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Structural dynamic conditional correlation

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Structural Dynamic Conditional Correlation

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Structural Dynamic Conditional Correlation¹

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Abstract

In the literature of identification through autoregressive conditional heteroscedasticity, Weber (2008) developed the structural constant conditional correlation (SCCC) model. Besides determining linear simultaneous influences between several variables, this model considers interaction in the structural innovations. Even though this allows for common fundamental driving forces, these cannot explain time variation in correlations of observed variables, which still have to rely on causal transmission effects. In this context, the present paper extends the analysis to structural *dynamic* conditional correlation (SDCC). The additional flexibility is shown to make an important contribution in the estimation of empirical real-data examples.

Keywords: Simultaneity, Identification, EGARCH, DCC

JEL classification: C32, G10

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1 Introduction

Identifying structural models that feature simultaneous effects between several variables is one of the key tasks of econometrics. The conventional method solving identification problems in multivariate time series analysis works through parametric (zero) constraints, which allow recovering the structural model from the estimated reduced form. However, it is often difficult to justify these restrictions as they naturally imply a certain *a priori* determination of structure and direction of causalities.

For heteroscedastic series, a small strand of recent literature introduced methods that exploit non-constant variances for identifying simultaneous models "through heteroscedasticity" (see Rigobon 2003). A shift in the structural volatility, which yields more additional determining equations from the reduced-form covariance-matrix than unknown coefficients, describes the basic idea. Building on this logic, further research for example in Sentana and Fiorentini (2001), Rigobon (2002) and Weber (2007) proposed estimating ARCH-type processes as to coherently describe the necessary volatility movements.

While the previously mentioned models are successful in identifying simultaneous transmission effects, they neglect fundamental driving forces, which might equally underlie the development of all included variables. This problem has been recently addressed by Weber (2008), who introduced the so-called structural constant conditional correlation (SCCC) model. Thereby, he allowed for common third-party influences in a framework that upholds identifiability by restricting these influences to result in constant correlations of the innovations. However, in conventional reduced-form approaches, time variation has usually been found to be a common feature of correlations of observed financial variables. Just like the overall level of correlation, this time variation in conditional correlation can logically be triggered by direct spillovers or by common grounds in the disturbances. Consequently, the present paper allows for both of these sources instead of solely relying on the first one. This is achieved in a simultaneous equation system featuring structural *dynamic* conditional correlation (SDCC), extending the concept of Engle (2002) to the unobservable fundamental shocks. The proposed setup makes it possible to identify general causal structures, whose flexibility paves the way to realistic interpretations in terms of financial economics.

Building on the work of Weber (2008), the next section sets out the new methodological proceeding. Thereafter, section 3 discusses the usefulness of the econometric approach in a small empirical example with Dow Jones Industrial and Nasdaq Composite stock index returns. The last section concludes.

2 Methodology

To begin with, let us establish the key features of the SCCC model introduced by Weber (2008). Therein, contemporaneous transmission effects between the n endogenous variables contained in the vector y_t are specified as

$$Ay_t = \varepsilon_t . \quad (1)$$

Here, the coefficients representing instantaneous impacts are included in the $n \times n$ matrix A , in which the diagonal elements are normalised to one. ε_t is a n -dimensional vector of structural innovations with unrestricted correlation matrix.

Treating some notation in preparation for the ARCH modelling, denote the conditional variances of the elements in ε_t by

$$\text{Var}(\varepsilon_{jt}|I_{t-1}) = h_{jt}^2 \quad j = 1, \dots, n , \quad (2)$$

where I_{t-1} stands for the whole set of available information at time $t - 1$.

Then, stack the conditional variances in the vector $H_t = \left(h_{1t}^2 \quad \dots \quad h_{nt}^2 \right)'$.

At last, denote the standardised white noise residuals by

$$\tilde{\varepsilon}_{jt} = \varepsilon_{jt}/h_{jt} \quad j = 1, \dots, n . \quad (3)$$

Then, the multivariate EGARCH(1,1)-process, as suggested by Weber (2007), is given by

$$\log H_t = C + G \log H_{t-1} + D|\tilde{\varepsilon}_{t-1}| + F\tilde{\varepsilon}_{t-1} , \quad (4)$$

where C is a n -dimensional vector of constants and G , D and F are $n \times n$ coefficient matrices. The absolute value operation is to be applied element by element and provides the pure magnitude of shocks.² In addition, the signed $\tilde{\varepsilon}_t$ introduce asymmetric volatility effects.

While the conditional variances are treated in (4), Weber (2008) recovers the covariances by the constant conditional correlation assumption as

$$\text{Cov}(\varepsilon_{it}, \varepsilon_{jt}|I_{t-1}) = h_{ijt} = \rho_{ij}h_{it}h_{jt} \quad i \neq j , \quad (5)$$

²Note that in the original univariate formulation of Nelson (1991), the unsigned shock was corrected for its mean as in $(|\tilde{\varepsilon}_{t-1}| - E(|\tilde{\varepsilon}_{t-1}|))$. The present specification merges the term $-D \cdot E(|\tilde{\varepsilon}_{t-1}|)$ into the constants C , but is completely conformable to the original version. The advantage is that no distributional assumption has to be made for calculating the expectation.

where ρ_{ij} denotes the correlation between the i th and j th residual. At this point, the present paper introduces a considerably more flexible setup by adopting a DCC specification for the structural innovations: Building on Engle (2002), define the conditional correlation matrix R_t as

$$R_t = \text{diag}\{Q_t\}^{-1/2} Q_t \text{diag}\{Q_t\}^{-1/2} . \quad (6)$$

Therein, Q_t follows the process

$$Q_t = (1 - \alpha - \beta)\bar{Q} + \alpha\tilde{\varepsilon}_{t-1}\tilde{\varepsilon}'_{t-1} + \beta Q_{t-1} . \quad (7)$$

(7) corresponds to a standard GARCH(1,1) in that Q_t is driven by the cross product of the shocks and a persistence term. \bar{Q} denotes the unconditional covariance matrix of the standardised residuals $\tilde{\varepsilon}_t$. Although α and β are defined as scalars for parsimony, more comprehensive solutions are possible, see Engle (2002).

With R_t at hand, the conditional covariance-matrix Ω_t of the structural disturbances ε_t is defined as

$$\Omega_t = \text{diag}\{H_t\}^{1/2} R_t \text{diag}\{H_t\}^{1/2} . \quad (8)$$

Accounting for the discussion in Engle (2002) and given positive variances from the log-linearised EGARCH, Ω_t is assured to be positive definite. This property carries over to the conditional covariance-matrix of the reduced-form residuals $A^{-1}\varepsilon_t$

$$\Sigma_t = A^{-1}\Omega_t(A^{-1})' \quad (9)$$

due to its quadratic form.

Identifiability can now be discussed as in Weber (2008), without loss of generality focusing on the bivariate case. The structural variance process (4) contains two parameters in C and four each in G , D and F . Together with the two parameters from the structural matrix A and one each from α , β and \bar{Q} , the sum adds up to 19 coefficients. This can be compared to the number arising from the reduced-form process for $\text{vech}(\Sigma_t)$, where the vech operator stacks the lower triangular portion of a matrix into a column vector. For the given example, this vector includes two variances and one covariance. Thus, in a general MGARCH, the equivalent of C has dimension 3×1 and those of G , D and F are 3×3 . Consequently, the number of parameters arrives at a total of $3 + 3(3 \cdot 3) = 30$, which exceeds 19 and hence satisfies the necessary summing-up constraint. In addition, a sufficient condition is given by linear independence of the conditional variances (as in Sentana and Fiorentini 2001), which should normally be met by ARCH-type processes.

The estimation can be done by Maximum Likelihood. For this purpose, the log-likelihood for a sample of T observations (complemented by an adequate number of pre-sample observations) under the assumption of conditional normality is constructed as

$$L(\theta) = -\frac{1}{2} \sum_{t=1}^T (n \log 2\pi + \log |\Sigma_t| + y_t' \Sigma_t^{-1} y_t), \quad (10)$$

where the vector θ stacks all free parameters from C , G , D , F , A , α , β and \bar{Q} . That is, maximisation of (10) yields estimates of both the EGARCH parameters and the structural coefficients governing spillovers and fundamental correlations. As the assumption of conditional normality is often problematic for financial markets data, the estimation relies on Quasi-Maximum-Likelihood (QML, see Bollerslev and Wooldridge 1992). This ensures consistency of the estimation, while standard errors are corrected for possible non-normality. Numerical likelihood optimisation is performed using the BHHH algorithm (Berndt et al. 1974).

3 Once again: Blue Chip vs. High Tech

Weber (2008) applies his SCCC model to the Dow Jones Industrial Average and Nasdaq Composite daily returns for the long sample from 2/5/1971 until 10/31/2007. For ρ , the constant conditional correlation coefficient, he finds an estimate of roughly 20%. Given the considerable proximity of the two stock segments, he argues that one might have expected a much higher coherence of shocks. Indeed, cutting the sample at the end of 1996 raises ρ to 42%. Weber (2008) ascribes this effect to the CCC assumption, which obviously fails to adequately describe the correlation structure through the whole sample including the extremely volatile period around the year 2000.

This discussion suggests that the SDCC model should be able to compensate for the depicted shortcoming: Namely, under an SCCC assumption, the time-varying part of the total correlation logically has to be picked up exclusively by the mutual transmission effects. Logically, above all in times of economic turbulences, the estimation might easily understate the influence of third-party common factors. In view of this problem, the SDCC model is likely to provide a more appropriate impression of the underlying financial processes.

As in Weber (2008), in a first step the returns are regressed on a constant and four day-of-the-week dummies, but no autoregressive lags. Starting values for the optimisation of the likelihood (10) were obtained as follows: The EGARCH parameters were estimated

in univariate models, whereas the off-diagonal elements were set to zero. The variance processes were started at the sample moments. A was initialised as the identity matrix, so that the off-diagonal element from \bar{Q} equalled the unconditional return correlation. However, putting more weight on A and less on \bar{Q} had no relevant impact on the outcome of the QML procedure. Starting values for α and β , which govern the development of the structural conditional correlation, were taken from a conventional reduced-form DCC model. The estimations were carried out in a Gauss programme employing the CML module.

Equations (11), (12) and (13) display the estimation outcome. The variable names denote close-to-close returns at time t , q_t is the off-diagonal element from Q_t , and QML standard errors are in parentheses.

$$\begin{aligned} \text{DJIA}_t &= \underset{(0.036)}{0.326} \text{NQC}_t + \hat{\varepsilon}_{1t} \\ \text{NQC}_t &= \underset{(0.037)}{0.328} \text{DJIA}_t + \hat{\varepsilon}_{2t} \end{aligned} \quad (11)$$

$$\begin{pmatrix} \log h_{1t}^2 \\ \log h_{2t}^2 \end{pmatrix} = \begin{pmatrix} -0.168 \\ (0.028) \\ -0.247 \\ (0.025) \end{pmatrix} + \begin{pmatrix} 0.984 & -0.008 \\ (0.005) & (0.003) \end{pmatrix} \begin{pmatrix} \log h_{1t-1}^2 \\ \log h_{2t-1}^2 \end{pmatrix} + \begin{pmatrix} 0.126 & 0.070 \\ (0.021) & (0.018) \\ 0.114 & 0.176 \\ (0.019) & (0.021) \end{pmatrix} \begin{pmatrix} |\tilde{\varepsilon}_{1t-1}| \\ |\tilde{\varepsilon}_{2t-1}| \end{pmatrix} + \begin{pmatrix} -0.036 & -0.030 \\ (0.010) & (0.007) \\ -0.045 & -0.039 \\ (0.010) & (0.008) \end{pmatrix} \begin{pmatrix} \tilde{\varepsilon}_{1t-1} \\ \tilde{\varepsilon}_{2t-1} \end{pmatrix} \quad (12)$$

$$q_t = \underset{(0.006)}{1} - \underset{(0.008)}{0.021} - \underset{(0.008)}{0.973} \cdot \underset{(0.081)}{0.361} + \underset{(0.006)}{0.021} \tilde{\varepsilon}_{1t-1} \tilde{\varepsilon}_{2t-1} + \underset{(0.008)}{0.973} q_{t-1} \quad (13)$$

The unconditional correlation of the structural innovations rises to more than 1/3, compared to less than 1/5 in the SCCC model. This confirms the presumption that a higher degree of coherence in shocks can be found by allowing for time variation, which is picked up by the SDCC specification. While Weber (2008) obtained a relatively high correlation coefficient only after sample shortening, the present approach achieves such a result over the whole sample including the period of economic and financial disturbances. Logically, the SDCC parameters $\hat{\alpha}$ and $\hat{\beta}$ are clearly significant, taking values that are common in the financial volatility literature. Restricting both of them to zero, that is applying the SCCC assumption, is clearly rejected with a decline in log-likelihood of 225.

Figure 1 shows the structural conditional correlation as the off-diagonal element in R_t as well as its reduced-form counterpart, which is calculated from the covariance-matrix (9). That is, the latter mirrors the correlation effects of causal transmission in addition to the fundamental commonalities in the structural innovations.

The most eye-catching drop in correlations appears in the year 2000, where the extreme spike and fall of the Nasdaq Composite index occurred. Thereafter, correlations jump up again coinciding with 9/11 and the US recession. In the years before, further turbulences

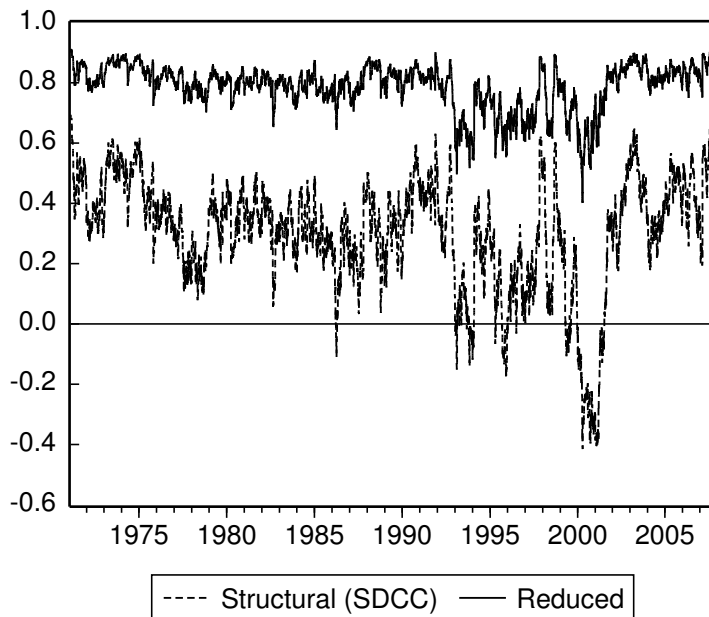


Figure 1: Conditional correlations

took place from the end of 1992 onwards, comprising numerous financial crises like those in Mexico, South-East Asia, Russia and Brazil. A similar pattern has as well been found by Engle (2002) in his reduced-form DCC approach within a shorter sample.

Concerning the direct spillovers, both coefficients in A are highly significant. Since the Dow effect is only marginally higher, the moderate dominance of the Dow found before hardly carries over to the SDCC model. Nevertheless, the causality-in-variance effects from (12) still reveal higher cross-segment influences of the Dow Jones as compared to the Nasdaq Composite. This transmission can be interpreted as a proxy for information flows between markets (Ross 1989). The negative parameters of the signed shocks represent the well-known asymmetric volatility effects. The negative off-diagonal coefficients in the autoregressive matrix indicate a certain dampening influence, which is however economically small. Being smaller than one, both eigenvalues of this matrix meet the stability criterion, even though the usual substantial persistence in variance can be found.

Finally, the model is subjected to several specification tests: As in Weber (2008), the autocorrelations of the squared standardised disturbances $\tilde{\varepsilon}_{jt}^2$ do not exceed the approximate 95% confidence bands, except for the Nasdaq first-order autocorrelation, which does however not reach significance at the 1% level. Furthermore, the autocorrelations of the cross product $\varepsilon_{1t}\varepsilon_{2t}$, standardised by the conditional covariance, are insignificant by the same criterion; again, this shows the benefit of *dynamic* correlations compared to the

SCCC model of Weber (2008), which could not absorb the whole time variation in the structural covariance. Even though the standardised residuals have excess kurtosis (4.1 and 1.6), allowing for heavy tails as in the Student-t-distribution does not relevantly alter the outcome of the maximum likelihood procedure.

4 Concluding Summary

Weber (2008) proposed the structural constant conditional correlation (SCCC) model, which complements non-restricted simultaneous effects between several variables by interaction in their fundamental innovations. That is, an observed correlation can be traced back to the sources direct causality and common shocks. This paper improved on the SCCC model by allowing for *dynamic* conditional correlations in the structural shocks (SDCC).

The methodological enhancement has the effect that even the *time variation* in correlations between financial variables can potentially be explained by unobserved third-party influences in addition to the direct mutual spillovers. In a system of Dow Jones and Nasdaq Composite returns, Weber (2008) found a 20% correlation of the fundamental structural shocks. This relatively low value increased to 36% when the new SDCC model was applied. The reason turned out to be the extremely volatile period in the second half of the 1990s, which could be picked up well by the dynamic specification for the conditional correlations.

In future research, the econometric progress of the SDCC model might be exploited for finding economic interpretations of structural systems, which have hitherto been treated in reduced form. By the same token, existing identification schemes could be checked for their consistence with empirical data.

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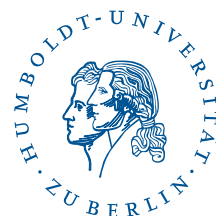
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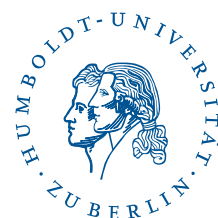
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