(e_1, e_2, \dots; \boldsymbol{\theta}^{(k)}) - \sigma_k(e_1, e_2, \dots; \boldsymbol{\theta}_0^{(k)})| \leq K(e_1, \dots)\|\boldsymbol{\theta}^{(k)} - \boldsymbol{\theta}_0^{(k)}\|,$$

and $E\left(\frac{K(\boldsymbol{\epsilon}_{t-1}, \boldsymbol{\epsilon}_{t-2}, \dots)}{\sigma_{kt}(\boldsymbol{\theta}_0^{(k)})}\right)^2 < \infty.$

A3: For some neighborhood $\mathcal{V}(\boldsymbol{\theta}_0^{(k)})$ of $\boldsymbol{\theta}_0^{(k)}$ we have $E\sup_{\boldsymbol{\theta}^{(k)} \in \mathcal{V}(\boldsymbol{\theta}_0^{(k)})}\left(\frac{\sigma_{kt}(\boldsymbol{\theta}_0^{(k)})}{\sigma_{kt}(\boldsymbol{\theta}^{(k)})}\right)^2 < \infty.$

A5: we have $\sigma_{kt}(\boldsymbol{\theta}_0^{(k)}) = \sigma_{kt}(\boldsymbol{\theta}^{(k)})$ $a.s.$ iff $\boldsymbol{\theta}^{(k)} = \boldsymbol{\theta}_0^{(k)}$.

Let $\Delta_{kt}(\boldsymbol{\theta}^{(k)}) = \tilde{\sigma}_{kt}(\boldsymbol{\theta}^{(k)}) - \sigma_{kt}(\boldsymbol{\theta}^{(k)})$, $a_t = \sup_k \sup_{\boldsymbol{\theta}^{(k)} \in \boldsymbol{\Theta}^{(k)}} |\Delta_{kt}(\boldsymbol{\theta}^{(k)})|$. Let C and ρ be generic constants with $C > 0$ and $0 < \rho < 1$. The "constant" C is allowed to depend on variables anterior to $t = 0$.

A6: We have $a_t \leq C\rho^t$, a.s.

A9: for any real sequence $(e_i)_{i\geq 1}$, the function $\boldsymbol{\theta}^{(k)} \mapsto \sigma_k(e_1, e_2, \dots; \boldsymbol{\theta}^{(k)})$ has continuous second-order derivatives.

A10: there exists a neighborhood $\mathcal{V}(\boldsymbol{\theta}_0^{(k)})$ of $\boldsymbol{\theta}_0^{(k)}$ such that

$$\sup_{\boldsymbol{\theta}^{(k)}\in\mathcal{V}(\boldsymbol{\theta}_0^{(k)})}\left\|\frac{1}{\sigma_{kt}(\boldsymbol{\theta}^{(k)})}\frac{\partial\sigma_{kt}(\boldsymbol{\theta}^{(k)})}{\partial\boldsymbol{\theta}^{(k)}}\right\|^{4(1+\frac{1}{\delta})}, \qquad \sup_{\boldsymbol{\theta}^{(k)}\in\mathcal{V}(\boldsymbol{\theta}_0^{(k)})}\left\|\frac{1}{\sigma_{kt}(\boldsymbol{\theta}^{(k)})}\frac{\partial^2\sigma_{kt}(\boldsymbol{\theta}^{(k)})}{\partial\boldsymbol{\theta}^{(k)}\partial\boldsymbol{\theta}^{(k)\prime}}\right\|^{2(1+\frac{1}{\delta})},$$

$$\sup_{\boldsymbol{\theta}^{(k)}\in\mathcal{V}(\boldsymbol{\theta}_0^{(k)})}\left|\frac{\sigma_{kt}(\boldsymbol{\theta}_0^{(k)})}{\sigma_{kt}(\boldsymbol{\theta}^{(k)})}\right|^4, \qquad \text{have finite expectations.}$$

A11: We have $b_t := \sup_k \sup_{\boldsymbol{\theta}^{(k)}\in\mathcal{V}(\boldsymbol{\theta}_0^{(k)})}\left\|\frac{\partial\Delta_{kt}(\boldsymbol{\theta}^{(k)})}{\partial\boldsymbol{\theta}^{(k)}}\right\| \leq C\rho^t, \quad a.s.$

A12: For $k = 1,\dots,m$ and for any $\boldsymbol{x}\in\mathbb{R}^{d_k}$, we have: $\quad \boldsymbol{x}'\frac{\partial\sigma^2_{kt}(\boldsymbol{\theta}_0^{(k)})}{\partial\boldsymbol{\theta}^{(k)}} = 0,\ a.s. \quad \Rightarrow \quad \boldsymbol{x} = 0.$

A10*: there exists a neighborhood $\mathcal{V}(\boldsymbol{\theta}_0^{(k)})$ of $\boldsymbol{\theta}_0^{(k)}$ such that

$$\sup_{\boldsymbol{\theta}^{(k)}\in\mathcal{V}(\boldsymbol{\theta}_0^{(k)})}\left\|\frac{1}{\sigma_{kt}(\boldsymbol{\theta}^{(k)})}\frac{\partial\sigma_{kt}(\boldsymbol{\theta}^{(k)})}{\partial\boldsymbol{\theta}^{(k)}}\right\|^{4}, \qquad \sup_{\boldsymbol{\theta}^{(k)}\in\mathcal{V}(\boldsymbol{\theta}_0^{(k)})}\left\|\frac{1}{\sigma_{kt}(\boldsymbol{\theta}^{(k)})}\frac{\partial^2\sigma_{kt}(\boldsymbol{\theta}^{(k)})}{\partial\boldsymbol{\theta}^{(k)}\partial\boldsymbol{\theta}^{(k)\prime}}\right\|^{2},$$

$$\sup_{\boldsymbol{\theta}^{(k)}\in\mathcal{V}(\boldsymbol{\theta}_0^{(k)})}\left|\frac{\sigma_{kt}(\boldsymbol{\theta}_0^{(k)})}{\sigma_{kt}(\boldsymbol{\theta}^{(k)})}\right|^4, \qquad \text{have finite expectations.}$$

B. Proofs

To save space, the proofs of Theorems 3.1, 3.2, 3.3 and 5.1, Propositions 3.1, 3.2 and 4.1 are displayed in the supplementary file.

B.1. Proof of Theorem 4.1

The proof consists in verifying the conditions required in Theorem 3.1. By Boussama, Fuchs and Stelzer (2011) and Aielli (2013), the conditions of the theorem ensure the existence of a strictly stationary, non anticipative and ergodic solution $[\mathrm{vech}(\boldsymbol{R}_t)', \boldsymbol{\eta}_t^{*'}]'$, that is a measurable function of $\{\boldsymbol{\eta}_{t-u}, u \geq 0\}$, to the equations $\boldsymbol{\eta}_t^* = \boldsymbol{R}_t^{1/2}\boldsymbol{\eta}_t$ and (4.2). Now, in view of $\epsilon_{\ell t} = \sigma_{\ell t}\eta^*_{\ell t}$, we have

$$\sigma^2_{\ell t} = \omega_\ell + a_\ell(\eta^*_{\ell,t-1})\sigma^2_{\ell,t-1} = \omega_\ell\left(1 + \sum_{s=1}^{\infty} a_\ell(\eta^*_{\ell,t-1})\dots a_\ell(\eta^*_{\ell,t-s})\right), \quad a.s.$$

where $a_\ell(x) = \alpha_\ell x^2 + \beta_\ell$. The a.s. convergence follows from the Cauchy rule for positive terms series, which can be applied because $E\log a_\ell(\eta^*_{\ell,t}) \leq \log Ea_\ell(\eta^*_{\ell,t}) = \log(\alpha_\ell + \beta_\ell) < 0$. It follows that $\boldsymbol{\epsilon}_t$ is a measurable function of $\{\boldsymbol{\eta}_{t-u}, u \geq 0\}$. Using the second equality in

(2.3), we have by the law of iterated conditional expectations $Ea_\ell(\eta^*_{\ell,t-1})\dots a_\ell(\eta^*_{\ell,t-s}) = (\alpha_\ell+\beta_\ell)^s$, from which we deduce that $E\sigma^2_{\ell t}<\infty$. Thus **A1** is satisfied. We note that

$$\eta^{*2}_{k,t} \le (\boldsymbol{\eta}^*_t)'\boldsymbol{\eta}^*_t = (\boldsymbol{\eta}_t)'\boldsymbol{R}_t\boldsymbol{\eta}_t \le \sum_{\ell,\ell'}|\eta_{\ell,t}\eta_{\ell',t}| = \left(\sum_{\ell,\ell'}|\eta_{\ell,t}|\right)^2,$$

from which **A8** follows, using $E\left\|\boldsymbol{\eta}_t\right\|^{4(1+\delta)}<\infty$. The other assumptions required to apply Theorem 3.1 can be shown to hold as in standard GARCH(1,1) models (see for instance Francq and Zakoïan (2004)).

B.2. Proof of Theorem 4.2

The existence of a (unique) ergodic, non anticipative, strictly and second-order stationary solution $(\boldsymbol{\epsilon}_t)$ of Model (4.3), under the conditions given in the corollary, follows from Boussama, Fuchs and Stelzer (2011), Theorem 2.4. Thus **A1** holds with $s=2$. By Proposition 4.5 of the same article, if the spectral radius of $\boldsymbol{A}_0+\boldsymbol{B}_0$ is less than 1, the spectral radius of $\sum_{j=1}^p \boldsymbol{H}_m(\boldsymbol{B}_{0j}\otimes\boldsymbol{B}_{0j})\boldsymbol{K}'_m$ is also less than 1. The latter matrix being diagonal, it can be seen that this entails $\sum_{j=1}^p b^2_{jk}<1$. Thus, under the strict stationarity condition, it is always possible to choose the compact set $\boldsymbol{\Theta}_k$ so that it contains the true parameter value. Assumption **A4** is satisfied by definition of $\boldsymbol{\Theta}^{(k)}$.

Now we turn to **A5**. Suppose $\sigma_t(\boldsymbol{\theta}^{(k)}_0)=\sigma_t(\boldsymbol{\theta}^{(k)})$, a.s. The polynomial $\mathcal{B}_{\boldsymbol{\theta}^{(k)}}(L)$ being invertible for any $\boldsymbol{\theta}^{(k)}\in\boldsymbol{\Theta}^{(k)}$, we have

$$\begin{aligned}
&\mathcal{B}^{-1}_{\boldsymbol{\theta}^{(k)}_0}(L)\sum_{i=1}^q\left(\sum_{\ell=1}^m \theta^{(k)}_{0,\ell+1+m(i-1)}\epsilon_{\ell,t-i}\right)^2 - \mathcal{B}^{-1}_{\boldsymbol{\theta}^{(k)}}(L)\sum_{i=1}^q\left(\sum_{\ell=1}^m \theta^{(k)}_{\ell+1+m(i-1)}\epsilon_{\ell,t-i}\right)^2\\
=\;& \mathcal{B}^{-1}_{\boldsymbol{\theta}^{(k)}}(1)\theta^{(k)}_1 - \mathcal{B}^{-1}_{\boldsymbol{\theta}^{(k)}_0}(1)\theta^{(k)}_{01}, \qquad a.s.
\end{aligned}$$

Then there exists some variables $c_{t-2}, a_{\ell,\ell',t-2}, \ell,\ell' = 1,\dots m$ belonging to the past of $\boldsymbol{\eta}_{t-1}$ such that $c_{t-2}+\sum_{\ell,\ell'=1}^m a_{\ell,\ell',t-2}\eta_{\ell,t-1}\eta_{\ell',t-1}=0$. Therefore, if those variables are not all equal to zero, the distribution of $\boldsymbol{\eta}_t$ conditional to the past is degenerate. Since $\boldsymbol{\eta}_t$ is independent from the past, this means that the unconditional distribution of $\boldsymbol{\eta}_t$ is degenerate, in contradiction with the existence of a density around zero. Thus $c_{t-2}=a_{1,1,t-2}=\cdots=a_{m,m,t-2}=0$, from which we deduce, by iterating the same argument, that $\boldsymbol{\theta}^{(k)}=\boldsymbol{\theta}^{(k)}_0$. Therefore, **A5** is verified. We omit the proof of **A6**, which can be done following the lines of proof of (4.6) in Francq and Zakoïan (2004). Hence the proof of the consistency of $\hat{\boldsymbol{\theta}}^{(k)}_n$ is complete.

Now we turn to the asymptotic normality. Assumption **A7** holds by assumption, and **A8** can be verified by the arguments used in the proof of Theorem 4.1. **A9** is obviously satisfied. The proofs of **A10-A11** being similar to those made for standard GARCH models in Francq and Zakoïan (2004), they will be omitted. To establish **A12**, let $\boldsymbol{x} \in \mathbb{R}^{1+p+qm}$ such that $\boldsymbol{x}'\frac{\partial\sigma^2_{kt}(\boldsymbol{\theta}_0^{(k)})}{\partial\boldsymbol{\theta}^{(k)}} = 0,\ a.s.$ It follows that

$$x_1 + \sum_{i=1}^{q} 2\left(\sum_{\ell=1}^{m} a_{ik\ell}\epsilon_{\ell,t-i}\right)\boldsymbol{x}'\begin{pmatrix}\mathbf{0}_{1+m(i-1)\times 1}\\ \boldsymbol{\epsilon}_{t-i}\\ \mathbf{0}_{p+m(q-i)\times 1}\end{pmatrix}$$
$$+\sum_{j=1}^{p} x_{1+qm+j}h_{kk,t-j} + \sum_{j=1}^{p} b_{jk}^2\boldsymbol{x}'\frac{\partial\sigma^2_{k,t-j}(\boldsymbol{\theta}_0^{(k)})}{\partial\boldsymbol{\theta}^{(k)}} = 0, \qquad a.s.$$

Note that the latter sum is equal to zero by stationarity. We thus have $x_1 + 2\left(\sum_{\ell=1}^{m} a_{1k\ell}\epsilon_{\ell,t-1}\right)\left(\sum_{\ell=1}^{m} x_{1+\ell}\epsilon_{\ell,t-1}\right) + z_{t-2} = 0, \qquad a.s.$ where z_{t-2} is a variable belonging to the past of $\boldsymbol{\epsilon}_{t-1}$. The arguments given for the proof of **A5** allow to conclude that $x_1 = \cdots = x_{m+1} = 0$. By iterating the argument we get $\boldsymbol{x} = 0$. The asymptotic normality follows from Theorem 3.1.

B.3. *Proof of Corollary 4.1*

The result is a straightforward consequence of the consistency of $\hat{\boldsymbol{A}}$ and $\hat{\boldsymbol{B}}$, obtained in Theorem 4.2, and the fact that Assumption **A1** with $s = 2$ is satisfied.

B.4. *Proof of Theorem 5.2*

Let $\dot{\boldsymbol{\ell}}_t(\boldsymbol{\theta}) = \left(\frac{\partial}{\partial\boldsymbol{\theta}^{(1)\prime}}\ell_{1t}(\boldsymbol{\theta}^{(1)}),\ldots,\frac{\partial}{\partial\boldsymbol{\theta}^{(m)\prime}}\ell_{mt}(\boldsymbol{\theta}^{(m)})\right)'$. For $\boldsymbol{\theta}^{(k)} \in \boldsymbol{\Theta}^{(k)}$ let $\tilde{\eta}^*_{kt}(\boldsymbol{\theta}^{(k)}) = \tilde{\sigma}^{-1}_{kt}(\boldsymbol{\theta}^{(k)})\epsilon_{kt}$ and $\eta^*_{kt}(\boldsymbol{\theta}^{(k)}) = \sigma^{-1}_{kt}(\boldsymbol{\theta}^{(k)})\epsilon_{kt}$. The proof relies on a set of preliminary results.

i) $E\left\|\dot{\boldsymbol{\ell}}_t(\boldsymbol{\theta}_0)\dot{\boldsymbol{\ell}}_t'(\boldsymbol{\theta}_0)\right\| < \infty$,

ii) $\sup_{\boldsymbol{\theta}\in\mathcal{V}(\boldsymbol{\theta}_0)}\left\|\frac{1}{n}\sum_{t=1}^{n}\frac{\partial\mathrm{vech}^0\left\{\boldsymbol{\eta}^*_t(\boldsymbol{\eta}^*_t)'\right\}}{\partial\boldsymbol{\theta}'} - \frac{\partial\mathrm{vech}^0\left\{\tilde{\boldsymbol{\eta}}^*_t(\tilde{\boldsymbol{\eta}}^*_t)'\right\}}{\partial\boldsymbol{\theta}'}\right\| \to 0$, in probability,

iii) $\frac{1}{n}\sum_{t=1}^{n}\left(\frac{\partial\mathrm{vech}^0\left\{\boldsymbol{\eta}^*_t(\boldsymbol{\eta}^*_t)'\right\}}{\partial\boldsymbol{\theta}'}\right)_{\boldsymbol{\theta}_n} \to -\frac{1}{2}\boldsymbol{\Lambda}'$, a.s. for any $\boldsymbol{\theta}_n$ between $\hat{\boldsymbol{\theta}}_n$ and $\boldsymbol{\theta}_0$,

iv) $\frac{1}{\sqrt{n}}\sum_{t=1}^{n}\begin{pmatrix}\dot{\boldsymbol{\ell}}_t(\boldsymbol{\theta}_0)\\ \mathrm{vech}^0\left\{\boldsymbol{\eta}^*_t(\boldsymbol{\eta}^*_t)' - \boldsymbol{R}\right\}\end{pmatrix} \xrightarrow{\mathcal{L}} \mathcal{N}\left(\mathbf{0},\begin{pmatrix}\boldsymbol{J}^* & \boldsymbol{L}\\ \boldsymbol{L}' & \boldsymbol{\Gamma}\end{pmatrix}\right)$, Point *i)* follows from the arguments given to prove i) in the proof of the asymptotic normality of $\hat{\boldsymbol{\theta}}_n$ (Theorem 3.1). Point ii) is equivalent to

$$\sup_{\boldsymbol{\theta}\in\mathcal{V}(\boldsymbol{\theta}_0)} \left\| \frac{1}{\sqrt{n}} \sum_{t=1}^{n} \frac{\partial\left(\eta^*_{kt}\eta^*_{\ell t}\right)}{\partial\boldsymbol{\theta}'}(\boldsymbol{\theta}) - \frac{\partial\left(\tilde{\eta}^*_{kt}\tilde{\eta}^*_{\ell t}\right)}{\partial\boldsymbol{\theta}'}(\boldsymbol{\theta}) \right\| \to 0\ , \quad \text{in probability.}$$

In view of

$$\frac{\partial}{\partial\boldsymbol{\theta}'}\left\{\eta^*_{kt}\eta^*_{\ell t}\right\}(\boldsymbol{\theta}) = -\frac{\epsilon_{kt}}{\sigma_{kt}(\boldsymbol{\theta}^{(k)})}\frac{\epsilon_{\ell t}}{\sigma_{\ell t}(\boldsymbol{\theta}^{(\ell)})}\left(\frac{1}{\sigma_{kt}}\frac{\partial\sigma_{kt}(\boldsymbol{\theta}^{(k)})}{\partial\boldsymbol{\theta}'} + \frac{1}{\sigma_{\ell t}}\frac{\partial\sigma_{\ell t}(\boldsymbol{\theta}^{(\ell)})}{\partial\boldsymbol{\theta}'}\right),$$

and the same equality for $\partial\left\{\tilde{\eta}^*_{kt}\tilde{\eta}^*_{\ell t}\right\}(\boldsymbol{\theta})/\partial\boldsymbol{\theta}'$, with σ_{kt} and $\sigma_{\ell t}$ replaced by $\tilde{\sigma}_{kt}$ and $\tilde{\sigma}_{\ell t}$, the conclusion follows by the arguments used to establish ii) in the proof of the asymptotic normality of $\hat{\boldsymbol{\theta}}_n$. Now we turn to iii). Note that

$$\frac{\partial}{\partial\boldsymbol{\theta}'}\left\{\eta^*_{kt}\eta^*_{\ell t}\right\}(\boldsymbol{\theta}_0) = -\eta^*_{kt}\eta^*_{\ell t}\left(\frac{1}{\sigma_{kt}}\frac{\partial\sigma_{kt}(\boldsymbol{\theta}_0^{(k)})}{\partial\boldsymbol{\theta}'} + \frac{1}{\sigma_{\ell t}}\frac{\partial\sigma_{\ell t}(\boldsymbol{\theta}_0^{(\ell)})}{\partial\boldsymbol{\theta}'}\right).$$

Thus, letting $d_{(k)} = \sum_{i=k+1}^{m} d_i$ and ${}_{(k)}d = \sum_{i=1}^{k-1} d_i$, with obvious conventions when $k=1$ or $k=m$,

$$E\left(\frac{\partial}{\partial\boldsymbol{\theta}'}\left\{\eta^*_{kt}\eta^*_{\ell t}\right\}(\boldsymbol{\theta}_0)\right) = -\frac{1}{2}R_{k\ell}[(\mathbf{0}_{1\times{}_{(k)}d}\ \boldsymbol{\Omega}'_k\ \mathbf{0}_{1\times d_{(k)}}) + (\mathbf{0}_{1\times{}_{(\ell)}d}\ \boldsymbol{\Omega}'_\ell\ \mathbf{0}_{1\times d_{(\ell)}})]$$

Therefore, we have

$$E\left(\frac{\partial}{\partial\boldsymbol{\theta}'}\left(\mathrm{vech}^0\left\{\boldsymbol{\eta}^*_t\left(\boldsymbol{\eta}^*_t\right)'\right\}\right)_{\boldsymbol{\theta}_0}\right) = -\frac{1}{2}\left(\mathbf{M}_{21}\boldsymbol{\Omega}\ \ \mathbf{M}_{31}\boldsymbol{\Omega}\ \dots\ \mathbf{M}_{m,m-1}\boldsymbol{\Omega}\right)' = -\frac{1}{2}\boldsymbol{\Lambda}'.$$

By the law of large numbers, it follows that

$$\frac{1}{n}\sum_{t=1}^{n}\left(\frac{\partial\mathrm{vech}^0\left\{\boldsymbol{\eta}^*_t\left(\boldsymbol{\eta}^*_t\right)'\right\}}{\partial\boldsymbol{\theta}'}\right)_{\boldsymbol{\theta}_0} \to -\frac{1}{2}\boldsymbol{\Lambda}', \quad \text{a.s.}$$

To complete the proof of iii), we will show that similarly to (**??**), for any $\varepsilon > 0$, there exists a neighborhood $\mathcal{V}(\boldsymbol{\theta}_0)$ of $\boldsymbol{\theta}_0$ such that, almost surely,

$$\lim_{n\to\infty}\frac{1}{n}\sum_{t=1}^{n}\sup_{\boldsymbol{\theta}\in\mathcal{V}(\boldsymbol{\theta}_0)}\left\|\left(\frac{\partial\mathrm{vech}^0\left\{\boldsymbol{\eta}^*_t\left(\boldsymbol{\eta}^*_t\right)'\right\}}{\partial\boldsymbol{\theta}'}\right)_{\boldsymbol{\theta}} - \left(\frac{\partial\mathrm{vech}^0\left\{\boldsymbol{\eta}^*_t\left(\boldsymbol{\eta}^*_t\right)'\right\}}{\partial\boldsymbol{\theta}'}\right)_{\boldsymbol{\theta}_0}\right\| \le \varepsilon.$$

The latter convergence is equivalent to

$$\lim_{n\to\infty}\frac{1}{n}\sum_{t=1}^{n}\sup_{\boldsymbol{\theta}^{(k)}\in\mathcal{V}(\boldsymbol{\theta}_0^{(k)})}\ \sup_{\boldsymbol{\theta}^{(\ell)}\in\mathcal{V}(\boldsymbol{\theta}_0^{(\ell)})}\left\|\frac{\epsilon_{kt}}{\sigma_{kt}(\boldsymbol{\theta}^{(k)})}\frac{\epsilon_{\ell t}}{\sigma_{\ell t}(\boldsymbol{\theta}^{(\ell)})}\frac{1}{\sigma_{kt}(\boldsymbol{\theta}^{(k)})}\frac{\partial\sigma_{kt}(\boldsymbol{\theta}^{(k)})}{\partial\boldsymbol{\theta}'}\right.$$
$$\left. - \frac{\epsilon_{kt}}{\sigma_{kt}(\boldsymbol{\theta}_0^{(k)})}\frac{\epsilon_{\ell t}}{\sigma_{\ell t}(\boldsymbol{\theta}_0^{(\ell)})}\frac{1}{\sigma_{kt}(\boldsymbol{\theta}_0^{(k)})}\frac{\partial\sigma_{kt}(\boldsymbol{\theta}_0^{(k)})}{\partial\boldsymbol{\theta}'}\right\| \le \varepsilon, \quad a.s. \qquad \text{(B.1)}$$

for any $k, \ell = 1, \ldots, m$. By the arguments used to prove iii) in the proof of the asymptotic normality of $\hat{\boldsymbol{\theta}}_n$, we have

$$E \sup_{\boldsymbol{\theta}^{(k)} \in \mathcal{V}(\boldsymbol{\theta}_0^{(k)})} \sup_{\boldsymbol{\theta}^{(\ell)} \in \mathcal{V}(\boldsymbol{\theta}_0^{(\ell)})} \left\| \frac{\epsilon_{kt}}{\sigma_{kt}(\boldsymbol{\theta}^{(k)})} \frac{\epsilon_{\ell t}}{\sigma_{\ell t}(\boldsymbol{\theta}^{(\ell)})} \frac{1}{\sigma_{kt}(\boldsymbol{\theta}^{(k)})} \frac{\partial \sigma_{kt}(\boldsymbol{\theta}^{(k)})}{\partial \boldsymbol{\theta}'} \right\| < \infty,$$

from which (B.1) follows. Thus, iii) is established. It remains to show iv). We note that

$$\boldsymbol{Z}_t := \begin{pmatrix} \dot{\boldsymbol{\ell}}_t(\boldsymbol{\theta}_0) \\ \mathrm{vech}^0 \left\{ \boldsymbol{\eta}_t^* \left(\boldsymbol{\eta}_t^*\right)' - \boldsymbol{R} \right\} \end{pmatrix} = \begin{pmatrix} F(\boldsymbol{\eta}_t^*)\boldsymbol{d}_t \\ \mathrm{vech}^0 \left\{ \boldsymbol{\eta}_t^* \left(\boldsymbol{\eta}_t^*\right)' - \boldsymbol{R} \right\} \end{pmatrix}$$

is measurable with respect to the σ-field $\mathcal{F}_t$ generated by $\{\boldsymbol{\eta}_u^*, u \leq t\}$. We have, using the independence of the sequence $(\boldsymbol{\eta}_t^*)$ under (2.9),

$$\begin{aligned} \mathrm{Var}(F(\boldsymbol{\eta}_t^*)\boldsymbol{d}_t) &= E\{F(\boldsymbol{\eta}_t^*)E(\boldsymbol{d}_t\boldsymbol{d}_t')F(\boldsymbol{\eta}_t^*)\} = \boldsymbol{J}^*, \\ \mathrm{Cov}\left[F(\boldsymbol{\eta}_t^*)\boldsymbol{d}_t, \mathrm{vech}^0\left\{\boldsymbol{\eta}_t^* \left(\boldsymbol{\eta}_t^*\right)'\right\}\right] &= E\left\{F(\boldsymbol{\eta}_t^*)\boldsymbol{\Omega}\left[\mathrm{vech}^0\left\{\boldsymbol{\eta}_t^* \left(\boldsymbol{\eta}_t^*\right)'\right\}\right]'\right\} = \boldsymbol{L}. \end{aligned}$$

Thus, $\forall \lambda \in \mathbb{R}^{d+m(m-1)/2}$, the sequence $\{\lambda' \boldsymbol{Z}_t, \mathcal{F}_t\}_t$ is an ergodic, stationary and square integrable martingale difference. The conclusion follows from the central limit theorem of Billingsley (1961).

We are now in a position to complete the proof of Theorem 5.2. Since $\hat{\boldsymbol{\theta}}_n^{(k)}$ converges to $\boldsymbol{\theta}_0^{(k)}$, which stands in the interior of the parameter space by **A7**, the derivative of the criterion $\tilde{Q}_n^{(k)}$ is equal to zero at $\hat{\boldsymbol{\theta}}_n^{(k)}$. In view of point ii), we thus have by a Taylor expansion of $Q_n^{(k)}$ at $\boldsymbol{\theta}_0^{(k)}$,

$$\sqrt{n}\left(\hat{\boldsymbol{\theta}}_n^{(k)} - \boldsymbol{\theta}_0^{(k)}\right) \overset{o_P(1)}{=} -\left(\frac{1}{n}\sum_{t=1}^n \frac{\partial^2 \ell_{kt}(\boldsymbol{\theta}_{ij}^{*(k)})}{\partial \theta_i^{(k)} \partial \theta_j^{(k)}}\right)^{-1} \frac{1}{\sqrt{n}} \sum_{t=1}^n \frac{\partial}{\partial \boldsymbol{\theta}^{(k)}} \ell_{kt}(\boldsymbol{\theta}_0^{(k)})$$

where the $\boldsymbol{\theta}_{ij}^{*(k)}$'s are between $\hat{\boldsymbol{\theta}}_n^{(k)}$ and $\boldsymbol{\theta}_0^{(k)}$. Thus we have, using iii) and iv),

$$\sqrt{n}\left(\hat{\boldsymbol{\theta}}_n - \boldsymbol{\theta}_0\right) \overset{o_P(1)}{=} -\boldsymbol{J}_0^{-1} \frac{1}{\sqrt{n}} \sum_{t=1}^n \dot{\boldsymbol{\ell}}_t(\boldsymbol{\theta}_0).$$

Another Taylor expansion around $\boldsymbol{\theta}_0$ yields,

$$\begin{aligned} &\sqrt{n}(\hat{\boldsymbol{\rho}}_n - \boldsymbol{\rho}_0) \\ = \;& \frac{1}{\sqrt{n}} \sum_{t=1}^n \mathrm{vech}^0 \left\{ \boldsymbol{\eta}_t^* \left(\boldsymbol{\eta}_t^*\right)' - \boldsymbol{R} \right\} + \frac{1}{n} \sum_{t=1}^n \frac{\partial}{\partial \boldsymbol{\theta}'} \left(\mathrm{vech}^0 \left\{ \tilde{\boldsymbol{\eta}}_t^* \left(\tilde{\boldsymbol{\eta}}_t^*\right)' \right\}\right)_{\tilde{\boldsymbol{\theta}}_n} \sqrt{n}\left(\hat{\boldsymbol{\theta}}_n - \boldsymbol{\theta}_0\right), \end{aligned}$$

where $\tilde{\boldsymbol{\eta}}_t^* = \tilde{\boldsymbol{\eta}}_t^*(\boldsymbol{\theta}) = \tilde{\boldsymbol{D}}_t^{-1}(\boldsymbol{\theta})\boldsymbol{\epsilon}_t$ and $\tilde{\boldsymbol{D}}_t(\boldsymbol{\theta}) = \mathrm{diag}\{\tilde{\sigma}_{1t}(\boldsymbol{\theta}^{(1)}), \ldots, \tilde{\sigma}_{mt}(\boldsymbol{\theta}^{(m)})\}$, and $\tilde{\boldsymbol{\theta}}_n$ is between $\hat{\boldsymbol{\theta}}_n$ and $\boldsymbol{\theta}_0$. It follows that, using v) and vi), denoting by $\boldsymbol{I}$ the identity matrix of size $m(m-1)/2$ and by $\mathbf{0}$ is null matrix of size $d \times m(m-1)/2$,

$$\begin{pmatrix} \sqrt{n}\left(\hat{\boldsymbol{\theta}}_n - \boldsymbol{\theta}_0\right) \\ \sqrt{n}(\hat{\boldsymbol{\rho}}_n - \boldsymbol{\rho}_0) \end{pmatrix} \stackrel{o_P(1)}{=} \begin{pmatrix} -\boldsymbol{J}_0^{-1} & \mathbf{0} \\ \frac{1}{2}\boldsymbol{\Lambda}'\boldsymbol{J}_0^{-1} & \boldsymbol{I} \end{pmatrix} \frac{1}{\sqrt{n}} \sum_{t=1}^{n} \boldsymbol{Z}_t. \tag{B.2}$$

The asymptotic distribution of Theorem 5.2 thus follows from iv).

It remains to establish that $\boldsymbol{\Sigma}$ is non singular. By (B.2), it suffices to show that $\mathrm{Var}(\boldsymbol{Z}_t)$ is nonsingular. We will show that for any $\boldsymbol{x} = (\boldsymbol{x}_i) \in \mathbb{R}^d$, where $\boldsymbol{x}_i \in \mathbb{R}^{d_i}$, for any $\boldsymbol{y} = (y_{k\ell}) \in \mathbb{R}^{m(m-1)/2}$ and any $c \in \mathbb{R}$,

$$\boldsymbol{x}'\dot{\boldsymbol{\ell}}_t(\boldsymbol{\theta}_0) + \boldsymbol{y}'\mathrm{vech}^0\left\{\boldsymbol{\eta}_t^*\left(\boldsymbol{\eta}_t^*\right)' - \boldsymbol{R}\right\} = c, \ a.s. \quad \Rightarrow \quad \boldsymbol{x} = 0 \text{ and } \boldsymbol{y} = 0. \tag{B.3}$$

Assume that the left-hand side of (B.3) holds. Then we have

$$\sum_{i=1}^{m}\left(1 - \eta_{it}^{*2}\right) z_{i,t-1} + \sum_{k \neq \ell} y_{k\ell}\left(\eta_{kt}^*\eta_{\ell t}^* - R_{k\ell}\right) = c$$

where $z_{i,t-1} = \frac{1}{\sigma_{it}^2}\boldsymbol{x}_i'\frac{\partial\sigma_{it}^2(\boldsymbol{\theta}_0^{(i)})}{\partial\boldsymbol{\theta}^{(i)}}$. It follows that $\boldsymbol{a}_{t-1}'\mathrm{vech}\left\{\boldsymbol{\eta}_t^*\left(\boldsymbol{\eta}_t^*\right)'\right\} = b_{t-1}$ for some vector $\boldsymbol{a}_{t-1}$ and some scalar b_{t-1} belonging to the past. By **A13** and (2.9) we must have $\boldsymbol{a}_{t-1} = \mathbf{0}$ and $b_{t-1} = 0$. Noting that the $z_{i,t-1}$ are components of $\boldsymbol{a}_{t-1}$, we must have $z_{i,t-1} = 0$ for $i = 1, \ldots, m$, in contradiction with **A12** unless if $\boldsymbol{x} = \mathbf{0}$. It is then straightforward to show that $\boldsymbol{y} = 0$ and the proof is complete.

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