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Spectral based methods to identify common trends and common cycles

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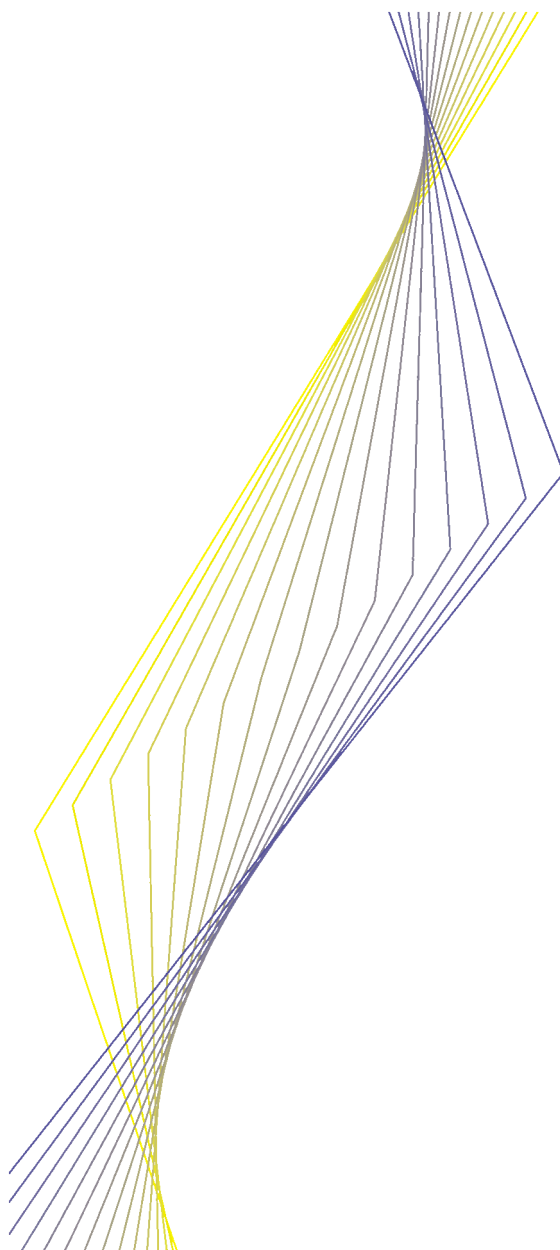
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Spectral Based Methods to Identify Common Trends and Common Cycles

Gonzalo Camba-Mendez*and George Kapetanios†

April 23, 2001

Abstract

The rank of the spectral density matrix conveys relevant information in a variety of modelling scenarios. Phillips (1986) showed that a necessary condition for cointegration is that the spectral density matrix of the innovation sequence at frequency zero is of a reduced rank. In a recent paper Forni and Reichlin (1998) suggested the use of a generalised dynamic factor model to explain the dynamics of a large set of macroeconomic series. Their method relied also on the computation of the rank of the spectral density matrix. This paper provides formal tests to estimate the rank of the spectral density matrix at any given frequency. The tests of rank at frequency zero are tests of the null of ‘cointegration’, complementary to those suggested by Phillips and Ouliaris (1988) which test the null of ‘no cointegration’.

Keywords: Tests of Rank, Spectral Density Matrix, Canonical Correlations.

JEL Classification: C12, C15, C32

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1 Introduction

The equivalence of time-domain and frequency-domain analysis of time series is well documented in the statistical and econometric literature. Nevertheless, the use of spectral densities is by far less widespread than the use of covariances in the econometric analysis of time series. This paper aims to apply both well known and new techniques of rank determination of matrices to the determination of the rank of the spectral density matrix. As the discussion below shows, the need for such techniques is clear from the existing econometric literature.

The idea that the movements in a number of economic variables can be represented by relatively few driving forces was first suggested by Stone (1947). Stone's use of principal components was purely static. Since then there has been a considerable amount of work using alternative data reduction techniques to promote parsimony in dynamic model specification. In a recent paper Forni and Reichlin (1998) suggested the use of a generalised dynamic factor model to describe the dynamics of sectoral industrial output and productivity for the US economy from 1958 to 1986. The number of common shocks driving those series is equal to the rank of their spectral density matrix. The foundations for this result are to be found in the literature on dynamic principal components, see Brillinger (1981). This issue is further explored in Forni, Hallin, Lippi, and Reichlin (1999a) and Forni, Hallin, Lippi, and Reichlin (1999b) where a 'generalised dynamic factor' model, novel to the literature, is proposed. Forni and Reichlin (1998) pointed that no standard test of the rank of the spectral density matrix was available, and consequently this issue was partially sidestepped.

Phillips (1986) showed that a necessary condition for cointegration is that the spectral density matrix of the innovation sequence of an $I(1)$ multivariate process has a reduced rank at frequency zero. Phillips and Ouliaris (1988) suggested two procedures for detecting the presence of cointegration. The drawback of their method was that they were tests of the null of 'no cointegration'. Namely a test of the hypothesis that the r smallest eigenvalues are greater than zero. The tests of rank of the spectral density matrix suggested in this paper may be thought of, under certain conditions, as tests of the null of 'cointegration', i.e. tests of the null that the r smallest eigenvalues are equal to zero.

In order to apply standard tests of rank, it is helpful to think of the estimate of the spectral density matrix as an estimate of a covariance matrix between two newly defined complex variate random processes. Two alternative approaches for determining the rank of the spectral density matrix are presented. On the one hand, the Bartlett

(1947) procedure is applicable in this context because the problem can be recast in terms of canonical correlations. On the other hand, the Cragg and Donald (1996) approach is more general as it only requires that an estimate of that matrix exists having a normal asymptotic distribution with a covariance matrix whose rank is known. An information criterion method suggested by Akaike (1976), which builds upon the Bartlett (1947) procedure is also described.

Section 2 presents the analytical framework. Expressions for two alternative estimates of the spectral density matrix are provided here, together with some background material. The extensions of the Bartlett (1947) test of rank, Akaike (1976) information criterion method and Cragg and Donald (1996) test to the case of the spectral density matrix are described in sections 3 and 4. Section 5 concludes.

2 Background Theory

2.1 The Complex Multivariate Normal Distribution

In this paper references are made to the complex normal distribution. A q -random variable \mathbf{y}_t with complex value components is complex multivariate normally distributed with mean $\boldsymbol{\mu}$ and covariance matrix $\boldsymbol{\Omega}$, and denoted as $N^C(\boldsymbol{\mu}, \boldsymbol{\Omega})$, if the $2q$ -random variable with real components $(\text{Re } \mathbf{y}_t', \text{Im } \mathbf{y}_t')'$ is distributed as

$$N \left(\begin{bmatrix} \text{Re } \boldsymbol{\mu} \\ \text{Im } \boldsymbol{\mu} \end{bmatrix}, \frac{1}{2} \begin{bmatrix} \text{Re } \boldsymbol{\Omega} & -\text{Im } \boldsymbol{\Omega} \\ \text{Im } \boldsymbol{\Omega} & \text{Re } \boldsymbol{\Omega} \end{bmatrix} \right)$$

where Re and Im are operators which extract the real and imaginary part of a complex variate written in its cartesian form, and N denotes the multivariate normal distribution.

If a set of vector random variables, $\mathbf{y}_1, \dots, \mathbf{y}_n$ are i.i.d zero mean complex multivariate normal with covariance $\boldsymbol{\Omega}$, then $\sum_{i=1}^n \mathbf{y}_i \bar{\mathbf{y}}_i'$ is said to have a complex Wishart distribution with n degrees of freedom, and is denoted by $W^C(n, \boldsymbol{\Omega})$.

2.2 Spectral Density Matrix Estimates

Denote a zero mean, wide sense stationary m -vector process by $\{\mathbf{x}_t\}_{t=1}^{\infty}$. The spectral density matrix of \mathbf{x}_t is defined as the infinite Fourier transform of the covariance matrix function,

$$\boldsymbol{\Sigma}(\theta) = (2\pi)^{-1} \sum_{k=-\infty}^{\infty} \boldsymbol{\Gamma}_k e^{-ik\theta}$$

for $\theta \in [-\pi, \pi]$ where $\mathbf{\Gamma}_k = E\{\mathbf{x}_t \mathbf{x}'_{t-k}\}$. Given a sample of T observations an obvious estimate of the spectral density matrix is given by:

$$\bar{\mathbf{\Sigma}}(\theta) = (2\pi)^{-1} \sum_{k=-(T-1)}^{T-1} \hat{\mathbf{\Gamma}}_k e^{-ik\theta}$$

where $\hat{\mathbf{\Gamma}}_k = \frac{1}{T} \sum_{t=1}^{T-|k|} \mathbf{x}_t \mathbf{x}'_{t-k}$. $2\pi \bar{\mathbf{\Sigma}}(\theta)$ is known as the periodogram. It is well known that the periodogram is an inconsistent but asymptotically unbiased estimate of the spectral density matrix, multiplied by 2π and is asymptotically distributed as a complex Wishart variable with 1 degree of freedom. This suggests some form of smoothing can provide a consistent estimate of the spectral density matrix¹. Two approaches are usually adopted. Firstly, methods that rely on ‘smoothing’ the autocovariance matrix function which take the form:

$$\hat{\mathbf{\Sigma}}(\theta) = \sum_{k=-M}^M \hat{\mathbf{\Gamma}}_k \omega(k, M) e^{-ik\theta} \quad (1)$$

where M is a bandwidth parameter and $\omega(k, M)$ is a kernel or spectral window satisfying certain regularity conditions, (see, e.g. Fuller (1996)). As long as $M \rightarrow \infty$ and $T^{-1}M \rightarrow 0$ as $T \rightarrow \infty$, $\hat{\mathbf{\Sigma}}(\theta)$ provides a consistent estimate of $\mathbf{\Sigma}(\theta)$, see, e.g., Fuller (1996, pp. 382-383).

Secondly methods that rely on ‘smoothing’ the periodogram itself over the frequencies, i.e. averaging adjacent frequency ordinates. These estimates take the form,

$$\tilde{\mathbf{\Sigma}}(\theta) = \frac{1}{2M+1} \sum_{k=-M}^M \bar{\mathbf{\Sigma}}(\theta + k/T) \quad (2)$$

For finite M this estimate is still inconsistent, asymptotically unbiased for the spectral density matrix and asymptotically distributed as $(2M+1)^{-1} W^C(2M+1, \mathbf{\Sigma}(\theta_j))$, (see Brillinger (1981, pp. 245)). This is the simplest form of a smoothed periodogram estimate for the spectral density matrix. Different weights can be assigned to the periodogram coordinates $\bar{\mathbf{\Sigma}}(\theta + k/T)$, see Brillinger (1981, Chapter 7). If we allow $M \rightarrow \infty$ as $T \rightarrow \infty$ but impose $M/T \rightarrow 0$ we get a consistent estimate. In particular we obtain the result that $\sqrt{M}(\text{vec}(\tilde{\mathbf{\Sigma}}(\theta)) - \text{vec}(\mathbf{\Sigma}(\theta)))$ is asymptotically complex normal² with a covariance matrix whose element giving the covariance between $\Sigma_{i,j}(\theta)$ and $\Sigma_{u,v}(\theta)$ is given by $\Sigma_{i,j}(\theta)\Sigma_{u,v}(\theta)$, where $\Sigma_{i,j}(\theta)$ is the (i, j) -th element of $\mathbf{\Sigma}(\theta)$. We will denote this covariance matrix by \mathbf{V} and its estimate, obtained by using the estimated spectral density matrix, by $\tilde{\mathbf{V}}$. More details may be found in e.g. Brillinger (1981, pp. 262) or Brockwell

¹As we are mainly interested with the rank of the spectral density matrix, in the rest of the discussion we drop the normalising constant 2π .

²Note that this result crucially depends on regulating the dependence properties of the series. For more details see Brillinger (1981, Chapter 2).

and Davis (1991, pp. 447). Alternative strategies to select the weighting parameters to compute $\hat{\Sigma}(\theta)$ and $\hat{\Sigma}(\theta_j)$ are well documented in the literature, see Priestley (1981) or Brockwell and Davis (1991) for further details.

3 Bartlett (1947) test of rank: Applicable to $\hat{\Sigma}(\theta)$

A well known result in canonical correlation analysis is that given two random vector series \mathbf{x}_{1t} and \mathbf{x}_{2t} , the rank of their covariance matrix is equal to the number of nonzero canonical correlations. Further details are in Anderson (1984). Note that for a given bandwidth M and a chosen kernel $\omega(k, M)$, the spectral density matrix estimate $\hat{\Sigma}(\theta)$ in (2) can be written as $\hat{\Sigma}(\theta) = \hat{\Pi}\Psi(\theta)$ where

$$\hat{\Pi} = \begin{bmatrix} \hat{\Gamma}_{-M} & \dots & \hat{\Gamma}_0 & \dots & \hat{\Gamma}_M \end{bmatrix} \quad \Psi(\theta) = \begin{bmatrix} \mathbf{I}_m \times \phi_{-M}^\theta \\ \dots \\ \mathbf{I}_m \times \phi_0^\theta \\ \dots \\ \mathbf{I}_m \times \phi_M^\theta \end{bmatrix}$$

\mathbf{I}_m is an identity matrix of order m , and $\phi_s^\theta = \omega(s, M)e^{is\theta}$.

The interesting feature is that the spectral density matrix estimate is equivalent to the estimate of the covariance matrix between two complex variates, and defined as $\Sigma(\theta) = E\{\mathbf{x}_{1t}\mathbf{x}_{2t}'\}$, where $\mathbf{x}_{1t} = \mathbf{x}_t$ and $\mathbf{x}_{2t} = \Psi(\theta)'(\mathbf{x}_{t+M}', \dots, \mathbf{x}_t', \dots, \mathbf{x}_{t-M}')'$. This implies that testing for the rank of $\hat{\Sigma}(\theta)$ above is equivalent to testing for the number of positive canonical correlations between \mathbf{x}_{1t} and \mathbf{x}_{2t} . To compute the canonical correlations between \mathbf{x}_{1t} and \mathbf{x}_{2t} we proceed as follows. Define the matrices \mathbf{X}_1 , \mathbf{X}_2 and $\Psi(\theta)$ as:

$$\mathbf{X}_1 = \begin{bmatrix} \mathbf{x}'_1 \\ \dots \\ \mathbf{x}'_M \\ \mathbf{x}'_{M+1} \\ \dots \\ \mathbf{x}'_{T-M} \\ \mathbf{x}'_{T-M+1} \\ \dots \\ \mathbf{x}'_T \end{bmatrix} \quad \mathbf{X}_2 = \begin{bmatrix} \mathbf{x}'_{M+1} & \dots & \mathbf{x}'_1 & \dots & 0 \\ \dots & \dots & \dots & \dots & \dots \\ \mathbf{x}'_{(2 \times M)} & \dots & \mathbf{x}'_M & \dots & 0 \\ \mathbf{x}'_{(2 \times M)+1} & \dots & \mathbf{x}'_{M+1} & \dots & \mathbf{x}'_1 \\ \dots & \dots & \dots & \dots & \dots \\ \mathbf{x}'_T & \dots & \mathbf{x}'_{T-M} & \dots & \mathbf{x}'_{T-(2 \times M)} \\ 0 & \dots & \mathbf{x}'_{T-M+1} & \dots & \mathbf{x}'_{T-(2 \times M)+1} \\ \dots & \dots & \dots & \dots & \dots \\ 0 & \dots & \mathbf{x}'_T & \dots & \mathbf{x}'_{T-M} \end{bmatrix}$$

Compute the QR decomposition of the matrices \mathbf{X}_1 and $\mathbf{X}_2\Psi(\theta)$ given above, i.e. $\mathbf{X}_1 = \mathbf{Q}_1\mathbf{R}_1$ and $\mathbf{X}_2\Psi(\theta) = \mathbf{Q}_2\mathbf{R}_2$. The canonical correlations between the vectors \mathbf{x}_{1t} and \mathbf{x}_{2t} , are the singular values of $\mathbf{Q}_1'\mathbf{Q}_2$. We denote the canonical correlations as ρ_i , $i = 1, \dots, m$. Bartlett (1947) provided a criterion for testing the null hypothesis that the last $m - r^*$

canonical correlations are zero, i.e., $H_{r^*} : \rho_{r^*+1} = \dots = \rho_m = 0$. Under the null hypothesis and assuming stationarity of the input-output multivariate system

$$BA = \left[\frac{2m+1}{2} - T \right] \ln \prod_{i=r^*+1}^m (1 - \hat{\rho}_i^2) \xrightarrow{d} \chi_{(m-r^*)^2}^2$$

Fujikoshi (1974) proved that this test procedure is based on the likelihood ratio method. His results can easily be extended to complex random variables providing justification for the application of the method in the current context. Bartlett's test was developed under independence and normality assumptions, but his result remains valid asymptotically following arguments by Kohn (1979) on the likelihood ratio tests for dependent observations.

3.1 Akaike (1976) Information Criterion

Akaike (1974) and Akaike (1976) showed that the number of linearly independent components of the projections of the previously defined \mathbf{x}_{1t} onto the linear space spanned by the components of \mathbf{x}_{2t} is identical to the number of nonzero canonical correlations between \mathbf{x}_{1t} and \mathbf{x}_{2t} . When \mathbf{x}_t is Gaussian, canonical correlation analysis between \mathbf{x}_{1t} and \mathbf{x}_{2t} is equivalent to maximum likelihood estimation of the linear model: $\mathbf{x}_{1t} = \Psi \mathbf{x}_{2t} + \varepsilon_t$, see Anderson (1984). The number of free parameters for this model is: $F(r^*) = \{[s_1(s_1+1)]/2\} + \{[s_2(s_2+1)]/2\} + r^*(s_1 + s_2 - r^*)$ where s_1 denotes the dimension of the vector \mathbf{x}_{1t} and s_2 denotes the dimension of \mathbf{x}_{2t} . The first two terms are the number of free parameters of the covariance matrices of \mathbf{x}_{1t} and \mathbf{x}_{2t} respectively, and the last term gives the number of free parameters in matrix Ψ . Akaike (1976) defined an information criterion for model fitting, and by extension rank determination, as:

$$AIC(r^*) = T \ln \prod_{i=1}^{r^*} (1 - \hat{\rho}_i^2) + 2F(r^*)$$

where $\hat{\rho}_i$ are the estimated canonical correlation coefficients previously defined. This criterion penalises models with a large number of parameters, and by extension large rank, and favours parsimonious representations. Schwarz (1978) suggested an alternative penalty on increasing the number of parameters. The information criterion suggested by Schwarz (1978) amounts to replace $2F(r^*)$ above by $\ln(T)F(r^*)$. The penalty used by this method is much more severe than that used by AIC.

4 Cragg and Donald test (1996)

We give an account of the general test of rank proposed by Cragg and Donald (1996) before discussing modifications for the problem at hand. For a general matrix Δ , the

procedure proposed by Cragg and Donald (1996) is based on the transformation of the matrix Δ using Gaussian elimination with complete pivoting³. r^* steps of Gaussian elimination with complete pivoting on matrix Δ amounts to the following operations:

$$\mathbf{Q}_{r^*} \mathbf{R}_{r^*} \mathbf{Q}_{r^*-1} \mathbf{R}_{r^*-1} \dots \mathbf{Q}_1 \mathbf{R}_1 \Delta \mathbf{C}_1 \dots \mathbf{C}_{r^*-1} \mathbf{C}_{r^*} = \begin{bmatrix} \Delta_{11}(r^*) & \Delta_{12}(r^*) \\ \mathbf{0} & \Delta_{22}(r^*) \end{bmatrix}$$

where \mathbf{R}_i and \mathbf{C}_i are pivoting matrices for step i and \mathbf{Q}_i are Gauss transformation matrices. The pivoting matrices used to perform the first r^* steps of Gaussian elimination are applied to Δ to obtain the following relation

$$\mathbf{R}_{r^*} \mathbf{R}_{r^*-1} \dots \mathbf{R}_1 \Delta \mathbf{C}_1 \dots \mathbf{C}_{r^*-1} \mathbf{C}_{r^*} = \mathbf{R} \Delta \mathbf{C} = \mathbf{F} = \begin{bmatrix} \mathbf{F}_{11}(r^*) & \mathbf{F}_{12}(r^*) \\ \mathbf{F}_{21}(r^*) & \mathbf{F}_{22}(r^*) \end{bmatrix}$$

where \mathbf{F} is partitioned accordingly, i.e. $\mathbf{F}_{11}(r^*)$ is of dimension $r^* \times r^*$. Note that in this case $\mathbf{F}_{11}(r^*)$ has full rank, under the null hypothesis that r^* is equal to the true rank. It then follows, see Cragg and Donald (1996), that $\mathbf{F}_{22}(r^*) - \mathbf{F}_{21}(r^*) \mathbf{F}_{11}^{-1}(r^*) \mathbf{F}_{12}(r^*) = \mathbf{0}$. The estimated counterpart of the above relation, i.e. $\hat{\mathbf{F}}_{22} - \hat{\mathbf{F}}_{21} \hat{\mathbf{F}}_{11}^{-1} \hat{\mathbf{F}}_{12} = \hat{\Lambda}_{22}(r^*)$, may be used as a test statistic of the hypothesis that the rank of Δ is r^* . Under regularity conditions, including the requirement that the covariance matrix of the asymptotically normally distributed matrix $\sqrt{T} \text{vec}(\hat{\Delta} - \Delta)$, denoted by \mathbf{V} , has full rank, the following result can be shown, under H_0 .

$$\sqrt{T} \text{vec}(\hat{\Lambda}_{22}(r^*)) \xrightarrow{d} N(\mathbf{0}, \Upsilon \mathbf{V} \Upsilon')$$

where $\Upsilon = \Phi_2 \otimes \Phi_1$ and $\Phi_1 = \begin{bmatrix} -\hat{\mathbf{F}}_{21} \hat{\mathbf{F}}_{11}^{-1} & \mathbf{I}_{m-r^*} \end{bmatrix} \mathbf{R}$, $\Phi_2 = \begin{bmatrix} -\hat{\mathbf{F}}'_{12} \hat{\mathbf{F}}_{11}^{-1'} & \mathbf{I}_{m-r^*} \end{bmatrix} \mathbf{C}'$ and \xrightarrow{d} denotes convergence in distribution. Then,

$$\hat{\xi} = T \text{vec} \hat{\Lambda}_{22}(r^*)' (\hat{\Upsilon} \hat{\mathbf{V}} \hat{\Upsilon}')^{-1} \text{vec} \hat{\Lambda}_{22}(r^*) \xrightarrow{d} \chi_{(m-r^*)^2}^2$$

where $\hat{\Upsilon}$ and $\hat{\mathbf{V}}$ are the sample estimates of Υ and \mathbf{V} and χ_l^2 denotes the χ^2 distribution with l degrees of freedom. The procedure uses the inverse of the estimated asymptotic covariance matrix of Δ , but this may not be available if the covariance matrix is of reduced rank. However, Camba-Mendez and Kapetanios (2000) proved that the use of the Moore-Penrose inverse, denoted as $(\hat{\Upsilon} \hat{\mathbf{V}} \hat{\Upsilon}')^+$, is valid if the rank of \mathbf{V} is known and $rk[\hat{\mathbf{V}}] = rk[\mathbf{V}]$, $\forall T > T_0$, for some finite T_0 . Note that by construction the estimate of the asymptotic covariance matrix estimator of $\tilde{\Sigma}(\theta)$, $\tilde{\mathbf{V}}$, has this property. A sequential application of the Cragg and Donald test of rank can provide a consistent estimate of

³For details on Gaussian elimination with complete pivoting see Cragg and Donald (1996) or Golub and Loan (1983).

the rank of Δ if the significance level used in the test converges to zero as the number of observations tends to infinity (See Hosoya (1989)).

We note that the Gaussian elimination results underlying the Cragg and Donald test remain valid for complex matrices such as $\Sigma(\theta)$ and its estimator $\tilde{\Sigma}(\theta)$. To see that we simply note that the r -th step of Gaussian elimination simply involves complete pivoting so as to bring the largest elements of the matrix in absolute value to the r -th place in the diagonal and subsequent zeroing out of the elements below that diagonal elements. Therefore, an $m \times m$ complex matrix with rank r subject to r steps of Gaussian elimination will be transformed into a matrix whose lower RHS $m - r \times m - r$ submatrix will be made up of zeros. The estimated counterpart of such a submatrix will have a multivariate complex normal with zero mean as long as the matrix is multivariate complex normal before the application of Gaussian elimination. This implies that the real and imaginary parts of the, typically complex in the case of spectral density estimates, elements of this submatrix, will be normally distributed with known covariance matrix, leading to a standard χ^2 test just like in the case of the Cragg and Donald test with real matrices.

Similar results may hold for the estimator $\hat{\Sigma}(\theta)$ conditional on an asymptotic normality result for it similar to that provided by Brillinger (1981, pp. 262) for $\tilde{\Sigma}(\theta)$. We do not know of a statement of such a result in the literature, although it seems safe to conjecture that such a result holds.

5 Conclusion

This paper has formulated a variety of rank determination procedures for the rank of the spectral density matrix at any frequency. Different estimators of the spectral density matrix have been considered. Both testing procedures and information criteria have been suggested and justified as valid rank determination methods. The need for such techniques becomes apparent in the econometric literature in areas such as multivariate factor models and cointegration. Phillips and Ouliaris (1988) suggested tests of the null of ‘no cointegration’ which amounted to a test of the hypothesis that the r smallest eigenvalues of the spectral density matrix of the innovation sequence at frequency zero are greater than zero. Phillips and Ouliaris (1990) expanded on the issue of choice of the null hypothesis in cointegration testing by pointing out that adopting the null hypothesis of cointegration may be more sensible from a methodological point of view given that cointegration is the focus of interest. However, it was also pointed out that

standard test statistics based on the spectral density matrix provided inconsistent tests under the null hypothesis of no cointegration. This paper has described tests of the rank of the spectral density matrix which may serve, at frequency zero, as tests of the null of ‘cointegration’. It is clear that, as long as a consistent estimate of the spectral density matrix of the innovation process exists and has an asymptotic complex normal distribution, the application of the Cragg and Donald test will provide a consistent testing procedure for cointegration. The tests of the rank of the spectral density matrix described in this paper are also relevant to identify the structure of a vector times series under the approach of dynamic principal components. Further use of the techniques may be envisaged in terms of restricting the dimensionality of cyclical components at individual frequencies, possibly motivated from economic theory, thereby extending the common cycle analysis discussed in, e.g., Vahid and Engle (1993) and Engle and Issler (1995).

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