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Simulated Method of Moments Estimation for Copula-Based Multivariate Models

Dong Hwan OH and Andrew J. PATTON

This article considers the estimation of the parameters of a copula via a simulated method of moments (MM) type approach. This approach is attractive when the likelihood of the copula model is not known in closed form, or when the researcher has a set of dependence measures or other functionals of the copula that are of particular interest. The proposed approach naturally also nests MM and generalized method of moments estimators. Drawing on results for simulation-based estimation and on recent work in empirical copula process theory, we show the consistency and asymptotic normality of the proposed estimator, and obtain a simple test of overidentifying restrictions as a specification test. The results apply to both iid and time series data. We analyze the finite-sample behavior of these estimators in an extensive simulation study. We apply the model to a group of seven financial stock returns and find evidence of statistically significant tail dependence, and mild evidence that the dependence between these assets is stronger in crashes than booms. Supplementary materials for this article are available online.

KEY WORDS: Correlation; Dependence; Inference; SMM.

1. INTRODUCTION

Copula-based models for multivariate distributions are widely used in a variety of applications, including actuarial science and insurance (Embrechts, McNeil, and Straumann 2002; Rosenberg and Schuermann 2006), economics (Brendstrup and Paarsch 2007; Bonhomme and Robin 2009), epidemiology (Clayton 1978; Fine and Jiang 2000), finance (Cherubini, Luciano, and Vecchiato 2004; Patton 2006a), geology and hydrology (Cook and Johnson 1981; Genest and Favre 2007) among many others. An important benefit they provide is the flexibility to specify the marginal distributions separately from the dependence structure, without imposing that they come from the same family of joint distributions.

While copulas provide a great deal of flexibility in theory, the search for copula models that work well in practice is an ongoing one. This search has spawned a number of new and flexible models (see Demarta and McNeil 2005; McNeil, Frey, and Embrechts 2005; Smith et al. 2010; Smith, Gan, and Kohn forthcoming; Oh and Patton 2011) among others. Some of these models are such that the likelihood of the copula is either not known in closed form, or is complicated to obtain and maximize, motivating the consideration of estimation methods other than maximum-likelihood estimation (MLE). Moreover, in many financial applications, the estimated copula model is used in pricing a derivative security, such as a collateralized debt obligation or a credit default swap (CDO or CDS), and it may be of interest to minimize the pricing error (the observed market price less the model-implied price of the security) in calibrating the parameters of the model. In some cases, the mapping from the parameter(s) of the copula to dependence measures (such as Spearman's or Kendall's rank correlation, for example) or to the price of the derivative contract is known in closed form, thus allowing for method of moments (MM) or generalized method of moments (GMM) estimation. In general, however, this mapping is unknown, and an alternative estimation method is required. We consider a simple yet widely applicable simulation-based approach to address this problem.

This article presents the asymptotic properties of a simulation-based estimator of the parameters of a copula model. We consider both iid and time series data, and the case that the marginal distributions are estimated using the empirical distribution function (EDF). The estimation method we consider shares features with the simulated method of moments (SMM) (see, e.g., McFadden 1989; Pakes and Pollard 1989); however, the presence of the EDF in the sample "moments" means that existing results on SMM are not directly applicable. We draw on well-known results on SMM estimators (see, e.g., Newey and McFadden 1994) and recent results from empirical process theory for copulas (see Fermanian, Radulović, and Wegkamp 2004; Chen and Fan 2006; Rémillard 2010) to show the consistency and asymptotic normality of simulation-based estimators of copula models. To the best of our knowledge, simulation-based estimation of copula models has not previously been considered in the literature. An extensive simulation study verifies that the asymptotic results provide a good approximation in finite samples. We illustrate the results with an application to a model of the dependence between the equity returns on seven financial firms during the recent crisis period.

In addition to maximum likelihood, several other estimation methods have been considered for copula-based multivariate models. First, multistage maximum likelihood, also known as "inference functions for margins," [see Joe and Xu (1996) and Joe (2005) for iid data, Patton (2006b) for time series data, and Song, Fan, and Kalbfleisch (2005) for an iterative multistage ML method with improved efficiency] is one of the most widely used estimation methods. Like MLE, this method only applies when the marginal distributions are parametric. When the marginal

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distribution models are correctly specified, this improves the efficiency of the estimator relative to the proposed SMM approach; however, it introduces the possibility of misspecified marginal distributions, which can have deleterious effects on the copula parameter estimates (see Kim, Silvapulle, and Silvapulle 2007). A second popular method is semi-parametric MLE [see Genest, Ghoudi, and Rivest (1995) for iid data and Chen and Fan (2006), Chan et al. (2009), and Chen, Fan, and Tsyrennikov (2006) for time series data]. This method yields efficient estimates of the copula parameters, however it requires, of course, the copula likelihood and for some more complicated models, the likelihood can be cumbersome to derive or to compute, for example, the "stochastic copula" model of Hafner and Manner (2012) or the high-dimension factor copula model of Oh and Patton (2011). An alternative, long-standing, estimator is the MM estimator [see Genest (1987) and Genest and Rivest (1993) for iid data and Rémillard (2010) for time series data]. This estimator exploits the known one-to-one mapping between the parameters of certain copulas and certain measures of dependence, and usually has the benefit of being very fast to compute. The SMM estimator proposed in this article generalizes MM to allow for the consideration of overidentified models (i.e., when we have more implied dependence measures than unknown parameters) and for dependence measures that are not known closed-form functions of the copula parameters, using simulations to obtain the mapping instead.

2. SIMULATION-BASED ESTIMATION OF COPULA MODELS

We consider the same class of data-generating processes (DGPs) as Chen and Fan (2006), Chan et al. (2009), and Rémillard (2010). This class allows each variable to have time-varying conditional mean and conditional variance, each governed by parametric models, with some unknown marginal distribution. As in those papers, and also earlier papers such as Genest and Rivest (1993) and Genest, Ghoudi, and Rivest (1995), we estimate the marginal distributions using the EDF. The conditional copula of the data is assumed to belong to a parametric family with unknown parameter $\boldsymbol{\theta}_0$. The DGP we consider is

$$[Y_{1t}, \ldots, Y_{Nt}]' \equiv \mathbf{Y}_t = \boldsymbol{\mu}_t\left(\boldsymbol{\phi}_0\right) + \boldsymbol{\sigma}_t\left(\boldsymbol{\phi}_0\right)\boldsymbol{\eta}_t,$$

where

$$\begin{aligned}\boldsymbol{\mu}_t(\boldsymbol{\phi}) &\equiv [\mu_{1t}(\boldsymbol{\phi}), \ldots, \mu_{Nt}(\boldsymbol{\phi})]',\\ \boldsymbol{\sigma}_t(\boldsymbol{\phi}) &\equiv \operatorname{diag}\{\sigma_{1t}(\boldsymbol{\phi}), \ldots, \sigma_{Nt}(\boldsymbol{\phi})\},\\ [\eta_{1t}, \ldots, \eta_{Nt}]' &\equiv \boldsymbol{\eta}_t \sim \text{iid } \mathbf{F}_{\boldsymbol{\eta}} = \mathbf{C}(F_1, \ldots, F_N; \boldsymbol{\theta}_0),\end{aligned} \quad (1)$$

where $\boldsymbol{\mu}_t$ and $\boldsymbol{\sigma}_t$ are $\mathcal{F}_{t-1}$-measurable and independent of $\boldsymbol{\eta}_t$. $\mathcal{F}_{t-1}$ is the sigma field containing information generated by $\{\mathbf{Y}_{t-1}, \mathbf{Y}_{t-2}, \ldots\}$. The $r \times 1$ vector of parameters governing the dynamics of the variables, $\boldsymbol{\phi}_0$, is assumed to be $\sqrt{T}$ consistently estimable, which holds under mild conditions for many commonly used models for multivariate time series, such as ARMA models, GARCH models, stochastic volatility models, etc. If $\boldsymbol{\phi}_0$ is known, or if $\boldsymbol{\mu}_t$ and $\boldsymbol{\sigma}_t$ are known constant, then the model becomes one for iid data. Our task is to estimate the $p \times 1$ vector of copula parameters, $\boldsymbol{\theta}_0 \in \boldsymbol{\Theta}$, based on the (estimated) standardized residual $\{\hat{\boldsymbol{\eta}}_t \equiv \boldsymbol{\sigma}_t^{-1}(\hat{\boldsymbol{\phi}})[\mathbf{Y}_t - \boldsymbol{\mu}_t(\hat{\boldsymbol{\phi}})]\}_{t=1}^T$ and simulations from the copula model, $\mathbf{C}(\cdot; \boldsymbol{\theta})$.

2.1 Definition of the SMM Estimator

We consider simulation from some parametric multivariate distribution, $\mathbf{F}_x(\boldsymbol{\theta})$, with marginal distributions $G_i(\boldsymbol{\theta})$, and copula $\mathbf{C}(\boldsymbol{\theta})$. This allows us to consider cases where it is possible to simulate directly from the copula model [in which case the G_i are all Unif(0, 1)] and also cases where the copula model is embedded in some joint distribution with unknown marginal distributions, such as the factor copula models of Oh and Patton (2011).

We use only "pure" dependence measures as those are unaffected by changes in the marginal distributions of simulated data. This rules out linear correlation, which contains information on the copula but is also affected by the marginal distributions. Dependence measures like Spearman's rank correlation, quantile dependence, and Kendall's tau are functions only of the copula (see, e.g., Joe 1997; Nelsen 2006). These measures for the pair (η_i, η_j) are defined as:

$$\rho^{ij} \equiv 12E[F_i(\eta_i)F_j(\eta_j)] - 3 = 12\iint uv\,dC_{ij}(u, v) - 3, \quad (2)$$

$$\lambda_q^{ij} \equiv \begin{cases} P[F_i(\eta_i) \leq q | F_j(\eta_j) \leq q] = \dfrac{C_{ij}(q, q)}{q}, & q \in (0, 0.5] \\ P[F_i(\eta_i) > q | F_j(\eta_j) > q] = \dfrac{1 - 2q + C_{ij}(q, q)}{1 - q}, & q \in (0.5, 1), \end{cases} \quad (3)$$

$$\tau^{ij} \equiv 4E[C_{ij}(F_i(\eta_i), F_j(\eta_j))] - 1, \quad (4)$$

where C_{ij} is the copula of (η_i, η_j). The sample counterparts are defined as

$$\hat{\rho}^{ij} \equiv \frac{12}{T}\sum_{t=1}^{T} \hat{F}_i(\hat{\eta}_{it})\hat{F}_j(\hat{\eta}_{jt}) - 3, \quad (5)$$

$$\hat{\lambda}_q^{ij} \equiv \begin{cases} \dfrac{1}{Tq}\displaystyle\sum_{t=1}^{T} 1\{\hat{F}_i(\hat{\eta}_{it}) \leq q, \hat{F}_j(\hat{\eta}_{jt}) \leq q\}, & q \in (0, 0.5] \\ \dfrac{1}{T(1-q)}\displaystyle\sum_{t=1}^{T} 1\{\hat{F}_i(\hat{\eta}_{it}) > q, \hat{F}_j(\hat{\eta}_{jt}) > q\}, & q \in (0.5, 1), \end{cases} \quad (6)$$

$$\hat{\tau}^{ij} \equiv \frac{4}{T}\sum_{t=1}^{T} \hat{C}_{ij}(\hat{F}_i(\hat{\eta}_{it}), \hat{F}_j(\hat{\eta}_{jt})) - 1, \quad (7)$$

where $\hat{F}_i(y) \equiv (T+1)^{-1}\sum_{t=1}^{T} 1\{\hat{\eta}_{it} \leq y\}$, and $\hat{C}_{ij}(u, v) \equiv (T+1)^{-1}\sum_{t=1}^{T} 1\{\hat{F}_i(\hat{\eta}_{it}) \leq u, \hat{F}_j(\hat{\eta}_{jt}) \leq v\}$. Counterparts based on simulations are denoted by $\tilde{\rho}^{ij}(\boldsymbol{\theta})$, $\tilde{\lambda}_q^{ij}(\boldsymbol{\theta})$, and $\tilde{\tau}^{ij}(\boldsymbol{\theta})$.

Let $\tilde{\mathbf{m}}_S(\boldsymbol{\theta})$ be a $(m \times 1)$ vector of dependence measures computed using S simulations from $\mathbf{F}_x(\boldsymbol{\theta})$, $\{\mathbf{X}_s\}_{s=1}^S$, and let $\hat{\mathbf{m}}_T$ be the corresponding vector of dependence measures computed using the standardized residuals $\{\hat{\boldsymbol{\eta}}_t\}_{t=1}^T$. These vectors can also contain linear combinations of dependence measures, a feature that is useful when considering estimation of high-dimension models. Define the difference between these as

$$\mathbf{g}_{T,S}(\boldsymbol{\theta}) \equiv \hat{\mathbf{m}}_T - \tilde{\mathbf{m}}_S(\boldsymbol{\theta}). \tag{8}$$

Our SMM estimator is based on searching across $\boldsymbol{\theta} \in \boldsymbol{\Theta}$ to make this difference as small as possible. The estimator is defined as

$$\hat{\boldsymbol{\theta}}_{T,S} \equiv \arg\min_{\boldsymbol{\theta}\in\boldsymbol{\Theta}} Q_{T,S}(\boldsymbol{\theta}),$$

where

$$Q_{T,S}(\boldsymbol{\theta}) \equiv \mathbf{g}'_{T,S}(\boldsymbol{\theta})\hat{\mathbf{W}}_T\mathbf{g}_{T,S}(\boldsymbol{\theta}), \tag{9}$$

and $\hat{\mathbf{W}}_T$ is some positive definite weight matrix, which may depend on the data. In the subsections below, we establish the consistency and asymptotic normality of this estimator, provide a consistent estimator of its asymptotic covariance matrix, and obtain a test based on overidentifying restrictions. The supplemental appendix (available online) presents details on the computation of the objective function.

2.2 Consistency of the SMM Estimator

The estimation problem here differs in two important ways from standard GMM or M-estimation: First, the objective function, $Q_{T,S}(\boldsymbol{\theta})$, is not continuous in $\boldsymbol{\theta}$, since $\tilde{\mathbf{m}}_S(\boldsymbol{\theta})$ will be a number in a set of discrete values as $\boldsymbol{\theta}$ varies on $\boldsymbol{\Theta}$, for example, $\{0, \frac{1}{Sq}, \frac{2}{Sq}, \ldots, \frac{S}{Sq}\}$ for a lower quantile dependence.This problem would vanish if, for the copula model being considered, we knew the mapping $\boldsymbol{\theta} \longmapsto m_0(\boldsymbol{\theta}) \equiv \lim_{S\to\infty}\tilde{\mathbf{m}}_S(\boldsymbol{\theta})$ in closed form. The second difference is that a law of large numbers is not available to show the pointwise convergence of $\mathbf{g}_{T,S}(\boldsymbol{\theta})$, as the functions $\hat{\mathbf{m}}_T$ and $\tilde{\mathbf{m}}_S(\boldsymbol{\theta})$ both involve EDFs. We use recent developments in empirical process theory to overcome this difficulty.

We now list some assumptions that are required for our results to hold.

Assumption 1.

(i) The distributions $\mathbf{F}_\eta$ and $\mathbf{F}_x$ are continuous.
(ii) Every bivariate marginal copula C_{ij} of $\mathbf{C}$ has continuous partial derivatives with respect to u_i and u_j.

If the data $\mathbf{Y}_t$ are iid, for example, if $\boldsymbol{\mu}_t$ and $\boldsymbol{\sigma}_t$ are known constant in Equation (1), or if $\boldsymbol{\phi}_0$ is known, then Assumption 1 is sufficient to prove Proposition 1 using the results of Fermanian, Radulović, and Wegkamp (2004). If, however, estimated standardized residuals are used in the estimation of the copula, then more assumptions are necessary to control the estimation error coming from the models for the conditional means and conditional variances. We combine assumptions A1–A6 in Rémillard (2010) in the following assumption. First, define $\boldsymbol{\gamma}_{0t} = \boldsymbol{\sigma}_t^{-1}(\hat{\boldsymbol{\phi}})\dot{\boldsymbol{\mu}}_t(\hat{\boldsymbol{\phi}})$ and $\boldsymbol{\gamma}_{1kt} = \boldsymbol{\sigma}_t^{-1}(\hat{\boldsymbol{\phi}})\dot{\boldsymbol{\sigma}}_{kt}(\hat{\boldsymbol{\phi}})$, where $\dot{\boldsymbol{\mu}}_t(\boldsymbol{\phi}) = \frac{\partial\boldsymbol{\mu}_t(\boldsymbol{\phi})}{\partial\boldsymbol{\phi}'}$, $\dot{\boldsymbol{\sigma}}_{kt}(\boldsymbol{\phi}) = \frac{\partial[\boldsymbol{\sigma}_t(\boldsymbol{\phi})]_{k\text{th column}}}{\partial\boldsymbol{\phi}'}$, $k = 1, \ldots, N$. Define $\mathbf{d}_t$ as

$$\mathbf{d}_t = \boldsymbol{\eta}_t - \hat{\boldsymbol{\eta}}_t - \left(\boldsymbol{\gamma}_{0t} + \sum_{k=1}^{N}\eta_{kt}\boldsymbol{\gamma}_{1kt}\right)(\hat{\boldsymbol{\phi}} - \boldsymbol{\phi}_0),$$

where η_{kt} is kth row of $\boldsymbol{\eta}_t$ and both $\boldsymbol{\gamma}_{0t}$ and $\boldsymbol{\gamma}_{1kt}$ are $\mathcal{F}_{t-1}$-measurable.

Assumption 2.

(i) $\frac{1}{T}\sum_{t=1}^{T}\boldsymbol{\gamma}_{0t} \xrightarrow{p} \boldsymbol{\Gamma}_0$ and $\frac{1}{T}\sum_{t=1}^{T}\boldsymbol{\gamma}_{1kt} \xrightarrow{p} \boldsymbol{\Gamma}_{1k}$, where $\boldsymbol{\Gamma}_0$ and $\boldsymbol{\Gamma}_{1k}$ are deterministic for $k = 1, \ldots, N$.
(ii) $\frac{1}{T}\sum_{t=1}^{T}E(\|\boldsymbol{\gamma}_{0t}\|)$, $\frac{1}{T}\sum_{t=1}^{T}E(\|\boldsymbol{\gamma}_{0t}\|^2)$, $\frac{1}{T}\sum_{t=1}^{T}E(\|\boldsymbol{\gamma}_{1kt}\|)$, and $\frac{1}{T}\sum_{t=1}^{T}E(\|\boldsymbol{\gamma}_{1kt}\|^2)$ are bounded for $k = 1, \ldots, N$.
(iii) There exists a sequence of positive terms $r_t > 0$, so that $\sum_{t\geq 1} r_t < \infty$ and such that the sequence $\max_{1\leq t\leq T}\|\mathbf{d}_t\|/r_t$ is tight.
(iv) $\max_{1\leq t\leq T}\|\boldsymbol{\gamma}_{0t}\|/\sqrt{T} = o_p(1)$ and $\max_{1\leq t\leq T}\eta_{kt}\|\boldsymbol{\gamma}_{1kt}\|/\sqrt{T} = o_p(1)$ for $k = 1, \ldots, N$.
(v) $(\alpha_T, \sqrt{T}(\hat{\boldsymbol{\phi}} - \boldsymbol{\phi}_0))$ weakly converges to a continuous Gaussian process in $[0,1]^N \times \mathbb{R}^r$, where α_T is the empirical copula process of uniform random variables:

$$\alpha_T = \frac{1}{\sqrt{T}}\sum_{t=1}^{T}\left\{\prod_{k=1}^{N}1(U_{kt} \leq u_k) - C(\mathbf{u})\right\}.$$

(vi) $\frac{\partial\mathbf{F}_\eta}{\partial\eta_k}$ and $\eta_k\frac{\partial\mathbf{F}_\eta}{\partial\eta_k}$ are bounded and continuous on $\bar{\mathbb{R}}^N = [-\infty, +\infty]^N$ for $k = 1, \ldots, N$.

With these two assumptions, sample rank correlation and quantile dependence converge in probability to their population counterparts [see Theorems 3 and 6 of Fermanian, Radulović, and Wegkamp (2004) for the iid case and combine with Corollary 1 of Rémillard (2010) for the time series case]. (See Lemma 1 of the supplemental appendix, available online, for details.) When applied to simulated data, this convergence holds pointwise for any $\boldsymbol{\theta}$. Thus, $\mathbf{g}_{T,S}(\boldsymbol{\theta})$ converges in probability to the population moment functions defined as follows:

$$\mathbf{g}_{T,S}(\boldsymbol{\theta}) \equiv \hat{\mathbf{m}}_T - \tilde{\mathbf{m}}_S(\boldsymbol{\theta}) \xrightarrow{p} \mathbf{g}_0(\boldsymbol{\theta}) \equiv \mathbf{m}_0(\boldsymbol{\theta}_0) - \mathbf{m}_0(\boldsymbol{\theta}),$$
$$\text{for } \forall\boldsymbol{\theta} \in \boldsymbol{\Theta} \text{ as } T, S \to \infty. \tag{10}$$

We define the population objective function as

$$Q_0(\boldsymbol{\theta}) = \mathbf{g}_0(\boldsymbol{\theta})'\mathbf{W}_0\mathbf{g}_0(\boldsymbol{\theta}), \tag{11}$$

where $\mathbf{W}_0$ is the probability limit of $\hat{\mathbf{W}}_T$. The convergence of $\mathbf{g}_{T,S}(\boldsymbol{\theta})$ and $\hat{\mathbf{W}}_T$ implies that

$$Q_{T,S}(\boldsymbol{\theta}) \xrightarrow{p} Q_0(\boldsymbol{\theta}) \quad \text{for } \forall\boldsymbol{\theta} \in \boldsymbol{\Theta} \text{ as } T, S \to \infty.$$

For consistency of our estimator, we need, as usual, *uniform* convergence of $Q_{T,S}(\boldsymbol{\theta})$, but as this function is not continuous in $\boldsymbol{\theta}$ and a law of large numbers is not available, the standard approach based on a uniform law of large numbers is not available. We instead use results on the stochastic equicontinuity of $\mathbf{g}_{T,S}(\boldsymbol{\theta})$ based on Andrews (1994) and Newey and McFadden (1994).

Assumption 3.

(i) $\mathbf{g}_0(\boldsymbol{\theta}) \neq 0$ for $\boldsymbol{\theta} \neq \boldsymbol{\theta}_0$.
(ii) $\boldsymbol{\Theta}$ is compact.
(iii) Every bivariate marginal copula $C_{ij}(u_i, u_j; \boldsymbol{\theta})$ of $\mathbf{C}(\boldsymbol{\theta})$ on $(u_i, u_j) \in (0,1) \times (0,1)$ is Lipschitz continuous on $\boldsymbol{\Theta}$.
(iv) $\hat{\mathbf{W}}_T$ is $O_p(1)$ and converges in probability to $\mathbf{W}_0$, a positive definite matrix.

Proposition 1. Suppose that Assumptions 1, 2, and 3 hold. Then, $\hat{\boldsymbol{\theta}}_{T,S} \xrightarrow{p} \boldsymbol{\theta}_0$ as $T, S \to \infty$.

A sketch of all proofs is presented in the appendix, and detailed proofs are in the supplemental appendix (available online). Assumption 3(iii) is needed to prove the stochastic Lipschitz continuity of $\mathbf{g}_{T,S}(\boldsymbol{\theta})$, which is a sufficient condition for the stochastic equicontinuity of $\mathbf{g}_{T,S}(\boldsymbol{\theta})$, and can easily be shown to be satisfied for many bivariate parametric copulas. Assumption 3(ii) requires compactness of the parameter space, a common assumption, and is aided by having outside information (such as constraints from economic arguments) that allow the researcher to bound the set of plausible parameters. While Pakes and Pollard (1989) and McFadden (1989) show the consistency of the SMM estimator for T, S diverging at the same rate, Proposition 1 shows that the copula parameter is consistent at any relative rate of T and S as long as both diverge. If we know the function $\mathbf{m}(\boldsymbol{\theta})$ in closed form, then GMM is feasible and is equivalent to our estimator with $S/T \to \infty$ as $T, S \to \infty$.

We focus on weak consistency of our estimator because it suffices for our asymptotic distribution theory, presented below. A corresponding strong consistency result, that is, $\hat{\boldsymbol{\theta}}_{T,S} \xrightarrow{\text{a.s.}} \boldsymbol{\theta}_0$, may be obtained by drawing on recent work by Bouzebda and Zari (2011). The key is to show uniform strong convergence of the sample objective function, from which strong consistency of the estimator easily follows (see, e.g., Newey and McFadden 1994). Uniform strong consistency of the objective function can be shown by combining minor changes in the above assumptions (e.g., $\hat{\mathbf{W}}_T$ must converge a.s. to $\mathbf{W}_0$) with pointwise strong convergence of the objective function, which can be obtained using results of Bouzebda and Zari (2011).

2.3 Asymptotic Normality of the SMM Estimator

As $Q_{T,S}(\boldsymbol{\theta})$ is nondifferentiable, the standard approach based on a Taylor expansion is not available; however, the asymptotic normality of our estimator can still be established with some further assumptions:

Assumption 4.

(i) $\boldsymbol{\theta}_0$ is an interior point of $\boldsymbol{\Theta}$.
(ii) $\mathbf{g}_0(\boldsymbol{\theta})$ is differentiable at $\boldsymbol{\theta}_0$ with derivative $\mathbf{G}_0$ such that $\mathbf{G}_0'\mathbf{W}_0\mathbf{G}_0$ is nonsingular.
(iii) $\mathbf{g}_{T,S}(\hat{\boldsymbol{\theta}}_{T,S})'\hat{\mathbf{W}}_T\mathbf{g}_{T,S}(\hat{\boldsymbol{\theta}}_{T,S}) \le \inf_{\boldsymbol{\theta}\in\boldsymbol{\Theta}} \mathbf{g}_{T,S}(\boldsymbol{\theta})' \hat{\mathbf{W}}_T\mathbf{g}_{T,S}(\boldsymbol{\theta}) + o_p(1/T + 1/S)$.

The first assumption above is standard, and the third assumption is standard in simulation-based estimation problems (see, e.g., Newey and McFadden 1994). The rate at which the o_p term vanishes in part (iii) turns out to depend on the smaller of T or S, as $o_p(1/T + 1/S) = o_p(\min(T, S)^{-1})$. The second assumption requires the population objective function, $\mathbf{g}_0$, to be differentiable even though its finite-sample counterpart is not, which is common in simulation-based estimation. The nonsingularity of $\mathbf{G}_0'\mathbf{W}_0\mathbf{G}_0$ is sufficient for local identification of the parameters of this model at $\boldsymbol{\theta}_0$ (see Rothenberg 1971; Hall 2005). With these assumptions in hand, we obtain the following result:

Proposition 2. Suppose that Assumptions 1, 2, 3, and 4 hold. Then,

$$\frac{1}{\sqrt{1/T + 1/S}}(\hat{\boldsymbol{\theta}}_{T,S} - \boldsymbol{\theta}_0) \xrightarrow{d} N(0, \boldsymbol{\Omega}_0) \text{ as } T, S \to \infty, \quad (12)$$

where $\boldsymbol{\Omega}_0 = (\mathbf{G}_0'\mathbf{W}_0\mathbf{G}_0)^{-1}\mathbf{G}_0'\mathbf{W}_0\boldsymbol{\Sigma}_0\mathbf{W}_0\mathbf{G}_0(\mathbf{G}_0'\mathbf{W}_0\mathbf{G}_0)^{-1}$, and $\boldsymbol{\Sigma}_0 \equiv \text{avar}[\hat{\mathbf{m}}_T]$.

The rate of convergence is thus shown to equal $\min(T, S)^{1/2}$. In general, one would like to set S very large to minimize the impact of simulation error and obtain a $\sqrt{T}$ convergence rate; however, if the model is computationally costly to simulate, then the result for $S \ll T$ may be useful. When S and T diverge at different rates, the asymptotic variance of $\min(T, S)^{1/2}(\hat{\boldsymbol{\theta}}_{T,S} - \boldsymbol{\theta}_0)$ is simply $\boldsymbol{\Omega}_0$. When S and T diverge at the same rate, say $S/T \to k \in (0, \infty)$, the asymptotic variance of $\sqrt{T}(\hat{\boldsymbol{\theta}}_{T,S} - \boldsymbol{\theta}_0)$ is $(1 + 1/k)\boldsymbol{\Omega}_0$, which incorporates efficiency loss from simulation error. As usual, we find that $\boldsymbol{\Omega}_0 = (\mathbf{G}_0'\boldsymbol{\Sigma}_0^{-1}\mathbf{G}_0)^{-1}$ if $\mathbf{W}_0$ is the efficient weight matrix, $\boldsymbol{\Sigma}_0^{-1}$.

The proof of the above proposition uses recent results for empirical copula processes presented in Fermanian, Radulović, and Wegkamp (2004) and Rémillard (2010) to establish the asymptotic normality of the sample dependence measures, $\hat{\mathbf{m}}_T$, and requires us to establish the stochastic equicontinuity of the moment functions, $\mathbf{v}_{T,S}(\boldsymbol{\theta}) = \sqrt{T}[\mathbf{g}_{T,S}(\boldsymbol{\theta}) - \mathbf{g}_0(\boldsymbol{\theta})]$. These are shown in Lemmas 6 and 7 in the supplemental appendix (available online).

Chen and Fan (2006), Chan et al. (2009), and Rémillard (2010) show that estimation error from $\hat{\boldsymbol{\phi}}$ does not enter the asymptotic distribution of the copula parameter estimator for maximum likelihood or (analytical) moment-based estimators, and the above proposition shows that this surprising result also holds for the SMM-type estimators proposed here. In applications based on *parametric* models for the marginal distributions, the asymptotic covariance matrix of the copula parameter is more complicated. In such cases, the model is fully parametric and the estimation approach here is a form of two-stage GMM (or SMM). In the absence of simulations, this can be treated using existing methods (see, e.g., White 1994; Gouriéroux, Monfort, and Renault 1996). If simulations are used in the copula estimation step, then the lemmas presented in the appendix can be combined with existing results on two-stage GMM to obtain the limiting distribution. This is straightforward but requires some additional detailed notation, and so is not presented here.

2.4 Consistent Estimation of the Asymptotic Variance

The asymptotic variance of our estimator has the familiar form of standard GMM applications; however, the components $\boldsymbol{\Sigma}_0$ and $\mathbf{G}_0$ require more care in their estimation than in standard applications. We suggest using an iid bootstrap to estimate $\boldsymbol{\Sigma}_0$, with the following steps: (i) sample with replacement from the standardized residuals $\{\hat{\boldsymbol{\eta}}_t\}_{t=1}^T$ to obtain a B bootstrap samples, $\{\hat{\boldsymbol{\eta}}_t^{(b)}\}_{t=1}^T$, $b = 1, 2, \ldots, B$; (ii) use $\{\hat{\boldsymbol{\eta}}_t^{(b)}\}_{t=1}^T$ to compute the sample moments and denote as $\hat{\mathbf{m}}_T^{(b)}$; (iii) calculate the sample covariance matrix of $\hat{\mathbf{m}}_T^{(b)}$ across the bootstrap replications, and scale it by the sample size:

$$\hat{\mathbf{\Sigma}}_{T,B} = \frac{T}{B}\sum_{b=1}^{B}\left(\hat{\mathbf{m}}_T^{(b)} - \hat{\mathbf{m}}_T\right)\left(\hat{\mathbf{m}}_T^{(b)} - \hat{\mathbf{m}}_T\right)'. \quad (13)$$

For the estimation of $\mathbf{G}_0$, we suggest a numerical derivative of $\mathbf{g}_{T,S}(\boldsymbol{\theta})$ at $\hat{\boldsymbol{\theta}}_{T,S}$; however, the fact that $\mathbf{g}_{T,S}$ is nondifferentiable means that care is needed in choosing the step size for the numerical derivative. In particular, Proposition 3 shows that we need to let the step size go to zero, as usual, but *slower* than the inverse of the rate of convergence of the estimator [i.e., $1/\min(\sqrt{T}, \sqrt{S})$]. Let $\mathbf{e}_k$ denote the kth unit vector whose dimension is the same as that of $\boldsymbol{\theta}$, and let $\varepsilon_{T,S}$ denote the step size. A two-sided numerical derivative estimator $\hat{\mathbf{G}}_{T,S}$ of $\mathbf{G}$ has kth column as follows:

$$\hat{\mathbf{G}}_{T,S,k} = \frac{\mathbf{g}_{T,S}(\hat{\boldsymbol{\theta}}_{T,S}+\mathbf{e}_k\varepsilon_{T,S}) - \mathbf{g}_{T,S}(\hat{\boldsymbol{\theta}}_{T,S}-\mathbf{e}_k\varepsilon_{T,S})}{2\varepsilon_{T,S}}, \quad k = 1, 2, \ldots, p. \quad (14)$$

Combine this estimator with $\hat{\mathbf{W}}_T$ to form

$$\hat{\mathbf{\Omega}}_{T,S,B} = (\hat{\mathbf{G}}_{T,S}'\hat{\mathbf{W}}_T\hat{\mathbf{G}}_{T,S})^{-1}\hat{\mathbf{G}}_{T,S}'\hat{\mathbf{W}}_T\hat{\mathbf{\Sigma}}_{T,B}\hat{\mathbf{W}}_T\hat{\mathbf{G}}_{T,S} \times (\hat{\mathbf{G}}_{T,S}'\hat{\mathbf{W}}_T\hat{\mathbf{G}}_{T,S})^{-1}. \quad (15)$$

Proposition 3. Suppose that all assumptions of Proposition 2 are satisfied, and that $\varepsilon_{T,S} \to 0$, $\varepsilon_{T,S} \times \min(\sqrt{T}, \sqrt{S}) \to \infty$, and $B \to \infty$ as $T, S \to \infty$. Then, $\hat{\mathbf{\Sigma}}_{T,B} \xrightarrow{p} \mathbf{\Sigma}_0$, $\hat{\mathbf{G}}_{T,S} \xrightarrow{p} \mathbf{G}_0$ and $\hat{\mathbf{\Omega}}_{T,S,B} \xrightarrow{p} \mathbf{\Omega}_0$ as $T, S \to \infty$.

2.5 A Test of Overidentifying Restrictions

If the number of moments used in estimation is greater than the number of copula parameters, then it is possible to conduct a simple test of the overidentifying restrictions, which can be used as a specification test of the model.

Proposition 4. Suppose that all assumptions of Proposition 3 are satisfied and that the number of moments (m) is greater than the number of copula parameters (p). Then,

$$J_{T,S} \equiv \min(T, S)\mathbf{g}_{T,S}(\hat{\boldsymbol{\theta}}_{T,S})'\,\hat{\mathbf{W}}_T\mathbf{g}_{T,S}(\hat{\boldsymbol{\theta}}_{T,S}) \xrightarrow{d} \mathbf{u}'\mathbf{A}_0'\mathbf{A}_0\mathbf{u} \text{ as } T, S \to \infty,$$

where $\mathbf{u} \sim N(0, \mathbf{I})$ and $\mathbf{A}_0 \equiv \mathbf{W}_0^{1/2}\mathbf{\Sigma}_0^{1/2}\mathbf{R}_0$, $\mathbf{R}_0 \equiv \mathbf{I} - \mathbf{\Sigma}_0^{-1/2}\mathbf{G}_0(\mathbf{G}_0'\mathbf{W}_0\mathbf{G}_0)^{-1}\mathbf{G}_0'\mathbf{W}_0\mathbf{\Sigma}_0^{1/2}$. If $\hat{\mathbf{W}}_T = \hat{\mathbf{\Sigma}}_{T,B}^{-1}$, then $J_{T,S} \xrightarrow{d} \chi^2_{m-p}$ as usual.

As in standard applications, the above test statistic has a chi-squared limiting distribution if the efficient weight matrix ($\hat{\mathbf{\Sigma}}_{T,B}^{-1}$) is used. When any other weight matrix is used, the test statistic has a sample-specific limiting distribution, and critical values in such cases can be obtained via a simple simulation: (i) compute $\hat{\mathbf{R}}$ using $\hat{\mathbf{G}}_{T,S}$, $\hat{\mathbf{W}}_T$, and $\hat{\mathbf{\Sigma}}_{T,B}$; (ii) simulate $\mathbf{u}^{(k)} \sim$ iid $N(\mathbf{0}, \mathbf{I})$, for $k = 1, 2, \ldots, K$, where K is large; (iii) for each simulation, compute $J_{T,S}^{(k)} = \mathbf{u}^{(k)\prime}\hat{\mathbf{R}}'\hat{\mathbf{\Sigma}}_{T,B}^{1/2\prime}\hat{\mathbf{W}}_T\hat{\mathbf{\Sigma}}_{T,B}^{1/2}\hat{\mathbf{R}}\mathbf{u}^{(k)}$; (iv) use the sample $(1-\alpha)$ quantile of $\{J_{T,S}^{(k)}\}_{k=1}^{K}$ as the critical value for this test statistic.

The need for simulations to obtain critical values from the limiting distribution is nonstandard but is not uncommon; this arises in many other testing problems (see, e.g., Wolak 1989; White 2000; Andrews 2001). Given that $\mathbf{u}^{(k)}$ is standard Normal, and that $\hat{\mathbf{R}}$ need only be computed once, obtaining critical values for this test is simple and fast.

2.6 SMM Under Model Misspecification

All of the above results hold under the assumption that the copula model is correctly specified. In the event that a specification test rejects a model as misspecified, one is led directly to the question of whether these results, or extensions of them, hold for misspecified models.

In the literature on GMM, there are two common ways to define misspecification. Newey (1985) defines a form of "local" misspecification (where the degree of misspecification vanishes in the limit), and in that case it is simple to show that the asymptotic behavior of the SMM estimator does not change except for the mean of limit distribution. Hall and Inoue (2003) consider "nonlocal" misspecification. Formally, a model is said to be misspecified if there is no value of $\boldsymbol{\theta} \in \mathbf{\Theta}$, which satisfies $\mathbf{g}_0(\boldsymbol{\theta}) = \mathbf{0}$. As Hall and Inoue (2003) note, misspecification is only a concern when the model is overidentified, and so in this section we assume $m > p$. The absence of a parameter that satisfies the population moment conditions means that we must instead consider a "pseudo-true" parameter:

Definition 1. The pseudo-true parameter is $\boldsymbol{\theta}_*(\mathbf{W}_0) \equiv \arg\min_{\boldsymbol{\theta}\in\mathbf{\Theta}} \mathbf{g}_0'(\boldsymbol{\theta})\mathbf{W}_0\mathbf{g}_0(\boldsymbol{\theta})$.

The true parameter, $\boldsymbol{\theta}_0$, when it exists, is determined only by the population moment condition $\mathbf{g}_0(\boldsymbol{\theta}_0) = \mathbf{0}$, while the pseudo-true parameter also depends on the weight matrix $\mathbf{W}_0$, and thus is denoted $\boldsymbol{\theta}_*(\mathbf{W}_0)$. With the additional assumptions below, consistency of the SMM estimator under misspecification can be proven. The following proposition extends the results for GMM under misspecification in Hall (2000) and Hall and Inoue (2003), as it is established under the discontinuity of the moment functions.

Assumption 5. (i) (Nonlocal misspecification) $\|\mathbf{g}_0(\boldsymbol{\theta})\| > 0$ for all $\boldsymbol{\theta} \in \mathbf{\Theta}$.

(ii) (Identification) There exists $\boldsymbol{\theta}_*(\mathbf{W}_0) \in \mathbf{\Theta}$ such that $\mathbf{g}_0(\boldsymbol{\theta}_*(\mathbf{W}_0))'\mathbf{W}_0\mathbf{g}_0(\boldsymbol{\theta}_*(\mathbf{W}_0)) < \mathbf{g}_0(\boldsymbol{\theta})'\mathbf{W}_0\mathbf{g}_0(\boldsymbol{\theta})$ for all $\boldsymbol{\theta} \in \mathbf{\Theta}\backslash\{\boldsymbol{\theta}_*(\mathbf{W}_0)\}$.

Proposition 5. Suppose that Assumptions 1, 2, 3(ii)–3(iv), and 5 hold. Then, $\hat{\boldsymbol{\theta}}_{T,S} \xrightarrow{p} \boldsymbol{\theta}_*(\mathbf{W}_0)$ as $T, S \to \infty$.

While consistency of $\hat{\boldsymbol{\theta}}_{T,S}$ under misspecification is easily obtained, establishing the limit distribution of $\hat{\boldsymbol{\theta}}_{T,S}$ is not straightforward. Hall and Inoue (2003) showed that the limit distribution of GMM (with smooth, differentiable moment functions) depends on the limit distribution of the weight matrix, not merely the probability limit of the weight matrix. In SMM applications, it is possible to show that the limit distribution will additionally depend on the limit distribution of the numerical derivative matrix $\hat{\mathbf{G}}_{T,S}$. Furthermore, under misspecification one needs an alternative approach to establish stochastic equicontinuity of the objective function, which is required to obtain the limit distribution of the estimator. We leave the study of this limit distribution for future research.

3. SIMULATION STUDY

In this section, we present a study of the finite sample properties of the SMM estimator studied in the previous section. We consider two widely known copula models, the Clayton and the Gaussian (or Normal) copulas (see Nelsen 2006 for discussion) and the "factor copula" proposed in Oh and Patton (2011) outlined below. A closed-form likelihood is available for the first two copulas, while the third copula requires a numerical integration step to obtain the likelihood (details on this are presented in the supplemental appendix, available online). In all cases, we contrast the finite-sample properties of the MLE with the SMM estimator. The first two copulas also have closed-form cumulative distribution functions, and so quantile dependence [defined in Equation (3)] is also known in closed form. For the Clayton copula, we have Kendall's tau in closed form ($\tau = \kappa/(2+\kappa)$) but not Spearman's rank correlation (see Nelsen 2006). For the Normal copula, we have both Spearman's rank correlation in closed form [$\rho_S = 6/\pi \arcsin(\rho/2)$] and Kendall's tau [$\tau = 2/\pi \arcsin(\rho)$] (see Demarta and McNeil 2005; Nelsen 2006). This allows us to also compare GMM with SMM for these copulas, to quantify the loss in accuracy from having to resort to simulations.

The factor copula we consider is based on the following structure:

$$\begin{aligned} \text{Let} \quad & X_i = Z + \varepsilon_i, \quad i = 1, 2, \ldots, N, \\ \text{where} \quad & Z \sim \text{Skew}\ t(0, \sigma^2, \nu^{-1}, \lambda),\ \varepsilon_i \sim \text{iid}\ t(\nu^{-1}), \\ & \text{and} \quad \varepsilon_i \perp\!\!\!\perp Z\ \forall i \\ [X_1, \ldots, X_N]' \equiv \mathbf{X} \sim \mathbf{F}_x = & \ \mathbf{C}(G_x, \ldots, G_x), \end{aligned} \tag{16}$$

where we use the skewed t distribution of Hansen (1994). We use the copula of $\mathbf{X}$ implied by the above structure as our "factor copula" model, and it is parameterized by $(\sigma^2, \nu^{-1}, \lambda)$. For the factor copula, we have none of the above dependence measures in closed form. For the simulation, we set the parameters to generate rank correlation of around 1/2, and so set the Clayton copula parameter to 1, the Gaussian copula parameter to 1/2, and the factor copula parameters to $\sigma^2 = 1$, $\nu^{-1} = 1/4$, and $\lambda = -1/2$.

We consider two different scenarios for the marginal distributions of the variables of interest. In the first case, we assume that the data are iid with standard Normal marginal distributions, meaning that the only parameters that need to be estimated are those of the copula. This case is contrasted with a scenario where the marginal distributions of the variables are assumed to follow an AR(1)-GARCH(1,1) process:

$$\begin{aligned} Y_{it} &= \phi_0 + \phi_1 Y_{i,t-1} + \sigma_{it}\eta_{it}, \quad t = 1, 2, \ldots, T, \\ \sigma_{it}^2 &= \omega + \beta\sigma_{i,t-1}^2 + \alpha\sigma_{i,t-1}^2\eta_{i,t-1}^2, \\ \boldsymbol{\eta}_t &\equiv [\eta_{1t}, \ldots, \eta_{Nt}]' \sim \text{iid}\ \mathbf{F}_\eta = \mathbf{C}(\Phi, \Phi, \ldots, \Phi), \end{aligned} \tag{17}$$

where Φ is the standard Normal distribution function and $\mathbf{C}$ can be Clayton, Gaussian, or the factor copula implied by Equation (16). We set the parameters of the marginal distributions as $[\phi_0, \phi_1, \omega, \beta, \alpha] = [0.01, 0.05, 0.05, 0.85, 0.10]$, which broadly matches the values of these parameters when estimated using daily equity return data. In this scenario, the parameters of the models for the conditional mean and variance are estimated, and then the estimated standardized residuals are obtained: $\hat{\eta}_{it} = \hat{\sigma}_{it}^{-1}(Y_{it} - \hat{\phi}_0 - \hat{\phi}_1 Y_{i,t-1})$. These residuals are used in a second stage to estimate the copula parameters. In all cases, we consider a time series of length $T = 1000$, corresponding to approximately 4 years of daily return data, and we use $S = 25 \times T$ simulations in the computation of the dependence measures to be matched in the SMM optimization. We use five dependence measures in estimation: Spearman's rank correlation, and the 0.05, 0.10, 0.90, 0.95 quantile dependence measures, averaged across pairs of assets. We repeat each scenario 100 times; and in the results below, we use the identity weight matrix for estimation. (Corresponding results based on the efficient weight matrix, $\hat{\mathbf{W}}_T = \hat{\boldsymbol{\Sigma}}_{T,B}^{-1}$, are comparable, and available in the supplemental appendix, available online.) We also report the computation times (per simulation) for each estimation.

Table 1 reveals that for all three dimensions ($N = 2$, 3, and 10) and for all three copula models, the estimated parameters are centered on the true values, with the average estimated bias being small relative to the standard deviation. Looking across the dimension size, we see that the copula model parameters are almost always more precisely estimated as the dimension grows. This is intuitive, given the exchangeable nature of all three models.

Comparing the SMM estimator with the ML estimator, we see that the SMM estimators suffer a loss in efficiency of around 50% for $N = 2$ and around 20% for $N = 10$. The loss is greatest for the ν^{-1} parameter of the factor copula, and is moderate and similar for the remaining parameters. Some loss is of course expected, and this simulation indicates that the loss is moderate overall. Comparing the SMM estimator to the GMM estimator, we find a loss in efficiency of zero to 3%, indicating only a slight loss in accuracy from having to estimate the population moment function via simulation. The simulation results in Table 2, where the copula parameters are estimated after the estimation of AR-GARCH models for the marginal distributions in a separate stage, are very similar to the case when no marginal distribution parameters are required to be estimated, consistent with Proposition 2.

In Table 3, we present the finite-sample coverage probabilities of 95% confidence intervals based on the asymptotic normality result from Proposition 2 and the asymptotic covariance matrix estimator presented in Proposition 3. As shown in that proposition, a critical input to the asymptotic covariance matrix estimator is the step size used in computing the numerical derivative matrix $\hat{\mathbf{G}}_{T,S}$. This step size, $\varepsilon_{T,S}$, must go to zero, but at a slower rate than $1/\sqrt{T}$. Ignoring constants, our simulation sample size of $T = 1000$ suggests setting $\varepsilon_{T,S} > 0.03$, which is much larger than standard step sizes used in computing numerical derivatives. We study the impact of the choice of step size by considering a range of values from 0.0001 to 0.1. Table 3 shows that when the step size is set to 0.01 or 0.1, the finite-sample coverage rates are close to their nominal levels. However, if the step size is chosen too small (0.001 or smaller), then the coverage rates are much lower than nominal levels.

Table 3 also presents the results of a study of the rejection rates for the test of overidentifying restrictions presented in Proposition 4. Given that we consider $\mathbf{W} = \mathbf{I}$ in this table, the test statistic has a nonstandard distribution, and we use $K = 10{,}000$ simulations to obtain critical values. The rejection rates are close to their nominal levels 95% for the all three copula models.

Table 1. Simulation results for iid data

	Clayton copula				*Normal copula*			*Factor copula*					
								MLE			SMM		
	MLE	GMM	SMM	SMM*	MLE	GMM	SMM						
	κ	κ	κ	κ	ρ	ρ	ρ	σ^2	ν^{-1}	λ	σ^2	ν^{-1}	λ
True	1.00	1.00	1.00	1.00	0.5	0.5	0.5	1.00	0.25	−0.50	1.00	0.25	−0.50
							$N = 2$						
Bias	0.001	−0.014	−0.006	−0.004	0.004	−0.001	−0.001	0.026	−0.002	−0.026	0.016	−0.012	−0.089
St dev	0.085	0.119	0.122	0.110	0.024	0.034	0.034	0.135	0.045	0.144	0.152	0.123	0.199
Median	1.011	0.982	0.991	0.998	0.503	0.497	0.498	1.027	0.251	−0.500	0.985	0.243	−0.548
90-10%	0.216	0.308	0.293	0.294	0.063	0.086	0.087	0.331	0.118	0.355	0.374	0.363	0.540
Time	0.061	0.060	512	49.5	0.021	0.292	0.483		254			103	
							$N = 3$						
Bias	0.015	0.008	0.006	0.008	0.003	−0.004	−0.005	0.014	−0.001	−0.012	0.032	−0.002	−0.057
St dev	0.061	0.090	0.092	0.091	0.020	0.025	0.026	0.120	0.028	0.109	0.124	0.111	0.157
Median	1.013	1.003	0.999	0.998	0.503	0.497	0.499	1.001	0.250	−0.502	1.031	0.256	−0.542
90-10%	0.155	0.226	0.219	0.216	0.049	0.064	0.068	0.297	0.073	0.222	0.297	0.293	0.395
Time	0.113	0.091	1360	56.2	0.023	0.293	0.815		263			136	
							$N = 10$						
Bias	0.008	0.007	0.008	0.004	0.003	−0.002	−0.002	0.011	0.000	0.006	0.026	0.001	−0.011
St dev	0.050	0.068	0.066	0.059	0.014	0.017	0.017	0.092	0.016	0.063	0.093	0.070	0.082
Median	1.005	1.002	1.005	0.999	0.504	0.498	0.499	1.005	0.248	−0.494	1.013	0.255	−0.508
90-10%	0.132	0.198	0.177	0.152	0.035	0.039	0.045	0.240	0.034	0.166	0.248	0.186	0.168
Time	0.409	0.998	22289	170	0.475	0.331	3.140		396			341	

NOTE: This table presents the results from 100 simulations of the Clayton copula, the Normal copula, and a factor copula. In the SMM and GMM estimation, all three copulas use five dependence measures, including four quantile dependence measures ($q = 0.05, 0.10, 0.90, 0, 95$). The Normal and factor copulas also use Spearman's rank correlation, while the Clayton copula uses either Kendall's (GMM and SMM) or Spearman's (SMM*) rank correlation. The marginal distributions of the data are assumed to be iid $N(0, 1)$. Problems of dimension $N = 2$, 3, and 10 are considered, the sample size is $T = 1000$, and the number of simulations used is $S = 25 \times T$. The first row of each panel presents the average difference between the estimated parameter and its true value. The second row presents the standard deviation of the estimated parameters. The third and fourth rows present the median and the difference between the 90th and 10th percentiles of the distribution of estimated parameters. The last row in each panel presents the average time in seconds to compute the estimator.

Table 2. Simulation results for AR-GARCH data

	Clayton copula				*Normal copula*			*Factor copula*					
								MLE			SMM		
	MLE	GMM	SMM	SMM*	MLE	GMM	SMM						
	κ	κ	κ	κ	ρ	ρ	ρ	σ^2	ν^{-1}	λ	σ^2	ν^{-1}	λ
True	1.00	1.00	1.00	1.00	0.5	0.5	0.5	1.00	0.25	−0.50	1.00	0.25	−0.50
							$N = 2$						
Bias	−0.005	−0.029	−0.028	−0.020	0.003	−0.001	−0.001	0.021	−0.009	−0.029	0.015	−0.012	−0.073
St dev	0.087	0.124	0.124	0.108	0.024	0.035	0.036	0.137	0.046	0.150	0.155	0.121	0.188
Median	0.998	0.977	0.975	0.982	0.503	0.497	0.499	1.021	0.245	−0.503	0.995	0.235	−0.558
90-10%	0.228	0.327	0.340	0.267	0.061	0.084	0.090	0.343	0.118	0.382	0.411	0.346	0.509
Time	0.026	0.059	525	52	0.030	0.299	0.505		234			95	
							$N = 3$						
Bias	0.006	−0.007	0.002	−0.008	0.003	−0.005	−0.006	0.007	−0.007	−0.011	0.013	−0.020	−0.052
St dev	0.060	0.087	0.088	0.080	0.020	0.026	0.026	0.118	0.028	0.110	0.121	0.106	0.148
Median	1.005	0.991	0.994	0.981	0.502	0.497	0.499	0.997	0.243	−0.502	1.005	0.238	−0.521
90-10%	0.145	0.205	0.213	0.195	0.050	0.065	0.068	0.315	0.074	0.224	0.311	0.297	0.357
Time	0.127	0.108	1577	73	0.022	0.288	1.009		232			119	
							$N = 10$						
Bias	−0.004	−0.002	−0.004	−0.003	0.002	−0.003	−0.004	0.005	−0.006	0.008	−0.004	−0.016	−0.012
St dev	0.049	0.067	0.064	0.059	0.014	0.016	0.017	0.091	0.015	0.063	0.085	0.079	0.071
Median	0.995	0.996	0.987	0.988	0.503	0.497	0.497	1.002	0.243	−0.493	0.990	0.238	−0.508
90-10%	0.134	0.179	0.170	0.152	0.034	0.041	0.045	0.240	0.037	0.169	0.210	0.209	0.165
Time	0.292	1.059	24549	171	1.099	0.392	3.437		430			309	

NOTE: This table presents the results from 100 simulations of the Clayton copula, the Normal copula, and a factor copula. In the SMM and GMM estimation, all three copulas use five dependence measures, including four quantile dependence measures ($q = 0.05, 0.10, 0.90, 0, 95$). The Normal and factor copulas also use Spearman's rank correlation, while the Clayton copula uses either Kendall's (GMM and SMM) or Spearman's (SMM*) rank correlation. The marginal distributions of the data are assumed to follow AR(1)-GARCH(1,1) processes as described in Section 3. Problems of dimension $N = 2$, 3, and 10 are considered, the sample size is $T = 1000$, and the number of simulations used is $S = 25 \times T$. The first row of each panel presents the average difference between the estimated parameter and its true value. The second row presents the standard deviation of the estimated parameters. The third and fourth rows present the median and the difference between the 90th and 10th percentiles of the distribution of estimated parameters. The last row in each panel presents the average time in seconds to compute the estimator.

Table 3. Simulation results on coverage rates

	Clayton copula		Normal copula		Factor copula			
	κ	J	ρ	J	σ^2	ν^{-1}	λ	J
				$N = 2$				
$\varepsilon_{T,S}$								
0.1	91	98	94	98	94	100	95	98
0.01	46	99	92	98	94	99	96	100
0.001	2	99	76	98	76	79	74	99
0.0001	1	99	21	98	54	75	57	97
				$N = 3$				
$\varepsilon_{T,S}$								
0.1	97	99	89	97	99	100	96	99
0.01	63	98	88	97	99	96	95	100
0.001	11	98	83	98	92	84	93	100
0.0001	2	100	38	99	57	70	61	99
				$N = 10$				
$\varepsilon_{T,S}$								
0.1	96	99	87	97	97	97	95	98
0.01	88	99	87	96	96	97	97	97
0.001	18	100	87	98	97	95	88	97
0.0001	0	98	71	97	73	85	81	98

NOTE: This table presents the results from 100 simulations of the Clayton copula, the Normal copula, and a factor copula, all estimated by SMM. The marginal distributions of the data are assumed to follow AR(1)-GARCH(1,1) processes as described in Section 3. Problems of dimension $N = 2$, 3, and 10 are considered, the sample size is $T = 1000$, and the number of simulations used is $S = 25 \times T$. The rows of each panel contain the step size, $\varepsilon_{T,S}$, used in computing the matrix of numerical derivatives, $\hat{\mathbf{G}}_{T,S}$. The numbers in column κ, ρ, σ^2, ν^{-1}, and λ present the percentage of simulations for which the 95% confidence interval based on the estimated covariance matrix contained the true parameter. The numbers in column J present the percentage of simulations for which the test statistic of overidentifying restrictions test described in Section 2 was smaller than its computed critical value under 95% confidence level.

We finally consider the properties of the estimator under model misspecification. In Table 4, we consider two scenarios: one where the true copula is Clayton but the model is Normal, and one where the true copula is Normal but the model is Clayton. The pseudo-true parameters for these two scenarios are not known in closed form, and we use a simulation of 10 million observations to estimate it. The pseudo-true parameters are reported in the top row of each panel of Table 4. Similar to the correctly specified case, we see here that the estimated parameters are centered on the pseudo-true values, with the average estimated bias being small relative to the standard deviation. These misspecified scenarios also provide some insight into the power of the specification test based on overidentifying restrictions. We find that for all three dimensions and for both iid and AR-GARCH data, the J-test rejected the null of correct specification across all 100 simulations, indicating this test has power to detect model misspecification.

These simulation results provide support for the proposed estimation method: for empirically realistic parameter values and sample size, the estimator is approximately unbiased, with estimated confidence intervals that have coverage close to their nominal level when the step size for the numerical derivative is chosen in line with our theoretical results, and the test for model misspecification has finite-sample rejection frequencies that are close to their nominal levels when the model is correctly specified, and has good power to reject misspecified models.

Table 4. Simulation results for misspecified models

	iid		AR-GARCH	
True copula Model	Clayton Normal	Normal Clayton	Clayton Normal	Normal Clayton
		$N = 2$		
Pseudo-true	0.542	0.599	0.543	0.588
Bias	−0.013	0.111	−0.007	0.046
St dev	0.050	0.173	0.035	0.120
Median	0.526	0.659	0.539	0.617
90-10%	0.130	0.433	0.091	0.265
Time	4	72	1	70
J-test prob.	0	0	0	0
		$N = 3$		
Pseudo-true	0.543	0.599	0.542	0.607
Bias	0.003	0.077	−0.002	0.006
St dev	0.039	0.164	0.027	0.088
Median	0.544	0.629	0.540	0.609
90–10%	0.107	0.432	0.072	0.198
Time	5	90	1	86
J-test prob.	0	0	0	0
		$N = 10$		
Pseudo-true	0.544	0.602	0.544	0.603
Bias	0.001	0.059	−0.001	0.047
St dev	0.033	0.118	0.016	0.116
Median	0.546	0.622	0.540	0.618
90-10%	0.086	0.307	0.043	0.314
Time	20	206	4	207
J-test prob.	0	0	0	0

NOTE: This table presents the results from 100 simulations when the true copula and the model are different (i.e., the model is misspecified). The parameters of the copula models are estimated using SMM based on rank correlation and four quantile dependence measures ($q = 0.05, 0.10, 0.90, 0, 95$). The marginal distributions of the data are assumed to be either iid $N(0, 1)$ or AR(1)-GARCH(1,1) processes as described in Section 3. Problems of dimension $N = 2$, 3, and 10 are considered, the sample size is $T = 1000$, and the number of simulations used is $S = 25 \times T$. The pseudo-true parameter is estimated using 10 million observations. The last row in each panel presents the proportion of tests of overidentifying restrictions that are smaller than the 95% critical value.

4. APPLICATION TO THE DEPENDENCE BETWEEN FINANCIAL FIRMS

This section considers models for the dependence between seven large financial firms. We use daily stock return data over the period January 2001 to December 2010, a total of $T = 2515$ trade days, on Bank of America, Bank of New York, Citigroup, Goldman Sachs, J.P. Morgan, Morgan Stanley, and Wells Fargo. Summary statistics for these returns are presented in Table S4 of the supplemental appendix (available online), and indicate that all series are positively skewed and leptokurtotic, with kurtosis ranging from 16.0 (J.P. Morgan) to 119.8 (Morgan Stanley).

To capture the impact of time-varying conditional means and variances in each of these series, we estimate the following autoregressive, conditionally heteroscedastic models:

$$r_{it} = \phi_{0i} + \phi_{1i} r_{i,t-1} + \phi_{2i} r_{m,t-1} + \varepsilon_{it}, \quad \varepsilon_{it} = \sigma_{it}\eta_{it},$$

where

$$\sigma_{it}^2 = \omega_i + \beta_i \sigma_{i,t-1}^2 + \alpha_{1i}\varepsilon_{i,t-1}^2 + \gamma_{1i}\varepsilon_{i,t-1}^2 \mathbf{1}_{[\varepsilon_{i,t-1} \le 0]} + \alpha_{2i}\varepsilon_{m,t-1}^2 + \gamma_{2i}\varepsilon_{m,t-1}^2 \mathbf{1}_{[\varepsilon_{m,t-1} \le 0]}, \quad (18)$$

where r_{it} is the return on one of these seven firms and r_{mt} is the return on the S&P 500 index. We include the lagged

Table 5. Sample dependence statistics

	Bank of America	Bank of New York	Citi group	Goldman Sachs	J.P. Morgan	Morgan Stanley	Wells Fargo
			Panel A: Correlation estimates				
BoA		0.586	0.691	0.556	0.705	0.602	0.701
BoNY	0.551		0.574	0.578	0.658	0.592	0.595
Citi	0.685	0.558		0.608	0.684	0.649	0.626
Goldman	0.564	0.565	0.609		0.655	0.759	0.548
JPM	0.713	0.633	0.694	0.666		0.667	0.683
Morgan S	0.604	0.587	0.650	0.774	0.676		0.578
Wells F	0.715	0.593	0.636	0.554	0.704	0.587	
			Panel B: Quantile dependence estimates				
BoA		0.219	0.239	0.219	0.398	0.298	0.358
BoNY	−0.048		0.179	0.199	0.159	0.219	0.199
Citi	−0.045	−0.004		0.199	0.318	0.219	0.199
Goldman	−0.068	0.000	0.032		0.239	0.378	0.199
JPM	−0.024	−0.056	−0.012	0.012		0.239	0.358
Morgan S	−0.060	−0.020	−0.064	−0.036	−0.008		0.219
Wells F	0.020	−0.052	0.044	−0.028	0.024	0.000	

NOTE: This table presents measures of dependence between the seven financial firms under analysis. The upper panel presents Spearman's rank correlation (upper triangle) and linear correlation (lower triangle), and the lower panel presents the difference between the 10% quantile dependence measures (lower triangle) and average 1% upper and lower quantile dependence (upper triangle). All dependence measures are computed using the standardized residuals from the models for the conditional mean and variance.

market index return in both the mean and variance specifications to capture any influence of lagged information in the model for a given stock, and in the model for the market index itself we set $\phi_1 = \alpha_1 = \gamma_1 = 0$. Estimated parameters from these models are presented in Table S5 of the supplemental appendix (available online), and are consistent with the values found in the empirical finance literature (see, e.g., Bollerslev, Engle, and Nelson 1994). From these models, we obtain the estimated standardized residuals, $\hat{\eta}_{it}$, which are used in the estimation of the dependence structure.

In Table 5, we present measures of dependence between these seven firms. The upper panel reveals that rank correlation between their standardized residuals is 0.63 on average, and ranges from 0.55 to 0.76. The lower panel of Table 5 presents measures of dependence in the tails between these series. The upper triangle presents the average of the 1% and 99% quantile dependence measures presented in Equation (6), and we see substantial dependence here, with values ranging between 0.16 and 0.40. The lower triangle presents the difference between the 90% and 10% quantile dependence measures, as a gauge of the degree of asymmetry in the dependence structure. These differences are mostly negative (14 out of 21), indicating greater dependence during crashes than during booms.

Table 6 presents the estimation results for three different copula models of these series. The first model is the well-known Clayton copula, the second is the Normal copula, and the third is a "factor copula" as proposed by Oh and Patton (2011). The first copula allows for lower tail dependence, but imposes that upper tail dependence is zero. The second copula implies zero tail dependence in both directions. The third copula allows for tail dependence in both tails, and allows the degree of dependence to differ across positive and negative realizations.

Table 6. Estimation results for daily returns on seven stocks

	Clayton copula			*Normal copula*			*Factor copula*		
	MLE κ	SMM κ	SMM-opt κ	MLE ρ	SMM ρ	SMM-opt ρ	MLE $\sigma^2, \nu^{-1}, \lambda$	SMM $\sigma^2, \nu^{-1}, \lambda$	SMM-opt $\sigma^2, \nu^{-1}, \lambda$
Estimate	0.907	1.274	1.346	0.650	0.682	0.659	1.995	2.019	1.955
Std err	0.028	0.048	0.037	0.007	0.010	0.008	0.020	0.077	0.069
Estimate	–	–	–	–	–	–	0.159	0.088	0.115
Std err	–	–	–	–	–	–	0.010	0.034	0.033
Estimate	–	–	–	–	–	–	−0.021	−0.015	−0.013
Std err	–	–	–	–	–	–	0.032	0.035	0.034
$Q_{SMM} \times 100$	–	19.820	19.872	–	0.240	0.719	–	0.040	0.187
J_{pval}	–	0.000	0.000	–	0.043	0.001	–	0.139	0.096
Time	0.7	344	360	0.5	6	6	1734	801	858

NOTE: This table presents estimation results for various copula models applied to seven daily stock returns in the financial sector over the period January 2001 to December 2010. Estimates and asymptotic standard errors for the copula model parameters are presented, as well as the value of the SMM objective function at the estimated parameters and the p-value of the overidentifying restriction test. Estimates labeled "SMM" are estimated using the identity weight matrix; estimates labeled "SMM-opt" are estimated using the efficient weight matrix.

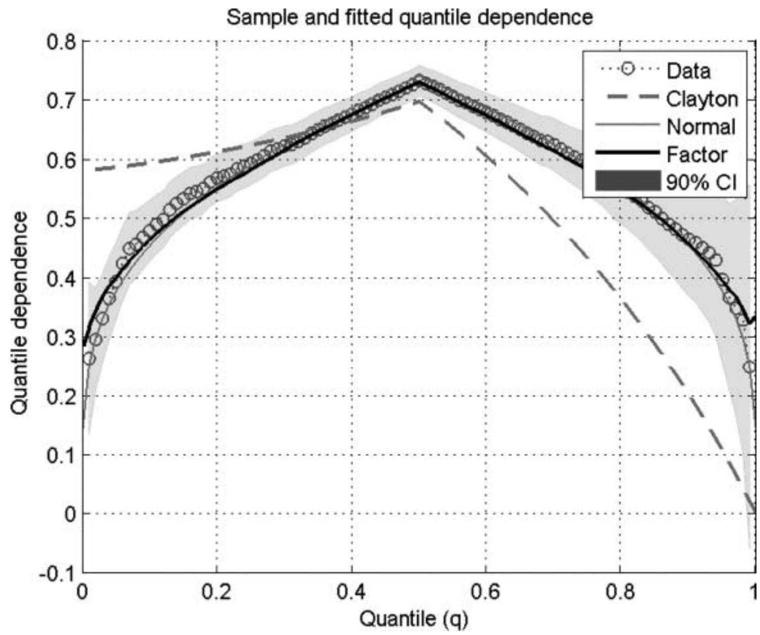

Figure 1. This figure plots the probability of both variables being less than their q quantile ($q < 0.5$) or greater than the q quantile ($q > 0.5$). For the data, this is averaged across all pairs, and a bootstrap 90% (pointwise) confidence interval is presented.

For all three copulas, we implement the SMM estimator proposed in Section 2, with the identity weight matrix and the efficient weight matrix, using five dependence measures: Spearman's rank correlation, and the 0.05, 0.10, 0.90, 0.95 quantile dependence measures, averaged across pairs of assets. We also implement the MLE for comparison. The value of the SMM objective function at the estimated parameters is presented for each model, along with the p-value from a test of the overidentifying restrictions based on Proposition 4. We use Proposition 3 to compute the standard errors, with $B = 1000$ bootstraps used to estimate $\mathbf{\Sigma}_{T,S}$, and $\varepsilon_{T,S} = 0.1$ used as the step size to compute $\hat{\mathbf{G}}_{T,S}$.

The parameter estimates for the Normal and factor copula models are similar for ML and SMM, while they are quite different for the Clayton copula. This may be explained by the results of the test of overidentifying restrictions: the Clayton copula is strongly rejected (with a p-value of less than 0.001 for both choices of weight matrix), while the Normal is less strongly rejected (p-values of 0.043 and 0.001). The factor copula is not rejected using this test for either choice of weight matrix. The improvement in fit from the factor copula appears to come from its ability to capture tail dependence: the parameter that governs tail dependence (ν^{-1}) is significantly greater than zero, while the parameter that governs asymmetric dependence (λ) is not significantly different from zero.

Given that our sample period spans the financial crisis, one may wonder whether the copula is constant throughout the period. To investigate this, we implement the copula structural break test proposed by Rémillard (2010), using 1000 simulations for the "multiplier" method, and find a p-value of 0.001, indicating strong evidence of a change in the copula over this period. Running this test on the last two years of our sample period (January 2009 to December 2010) results in a p-value of 0.191, indicating no evidence of a change in the copula over this subperiod. In Table S6 of the supplemental appendix (available online), we present results from the estimation of these copula models for this subperiod. The estimated parameters all indicate a slight increase in dependence relative to the full sample estimates, including an increase in the degree of tail dependence between these firms. The results of the specification tests for this subsample are very similar to the full sample results.

Figure 1 sheds some further light on the relative performance of these copula models, over the full sample. This figure compares the empirical quantile dependence function with those implied by the three copula models. An iid bootstrap with $B = 1000$ replications is used to construct pointwise confidence intervals for the sample quantile dependence estimates. We see here that the Clayton copula is "too asymmetric" relative to the data, while both the Normal and the factor copula models appear to provide a reasonable fit.

5. CONCLUSION

This article presents the asymptotic properties of a new simulation-based estimator of the parameters of a copula model, which matches measures of rank dependence implied by the model to those observed in the data. The estimation method shares features with the SMM (see, e.g., McFadden 1989; Newey and McFadden 1994); however, the use of rank dependence measures as "moments" means that existing results on SMM cannot be used. We extend results on SMM estimators using empirical process theory for copula estimation (see Fermanian, Radulović, and Wegkamp 2004; Chen and Fan 2006; Rémillard 2010) to show the consistency and asymptotic normality of SMM-type estimators of copula models. We also provide a method for obtaining a consistent estimate of the asymptotic covariance matrix and a test of the overidentifying restrictions. Our results apply to both iid and time series data, and an extensive simulation study verifies that the asymptotic results provide a good approximation in finite samples. We illustrate the results with an application to a model of the dependence between the equity returns on seven financial firms.

APPENDIX: SKETCH OF PROOFS

Detailed proofs are available in the supplemental appendix (available online) to this article.

Proof of Proposition 1. First note that: (a) $Q_0(\boldsymbol{\theta})$ is uniquely minimized at $\boldsymbol{\theta}_0$ by Assumption 3(i) and positive definite $\mathbf{W}_0$ of Assumption 3(iv), (b) $\boldsymbol{\Theta}$ is compact by Assumption 3(ii), (c) $Q_0(\boldsymbol{\theta})$ consists of linear combinations of rank correlations and quantile dependence measures that are functions of pairwise copula functions, so $Q_0(\boldsymbol{\theta})$ is continuous by Assumption 3(iii). The main part of the proof requires establishing that $Q_{T,S}$ uniformly converges in probability to Q_0, which we show using five lemmas in the supplemental appendix (available online): pointwise convergence of $\mathbf{g}_{T,S}(\boldsymbol{\theta})$ to $\mathbf{g}_0(\boldsymbol{\theta})$ and stochastic Lipschitz continuity of $\mathbf{g}_{T,S}(\boldsymbol{\theta})$ are shown using results from Fermanian, Wegkamp, and Radulović (2004) and Rémillard (2010) combined with Assumption 3(iii). This is sufficient for the stochastic equicontinuity of $\mathbf{g}_{T,S}$ and for the uniform convergence in probability of $\mathbf{g}_{T,S}$ to $\mathbf{g}_0$by Lemmas 2.8 and 2.9 of Newey and McFadden (1994). Using the triangle and Cauchy–Schwarz inequalities, this implies that $Q_{T,S}$ uniformly converges in probability to Q_0. We have thus verified that the conditions of Theorem 2.1 of Newey and McFadden (1994) hold, and we have $\hat{\boldsymbol{\theta}} \xrightarrow{p} \boldsymbol{\theta}_0$ as claimed. □

Proof of Proposition 2. We prove this proposition by verifying the five conditions of Theorem 7.2 of Newey and McFadden (1994) for our problem: (i) $\mathbf{g}_0(\boldsymbol{\theta}_0) = 0$ by construction of $\mathbf{g}_0(\boldsymbol{\theta}) = \mathbf{m}(\boldsymbol{\theta}_0) - \mathbf{m}(\boldsymbol{\theta})$. (ii) $\mathbf{g}_0(\boldsymbol{\theta})$ is differentiable at $\boldsymbol{\theta}_0$ with derivative $\mathbf{G}_0$ such that $\mathbf{G}_0'\mathbf{W}_0\mathbf{G}_0$ is nonsingular by Assumption 4(ii). (iii) $\boldsymbol{\theta}_0$ is an interior point of $\boldsymbol{\Theta}$ by Assumption 4(i). (iv) This part requires showing the asymptotic normality of $\sqrt{T}\mathbf{g}_{T,S}(\boldsymbol{\theta}_0)$. We will present the result only for $S/T \to k \in (0, \infty)$. The results for the cases that $S/T \to 0$ or $S/T \to \infty$ are similar. In Lemma 6 of the supplemental appendix (available online), we show that $\sqrt{T}(\hat{\mathbf{m}}_T - \mathbf{m}_0(\boldsymbol{\theta}_0)) \xrightarrow{d} N(0, \boldsymbol{\Sigma}_0)$ as $T \to \infty$ and $\sqrt{S}(\tilde{\mathbf{m}}_S(\boldsymbol{\theta}_0) - \mathbf{m}_0(\boldsymbol{\theta}_0)) \xrightarrow{d} N(0, \boldsymbol{\Sigma}_0)$ as $S \to \infty$ using Theorem 3 and Theorem 6 of Fermanian, Radulović, and Wegkamp (2004) and Corollary 1, Proposition 2, and Proposition 4 of Rémillard (2010). This implies that

$$\sqrt{T}\mathbf{g}_{T,S}(\boldsymbol{\theta}_0) = \underbrace{\sqrt{T}(\hat{\mathbf{m}}_T - \mathbf{m}_0(\boldsymbol{\theta}_0))}_{\xrightarrow{d} N(0,\boldsymbol{\Sigma}_0)} - \underbrace{\sqrt{\frac{T}{S}}}_{\to 1/\sqrt{k}} \underbrace{\sqrt{S}(\tilde{\mathbf{m}}_S(\boldsymbol{\theta}_0) - \mathbf{m}_0(\boldsymbol{\theta}_0))}_{\xrightarrow{d} N(0,\boldsymbol{\Sigma}_0)},$$

and so $\sqrt{T}\mathbf{g}_{T,S}(\boldsymbol{\theta}_0) \xrightarrow{d} N(\mathbf{0}, (1 + 1/k)\boldsymbol{\Sigma}_0)$ as $T, S \to \infty$. (v) This part requires showing that $\sup_{\|\theta-\theta_0\|<\delta} \sqrt{T}\|\boldsymbol{g}_{T,S}(\boldsymbol{\theta}) - \boldsymbol{g}_{T,S}(\boldsymbol{\theta}_0) - \boldsymbol{g}_0(\boldsymbol{\theta})\|/[1 + \sqrt{T}\|\boldsymbol{\theta} - \boldsymbol{\theta}_0\|] \xrightarrow{p} 0$. The main part of this proof involves showing the stochastic equicontinuity of $\mathbf{v}_{T,S}(\boldsymbol{\theta}) = \sqrt{T}[\boldsymbol{g}_{T,S}(\boldsymbol{\theta}) - \boldsymbol{g}_0(\boldsymbol{\theta})]$. This is shown in Lemma 7 of the supplemental appendix (available online) by showing that $\{\mathbf{g}_{\cdot,\cdot}(\boldsymbol{\theta}) : \theta \in \Theta\}$ is a type II class of functions in Andrews (1994), and then using that article's Theorem 1. □

Proof of Proposition 3. If $\boldsymbol{\mu}_t$ and $\boldsymbol{\sigma}_t$ are known constant, or if $\boldsymbol{\phi}_0$ is known, then the consistency of $\hat{\boldsymbol{\Sigma}}_{T,B}$ follows from Theorems 5 and 6 of Fermanian, Radulović, and Wegkamp (2004). When $\boldsymbol{\phi}_0$ is estimated, the result is obtained by combining the results in Fermanian, Radulović, and Wegkamp with those of Rémillard (2010): for simplicity, assume that only one dependence measure is used. Let $\hat{\rho}_{ij}$ and $\hat{\rho}_{ij}^{(b)}$ be the sample rank correlations constructed from the standardized residuals $\{\hat{\eta}_t^i, \hat{\eta}_t^j\}_{t=1}^T$ and from the bootstrap counterpart $\{\hat{\eta}_t^{(b)i}, \hat{\eta}_t^{(b)j}\}_{t=1}^T$. Also, define the corresponding estimates, $\ddot{\rho}_{ij}$ and $\ddot{\rho}_{ij}^{(b)}$, using the true innovations $\{\eta_t^i, \eta_t^j\}_{t=1}^T$ and the bootstrapped true innovations $\{\eta_t^{(b)i}, \eta_t^{(b)j}\}_{t=1}^T$ (where the same bootstrap time indices are used for both $\{\hat{\eta}_t^{(b)i}, \hat{\eta}_t^{(b)j}\}_{t=1}^T$ and $\{\eta_t^{(b)i}, \eta_t^{(b)j}\}_{t=1}^T$). Define true ρ as ρ_0. Theorem 5 of Fermanian, Radulović, and Wegkamp (2004) shows that

$$\sqrt{T}(\ddot{\rho}_{ij} - \rho_0) = \sqrt{T}(\ddot{\rho}_{ij}^{(b)} - \ddot{\rho}_{ij}) + o_p(1).$$

Corollary 1 and Proposition 4 of Rémillard (2010) show, under Assumption 2, that

$$\sqrt{T}(\hat{\rho}_{ij} - \ddot{\rho}_{ij}) = o_p(1) \quad \text{and} \quad \sqrt{T}\left(\hat{\rho}_{ij}^{(b)} - \ddot{\rho}_{ij}^{(b)}\right) = o_p(1).$$

Combining those three equations, we obtain

$$\sqrt{T}(\hat{\rho}_{ij} - \rho_0) = \sqrt{T}\left(\hat{\rho}_{ij}^{(b)} - \hat{\rho}_{ij}\right) + o_p(1), \text{ as } T, B \to \infty,$$

and so Equation (13) is a consistent estimator of $\boldsymbol{\Sigma}_0$. Consistency of the numerical derivatives $\hat{\mathbf{G}}_{T,S}$ can be established using a similar approach to Theorem 7.4 of Newey and McFadden (1994), and since $\hat{\mathbf{W}}_T \xrightarrow{p} \mathbf{W}_0$ by Assumption 3(iv), we thus have $\hat{\boldsymbol{\Omega}}_{T,S,B} \xrightarrow{p} \boldsymbol{\Omega}_0$. □

Proof of Proposition 4. We consider only the case where $S/T \to \infty$ or $S/T \to k > 0$. The case for $k = 0$ is analogous. A Taylor expansion of $\mathbf{g}_0(\hat{\boldsymbol{\theta}}_{T,S})$ around $\boldsymbol{\theta}_0$ yields

$$\sqrt{T}\mathbf{g}_0(\hat{\boldsymbol{\theta}}_{T,S}) = \sqrt{T}\mathbf{g}_0(\boldsymbol{\theta}_0) + \mathbf{G}_0 \cdot \sqrt{T}(\hat{\boldsymbol{\theta}}_{T,S} - \boldsymbol{\theta}_0) + o(\sqrt{T}\|\hat{\boldsymbol{\theta}}_{T,S} - \boldsymbol{\theta}_0\|),$$

and since $\mathbf{g}_0(\boldsymbol{\theta}_0) = 0$ and $\sqrt{T}\|\hat{\boldsymbol{\theta}}_{T,S} - \boldsymbol{\theta}_0\| = O_p(1)$

$$\sqrt{T}\mathbf{g}_0(\hat{\boldsymbol{\theta}}_{T,S}) = \mathbf{G}_0 \cdot \sqrt{T}(\hat{\boldsymbol{\theta}}_{T,S} - \boldsymbol{\theta}_0) + o_p(1). \tag{19}$$

Then, consider the following expansion of $\mathbf{g}_{T,S}(\hat{\boldsymbol{\theta}}_{T,S})$ around $\boldsymbol{\theta}_0$

$$\sqrt{T}\mathbf{g}_{T,S}(\hat{\boldsymbol{\theta}}_{T,S}) = \sqrt{T}\mathbf{g}_{T,S}(\boldsymbol{\theta}_0) + \hat{\mathbf{G}}_{T,S} \cdot \sqrt{T}(\hat{\boldsymbol{\theta}}_{T,S} - \boldsymbol{\theta}_0) + \mathbf{R}_{T,S}(\hat{\boldsymbol{\theta}}_{T,S}), \tag{20}$$

where the remaining term is captured by $\mathbf{R}_{T,S}(\hat{\boldsymbol{\theta}}_{T,S})$. Combining Equations (19) and (20), we obtain

$$\begin{aligned}&\sqrt{T}[\mathbf{g}_{T,S}(\hat{\boldsymbol{\theta}}_{T,S}) - \mathbf{g}_{T,S}(\boldsymbol{\theta}_0) - \mathbf{g}_0(\hat{\boldsymbol{\theta}}_{T,S})]\\ &\quad = (\hat{\mathbf{G}}_{T,S} - \mathbf{G}_0) \cdot \sqrt{T}(\hat{\boldsymbol{\theta}}_{T,S} - \boldsymbol{\theta}_0) + \mathbf{R}_{T,S}(\hat{\boldsymbol{\theta}}_{T,S}) + o_p(1).\end{aligned}$$

The stochastic equicontinuity of $\mathbf{v}_{T,S}(\boldsymbol{\theta}) = \sqrt{T}[\mathbf{g}_{T,S}(\boldsymbol{\theta}) - \mathbf{g}_0(\boldsymbol{\theta})]$ is established in the proof of Proposition 2, which implies (see proof of Proposition 2) that

$$\sqrt{T}[\mathbf{g}_{T,S}(\hat{\boldsymbol{\theta}}_{T,S}) - \mathbf{g}_{T,S}(\boldsymbol{\theta}_0) - \mathbf{g}_0(\hat{\boldsymbol{\theta}}_{T,S})] = o_p(1).$$

By Proposition 3, $\hat{\mathbf{G}}_{T,S} - \mathbf{G}_0 = o_p(1)$, which implies $\mathbf{R}_{T,S}(\hat{\boldsymbol{\theta}}_{T,S}) = o_p(1)$. Thus, we obtain the expansion of $\mathbf{g}_{T,S}(\hat{\boldsymbol{\theta}}_{T,S})$ around $\boldsymbol{\theta}_0$:

$$\sqrt{T}\mathbf{g}_{T,S}(\hat{\boldsymbol{\theta}}_{T,S}) = \sqrt{T}\mathbf{g}_{T,S}(\boldsymbol{\theta}_0) + \hat{\mathbf{G}}_{T,S} \cdot \sqrt{T}(\hat{\boldsymbol{\theta}}_{T,S} - \boldsymbol{\theta}_0) + o_p(1). \tag{21}$$

The remainder of the proof is the same as in standard GMM applications (see, e.g., Hall 2005). □

Proof of Proposition 5. Lemmas 1, 2, 3, and 4 are not affected by misspecification. Lemma 5(i) is replaced by Assumption 5(ii). Therefore, $\hat{\boldsymbol{\theta}}_{T,S} \xrightarrow{p} \boldsymbol{\theta}_*(\mathbf{W}_0)$. □

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