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**DETECTING CHANGES  
IN THE A. R. PARAMETERS  
OF A NONSTATIONARY  
A. R. M. A. PROCESS**

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## DETECTING CHANGES IN THE A.R. PARAMETERS OF A NONSTATIONARY A.R.M.A. PROCESS.

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### RESUME

Nous présentons une méthode pour détecter des changements dans les paramètres AR (supposés constants avant le changement) d'un processus ARMA non stationnaire. La méthode de détection est fondée sur une approche asymptotique locale, et trouve son origine dans une procédure d'identification dont on a montré qu'elle était consistante en présence de non stationnarités. Les démonstrations sont détaillées.

### ABSTRACT

We present a method for detecting changes in the AR parameters (assumed to be time invariant) of a nonstationary ARMA process. The detection method is derived by using a local asymptotic approach, and is based upon an estimation procedure which was shown to be consistent under nonstationarities.

Ce travail se situe dans le cadre du GRECO-CNRS "systèmes adaptatifs".

## I. INTRODUCTION.

The problem of detecting changes in the spectral parameters of processes is frequently encountered in practice. Most methods (likelihood or innovation based techniques) solve this problem by using a complete description of the spectral parameters, (see for example [2,8,11]). Sometimes this description is not possible for all parameters, this is for instance the case when nonstationarities are present. In such situations one is mainly interested in those parameters that can be described, while regarding the others as nuisance. A problem of this type is the problem of vibration monitoring. Here one is interested in detecting changes in the vibrating modes of a system, without taking into account the cause that generates these vibrations. If we use an ARMA model to describe the system then the problem translates into detecting changes in the AR parameters (assumed time invariant) while the MA part is assumed nonstationary and unknown. Clearly the classical techniques do not apply directly here because of the high variability of the MA part. In this paper we give a solution to this problem. The method that we present is based on an estimation method for the AR parameters that is known to be consistent even under nonstationarity. To derive the detection scheme we use asymptotic techniques assuming that the number of observations goes to infinity and the magnitude of the change goes to zero. Similar asymptotic techniques are used in [5,9], but they are mainly applied to the likelihood function, something which is not possible here.

## II. PROBLEM STATEMENT.

Let us consider the following system of difference equations

$$\begin{aligned} X_{k+1} &= FX_k + W_{k+1} \quad X_0 = 0 \\ y_k &= c^T X_k \end{aligned} \tag{1}$$

where  $F$  is a real square matrix of dimension  $m$ ,  $c$  is a real vector of dimension  $m$  and  $\{W_k\}$  is a sequence of zero mean independent nonstationary vectors. With the superscript "  $r$  " we denote the transpose. The model in (1) is often used to model real systems. The vector  $X_k$  is called the state of the system,  $W_k$  the input and the scalar  $y_k$  the output (observations). The process  $y_k$  is an ARMA process and if we write it in this form then we have

$$y_k - \alpha_1 y_{k-1} - \dots - \alpha_m y_{k-m} = \beta_k^0 e_k + \dots + \beta_k^{m-1} e_{k-m+1} \quad (2)$$

where  $\{e_k\}$  is a standard i.i.d. sequence. The vector  $\alpha^T = [\alpha_1 \dots \alpha_m]$  is the vector of the AR parameters and the vector  $\beta_k^T = [\beta_k^0 \dots \beta_k^{m-1}]$  is the vector of the MA parameters. The vector  $\alpha$  is related only to the matrix  $F$ , its components are the coefficients of the characteristic polynomial of this matrix. Since the roots of this polynomial are the spectral modes of the system, all the spectral information is contained in the vector  $\alpha$ . The vector  $\beta_k$  is related to the input  $W_k$  thus it is time varying. Notice that with the model in (1) we can have a MA part in (2) of order at most  $m$ .

Since any change in the spectral modes reflects into a change in the vector  $\alpha$  we concentrate on this vector. Thus we are interested in detecting changes in the vector  $\alpha$  in a *nonsequential way*, assuming that we have available the observation sequence  $\{y_k\}$ . We will assume that a nominal vector  $\alpha = \alpha_0$  is known and that we do not know the vector  $\alpha$  after the change. More specifically we will assume that a collection  $\{y_1, \dots, y_n\}$  is given and we would like to decide between the two hypotheses,  $H_0 : \alpha = \alpha_0$  and  $H_1 : \alpha \neq \alpha_0$ . As we said in the introduction we will follow an asymptotic approach, thus we suppose that under change,  $\alpha$  is of the form  $\alpha = \alpha_0 + \vartheta/\sqrt{n}$ , where  $\vartheta$  is an unknown direction of change. Although for our test we will never use the matrix  $F$  we will assume that  $\alpha_0$  corresponds to some nominal  $F_0$  and  $\alpha_0 + \vartheta/\sqrt{n}$  to  $F_0 + \Theta/\sqrt{n}$ . Let us now consider the following vector:

$$U_n = \frac{1}{\sqrt{n}} \sum_{k=1}^n (y_{k+1} - Y_k^T \alpha_0) Z_k \quad (3)$$

where  $Y_k^T = [y_k \cdots y_{k-m+1}]$  and  $Z_k^T = [y_{k-m+1} \cdots y_{k-m-M+2}]$ . Equation (3) can be used to estimate the AR parameters of the observation process. Since  $U_n$  (disregarding  $1/\sqrt{n}$ ) is the sample correlation of the MA part with delayed enough observation and since this correlation is zero, we can estimate  $\alpha$  by solving the system  $U_n = 0$  (in the mean square sense if  $M > m$ ). This estimate is known to be consistent even under nonstationarities [4]. We use a similar idea to derive a test. We expect that under no change the vector  $U_n$  will have a zero mean and under change a nonzero mean. Indeed we have

$$\begin{aligned} \Sigma_n^{-1/2} U_n &\xrightarrow{d} N(0, I) && \text{for } H_0 \\ \Sigma_n^{-1/2} (U_n - H_n \vartheta) &\xrightarrow{d} N(0, I) && \text{for } H_1 \end{aligned} \quad (4)$$

where  $\xrightarrow{d}$  means convergence in distribution,  $N(0, I)$  is the standard vector Gaussian density and  $\Sigma_n$  is the covariance matrix of  $U_n$ . The matrix  $H_n$  is defined as

$$H_n = \frac{1}{n} \sum_{k=1}^n Z_k Y_k^T \quad (5)$$

We can see that under both hypotheses  $U_n$  has asymptotically the same variance  $\Sigma_n$  but not the same mean. Next step is to use  $U_n$  as if we had equality in (4) and define a log-likelihood

$$T_n' = U_n^T \Sigma_n^{-1} H_n \vartheta - \frac{1}{2} \vartheta^T H_n^T \Sigma_n^{-1} H_n \vartheta \quad (6)$$

Notice that  $Z_k$  is of length  $M$  thus  $H_n$  is of dimension  $M \times m$ , assuming  $M \geq m$  let us substitute  $\vartheta$  by its maximum likelihood estimate, this yields

$$T_n = U_n^T \Sigma_n^{-1} H_n (H_n^T \Sigma_n^{-1} H_n)^{-1} H_n^T \Sigma_n^{-1} U_n \quad (7)$$

The quantity  $T_n$  will be our test statistic. To define the threshold we have that under no change  $T_n$  is asymptotically  $\chi^2$  with  $m$  degrees of freedom. Under  $H_1$ ,  $T_n$  is a noncentered  $\chi^2$  with noncentricity equal to  $\vartheta H_n^T \Sigma_n^{-1} H_n \vartheta$ . Our test will be able to detect the change if the noncentricity is nonzero. As we will show in Section III, with our assumptions this is always the case. The covariance matrix  $\Sigma_n$  is not known and thus we must estimate it. We can for example use the sample covariance matrix defined as

$$\hat{\Sigma}_n = \frac{1}{n} \sum_{k=1}^n \nu_k^2 Z_k Z_k^T + \frac{1}{n} \sum_{j=1}^{m-1} \sum_{k=1}^n \nu_k \nu_{k-j} (Z_k Z_{k-j}^T + Z_{k-j} Z_k^T) \quad (8)$$

where  $\nu_k = y_k - Y_{k-1}^T \alpha_0$ . Thus everything is defined in terms of known things.

### III. ASSUMPTIONS AND PROOFS.

In this section we will prove the validity of (4), i.e. that the Central Limit Theorem (CLT) holds for  $U_n$  under both hypotheses. Notice also that in (7) we take the inverse of the matrices  $\Sigma_n$  and  $H_n^T \Sigma_n^{-1} H_n$ . We will show that these two matrices are full rank. For  $\Sigma_n$  this will be necessary for the proof of the CLT, for the other matrix it is important in showing that the noncentricity factor is nonzero. Notice that the second matrix will be full rank if  $H_n$  is full rank and  $\Sigma_n$  is nonsingular and bounded from above. The proofs will be presented in several steps. Let us first state our assumptions. Let  $\{y_k\}$  be an ARMA process generated by a system of the form of (1) or (2). Let  $F_0, \alpha_0$  be the nominal values of  $F$  and  $\alpha$ . We assume:

A1. The matrix  $F_0$  is full rank and has all its eigenvalues distinct and strictly inside the unit circle.

A2. There exists a real  $p > 0$  such that for every vector  $\lambda$  and every integer

$k$  we have  $E\{[\lambda^T W_k]^4\} \leq p(\lambda^T \lambda)^2$

A3. The matrix  $O = [c \ F_0^T c \dots \ (F_0^T)^{m-1} c]$  is of full rank.

A4. If  $Q_k = E\{W_k W_k^T\}$  is the input covariance matrix at time  $k$ , then we assume that there exists a nonzero vector  $s$  such that for every  $k$  we have  $Q_k \geq ss^T$ .

A5. If  $\bar{Q}_k = \frac{1}{n} \sum_{k=1}^n Q_k$  is the mean input covariance, we assume that there exists a real  $\delta > 0$  such that for every eigenvalue  $\tau_i, i = 1, \dots, m$  of  $F_0$  we have

$$| \lim_{z \rightarrow \tau_i} \{ (z - \tau_i) c^T (zI - F_0)^{-1} \bar{Q}_n (z^{-1} - F_0)^{-1} c \} | \geq \delta$$

By  $E\{ \}$  we denote expectation. Assumption A1 is to ensure the stability of our system, i.e. bounded inputs will produce bounded outputs. Also the assumption that the eigenvalues are distinct is made only to facilitate certain proofs. Assumption A2 is rather technical it requires that the input has uniformly bounded fourth order moments. A3 is an observability condition, it ensures that any mode excited by the input will be observed in the output. Assumption A4 is to ensure that the input excites constantly at least at the direction of the vector  $s$ . Finally A5 means that in the mean the input excites all the modes of the system and that these modes are present in the second order statistics. This assumption is very important since our test is based on second order statistics. If for example A5 was not true for some mode, in other words, if a mode even though excited by the input was not present in the second order statistics (this is the case when the MA part of a stationary ARMA process has zeros that are mirror images, with respect to the unit circle, of poles of the AR part), then any change on this mode cannot be detected by simply using second order statistics. Notice also that A5 is referred only to the mean input, that is, instantaneously we can have cancellation of modes. As one can see this assumption will be



the base for proving that  $H_n$  is of full rank.

Before going to the proof of the CLT we will first present some lemmas that will be useful for this proof. Let us denote by  $A$  and  $B$  two real square matrices of dimension  $m$ , with distinct eigenvalues. Let  $\xi_X$  denote the eigenvalue of a square matrix  $X$  with the maximum magnitude and  $D_X$  a constant that depends only on  $X$ .

**Lemma 1.** There exist constants  $D_A, D_{A,B}$  such that for any two vectors  $a, b$  and any integer  $k \geq 0$  we have

$$i. \quad |a^T A^k b| \leq D_A |\xi_A|^k (a^T a)^{\frac{1}{2}} (b^T b)^{\frac{1}{2}}$$

$$ii. \quad |a^T (A^k - B^k) (A^k - B^k)^T a|^{\frac{1}{2}} \leq D_{A,B} |\xi_{A,B}|^k \|A - B\| (a^T a)^{\frac{1}{2}}$$

where by  $\|X\|$  we denote the maximum singular value of the matrix  $X$ .

**Proof.** The proof of i. is easy. Notice that it is obvious when  $A$  is diagonal. When it is not diagonal we make a diagonalization and proceed in a similar way. Inequality ii. can be proved by using i. and that  $A^k - B^k = A^{k-1}(A - B) + A^{k-2}(A - B)B + \dots + (A - B)B^{k-1}$ , we omit the details.

**Lemma 2.** Let  $F, W_k$  satisfy A1 and A2, then we also have that  $X_k$  has uniformly bounded forth order moments.

**Proof.** From (1) we have that

$$X_k = \sum_{i=0}^{k-1} F^i W_{k-i}$$

thus using the fact that  $\{W_k\}$  is a sequence of independent vectors, we have for every vector  $\lambda$  that

$$E\{[\lambda^T X_k]^4\} = E\left\{\sum_{i=0}^{k-1} \sum_{j=0}^{k-1} \sum_{l=0}^{k-1} \sum_{s=0}^{k-1} [\lambda^T F^i W_{k-i}][\lambda^T F^j W_{k-j}][\lambda^T F^l W_{k-l}][\lambda^T F^s W_{k-s}]\right\} =$$

$$\sum_{i=0}^{k-1} E\{[\lambda^T F^i W_{k-i}]^4\} + 6 \sum_{j=0}^{k-1} \sum_{\substack{l=0 \\ j \neq l}}^{k-1} E\{[\lambda^T F^j W_{k-j}]^2\} E\{[\lambda^T F^l W_{k-l}]^2\} \leq$$

$$\sum_{i=0}^{k-1} E\{[\lambda^T F^i W_{k-i}]^4\} + 6 \left[ \sum_{i=0}^{k-1} E\{[\lambda^T F^i W_{k-i}]^2\} \right]^2$$

Using Lemma 1(i) and A2 and the fact that  $|\xi_F| < 1$  (from A1), we have that there exists constant  $D$  such that for every  $k$  we have

$$E\{[\lambda^T X_k]^4\} \leq D(\lambda^T \lambda)^2$$

**Lemma 3.** Let  $F, W_k$  satisfy A1 and A2. Consider the following sequence of covariance matrices  $P_k$

$$P_{k+1} = FP_k F^T + Q_{k+1}, \quad P_0 = 0 \quad (9)$$

The matrix  $P_k$  is the covariance matrix of the state vector  $X_k$ . Let now  $\bar{Q}_n = \frac{1}{n} \sum_{k=1}^n Q_k$  and  $\bar{P}_n = \frac{1}{n} \sum_{k=1}^n P_k$ . Let  $P$  be the solution of the equation  $P = FPF^T + \bar{Q}_n$ , then  $\|\bar{P}_n - P\| \rightarrow 0$ . The matrix  $P$  is nothing but the state covariance matrix of a stationary system with covariance input  $\bar{Q}_n$  (assumed constant).

**Proof.** Since  $W_k$  has uniformly bounded fourth-order moments, it will also have uniformly bounded second order moments, thus from (9), assuming  $Q_0 = 0$  we have

$$\bar{P}_n = \frac{1}{n} \sum_{k=1}^n P_k = \frac{1}{n} \sum_{k=1}^n \sum_{j=0}^{k-1} F^j Q_{k-j} (F^T)^j = \sum_{j=0}^{n-1} F^j \left( \frac{1}{n} \sum_{k=j+1}^n Q_{k-j} \right) (F^T)^j =$$

$$\sum_{j=0}^{\infty} F^j \bar{Q}_n (F^T)^j - \sum_{j=n}^{\infty} F^j \bar{Q}_n (F^T)^j - \frac{1}{n} \sum_{j=0}^{n-1} F^j \left( \sum_{k=n-j+1}^n Q_k \right) (F^T)^j \quad (10)$$

The first term in (10) is equal to  $P$ , the second is uniformly of order  $|\xi_F|^{2n}$  and the third term of order  $\frac{1}{n}$ . Thus both last terms tend to zero. And this

concludes the proof. Using this lemma we can show that we can approximate (order  $\frac{1}{n}$ ) any average of matrices of the form  $\bar{V}_n = \frac{1}{n} \sum_{k=1}^n E\{[y_k \cdots y_{k-j}]^T [y_{k-j} \cdots y_{k-j-i}]\}$ , by the corresponding matrix  $V$  of a stationary system with input covariance  $\bar{Q}_n$  (assumed constant). This is true because  $E\{y_k y_{k+j}\} = c^T F^j P_k c$  and thus we can approximate every term in  $\bar{V}_n$  by the corresponding stationary term.

**Lemma 4.** Let  $Q, Q'$  be two nonnegative definite matrices with  $Q \geq Q'$ . Let  $F$  satisfy A1. Consider the following two systems in stationary situation.

$$\begin{aligned} X_{k+1} &= FX_k + W_{k+1} & y_k &= c^T X_k \\ X'_{k+1} &= FX'_k + W'_{k+1} & y'_k &= c^T X'_k \end{aligned} \quad (11)$$

with  $W_k$  having covariance  $Q$  and  $W'_k$  covariance  $Q'$ . For any integer  $N$  we then have

$$E\{[y_k \cdots y_{k-N}]^T [y_k \cdots y_{k-N}]\} \geq E\{[y'_k \cdots y'_{k-N}]^T [y'_k \cdots y'_{k-N}]\} \quad (12)$$

**Proof.** The proof is easy. Since we consider only second order statistics we can decompose  $W_k$  into two independent processes  $V_k, R_k$  such that  $E\{V_k V_k^T\} = Q'$ ,  $E\{R_k R_k^T\} = Q - Q'$  and  $E\{V_k R_k^T\} = 0$ . Because of linearity, the process  $y_k$  can be also decomposed into two independent processes one due to  $V_k$  and the other to  $R_k$ . The process due to  $V_k$  will have exactly the same second order properties with  $y'_k$ . Thus we conclude that the left hand side of (12) is equal to the right hand side plus another nonnegative term due to  $R_k$ .

**Lemma 5.** Let  $F, F'$  be two matrices satisfying A1 and  $W_k$  a process satisfying A2. Let  $X_k$  be the state vector of a system of the form of (1) and  $X'_k$

another state vector of a similar system but  $F$  replaced by  $F'$ . For any vector  $\lambda$  with  $\lambda^T \lambda = 1$  we have

$$E\{[\lambda^T(X_k - X'_k)]^2\} = o(\|F - F'\|^2)$$

uniformly in  $k$ . Where by  $z = o(x)$  we mean  $z$  is of the order of  $x$ .

*Proof.* By writing  $X_k = \sum_{i=0}^{k-1} F^i W_{k-i}$ , using uniform boundedness of  $\{Q_k\}$  and Lemma 1(ii) we have

$$\begin{aligned} E\{[\lambda^T(X_k - X'_k)]^2\} &= \sum_{i=0}^{k-1} \lambda^T (F^i - (F')^i) Q_{k-i} (F^i - (F')^i)^T \lambda \leq \\ & p \sum_{i=0}^{k-1} \lambda^T (F^i - (F')^i) (F^i - (F')^i)^T \lambda \leq Dp \|F - F'\|^2 \sum_{i=0}^{\infty} |\xi_{F,F'}|^{2i} = \\ & D' \|F - F'\|^2 \end{aligned}$$

And this concludes the proof of the lemma.

**Lemma 6.** Let  $s$  be a nonzero vector, let  $F$  satisfy A1 and A3 and let  $E\{W_k W_k^T\} = ss^T$ . Consider the system in (1) in the stationary situation, then for any fixed integer  $N$  the covariance matrix

$$V = E\{[y_k \cdots y_{k-N}]^T [y_k \cdots y_{k-N}]\}$$

is of full rank.

*Proof.* Let  $\lambda = [\lambda_1 \cdots \lambda_N]^T$  be any vector, then

$$\lambda^T V \lambda = \frac{1}{2\pi} \int_0^{2\pi} |\lambda_1 + \cdots + \lambda_N e^{j\omega(N-1)}|^2 |c^T (e^{-j\omega I} - F)^{-1} s|^2 d\omega$$

If  $\lambda^T V \lambda = 0$  then we will also have that the product under the integral is identically zero. Since the first term in this product is a trigonometric polynomial, if it is not identically zero, it can have only a finite number of zeros in the interval

$[0, 2\pi]$ . Thus if we assume that  $\lambda \neq 0$  we conclude that  $c^T(e^{-j\omega I} - F)^{-1}s$  must be identically zero. But we have that

$$c^T(e^{-j\omega I} - F)^{-1}s = \sum_{i=0}^{\infty} e^{j\omega(i+1)} c^T F^i s$$

The above can be identically zero only if  $c^T F^i s = 0$  for every  $i$ . This is not possible since because of A3 we have that for at least one  $i$  between zero and  $m-1$  we have  $c^T F^i s \neq 0$ , contradiction. Thus  $V$  is of full rank.

With the next theorem we prove the Law of Large Numbers for expressions that appear in the proof of the CLT.

**Theorem 1.** Let  $F, W_k$  satisfy A1 and A2, let  $X_k$  be the state vector of a system of the form of (1). If  $\{a_k\}, \{b_k\}$  are two uniformly bounded real vector sequences, then for any fixed integer  $j$

$$A_n = \frac{1}{n} \sum_{k=1}^n b_k^T [X_{k+j} X_k^T - E\{X_{k+j} X_k^T\}] a_k \rightarrow 0 \quad \text{a.s.}$$

**Proof.** Since  $X_{k+j} = F^j X_k + \sum_{i=1}^j F^{j-i} W_{k+i}$  and  $W_{k+i}$  is independent of  $X_k$

for  $i \geq 1$  we have

$$A_n = \sum_{i=1}^j \frac{1}{n} \sum_{k=1}^n b_k^T F^{j-i} W_{k+i} X_k^T a_k + \frac{1}{n} \sum_{k=1}^n b_k^T F^j [X_k X_k^T - E\{X_k X_k^T\}] a_k \quad (13)$$

The sums of the form  $\sum_{k=1}^n b_k^T F^{j-i} W_{k+i} X_k^T a_k$  are martingales. Following Feller ([6], page 243) in order to show that these sums normalized by  $n$  go a.s. to zero, it is enough to show that

$$\sum_{n=1}^{\infty} \frac{E\{[b_n^T F^{j-i} W_{n+i} X_n^T a_n]^2\}}{n^2} < \infty \quad (14)$$

But (14) is true because

$$E\{[b_n^T F^{j-1} W_{n+i} X_n^T a_n]^2\} = E\{[b_n^T F^{j-1} W_{n+i}]^2\} E\{[X_n^T a_n]^2\}$$

and from Lemma 2 we have that both terms in the above product are uniformly bounded. Thus the first  $j$  terms in (13) go to zero a.s. To show that this is also true for the last term we will use the theory of mixingales (see Hall [7], page 41). If  $\{\mathcal{F}_n\}_{n \geq 1}$  is a sequence of  $\sigma$ -algebras with  $\mathcal{F}_n$  the  $\sigma$ -algebra generated by  $\{W_1 \cdots W_n\}$  and if  $V_n = b_n^T F^j [X_n X_n^T - E\{X_n X_n^T\}] a_n$  then if we can find  $\psi_k \rightarrow 0$  and  $f_n \geq 0$  such that

1.  $\|V_n - E\{V_n / \mathcal{F}_{n+k}\}\|_2 \leq \psi_{k+1} f_n$
2.  $\|E\{V_n / \mathcal{F}_{n-k}\}\|_2 \leq \psi_k f_n$
3.  $\sum_{n=1}^{\infty} \frac{f_n}{n^2} < \infty$  and  $\psi_k = o(k^{-k}(\log k)^{-2})$

then  $\frac{1}{n} \sum_{k=1}^n V_k \rightarrow 0$  a.s. For a proof of this statement see Hall ([7], page 41). We will show that we can define  $f_n$  and  $\psi_k$  to satisfy conditions 1 through 3. First notice that 1 is trivially satisfied for any nonnegative  $\psi_k, f_n$  because  $E\{V_n / \mathcal{F}_{n+k}\} = V_n$ . Let us now for simplicity denote  $d_n^T = b_n^T F^j$ , then  $d_n$  is also uniformly bounded. We can see that

$$E\{V_n / \mathcal{F}_{n-k}\} = d_n [F^k (X_{n-k} X_{n-k}^T - E\{X_{n-k} X_{n-k}^T\}) (F^T)^k] a_n \quad (15)$$

Using the fact that the second moment of a random variable is larger than the variance we have from (15)

$$\begin{aligned} \|E\{V_n / \mathcal{F}_{n-k}\}\|_2^2 &\leq E\{[d_n^T F^k X_{n-k}]^2 [a_n^T F^k X_{n-k}]^2\} \leq \\ &E\{[d_n^T F^k X_{n-k}]^4\} + E\{[a_n^T F^k X_{n-k}]^4\} \end{aligned} \quad (16)$$

Using now Lemma 1(i) and Lemma 2 we have from (16) that there exists a constant  $D$  such that  $\|E\{V_n / \mathcal{F}_{n-k}\}\|_2^2 \leq D^2 |\xi_F|^{4k}$ . We can thus define  $f_n = D$  and  $\psi_k = |\xi_F|^{2k}$ . Clearly with this definition we have validity of conditions 2 and

3 and thus the Strong Law of Large Numbers holds also for the last term in (13).

We are now ready to prove the CLT defined by (4). This is done in the following theorem.

**Theorem 2.** Let  $\{y_k\}$  be the observation process defined by a system of the form of (1). If  $F_0$  and  $W_k$  satisfy conditions A1 through A5, then (4) holds.

**Proof.** We first show the CLT under hypothesis  $H_0$ . We will use a version of the martingale CLT, thus we will put  $U_n$  defined in (3) under a martingale form. Let  $\lambda$  be a vector with  $\lambda^T \lambda = 1$ . What we would like to show is that  $\lambda^T \Sigma_n^{-1/2} U_n \xrightarrow{d} N(0,1)$ . Notice that

$$E\{[\lambda^T \Sigma_n^{-1/2} U_n]^2\} = 1 \quad (17)$$

this is true because  $\Sigma_n$  was defined as the covariance of  $U_n$ . From (17) we see that we have the right variance required by the CLT. Dropping for simplicity the subscript "0" we have from (1)

$$y_{k-i} = c^T F^{m-i-1} X_{k-m+1} + c^T \sum_{j=0}^{m-i-2} F^j W_{k-j} \quad (18)$$

Substituting (18) in (3) and using the Cayley - Hamilton theorem we have

$$\lambda^T \Sigma_n^{-1/2} U_n = \frac{\lambda_n^T}{\sqrt{n}} \sum_{k=1}^n [c_1^T W_{k+1} + c_2^T W_k + \dots + c_m^T W_{k-m+2}] Z_k \quad (19)$$

where  $\lambda_n^T = \lambda^T \Sigma_n^{-1/2}$  and  $c_i^T = c^T [F^{i-1} - \alpha_1 F^{i-2} - \dots - \alpha_{i-1} I]$ . Let us for the moment assume that the covariance  $\Sigma_n$  is uniformly bounded away from zero for large enough  $n$ . We will prove this statement in Theorem 3. With this assumption  $\lambda_n$  is uniformly bounded. Now rearranging the sum in (19) we have

$$\lambda^T \Sigma_n^{-1/2} U_n = \frac{\lambda_n^T}{\sqrt{n}} \sum_{k=1}^n [Z_k c_1^T + \dots + Z_{k+m-1} c_m^T] W_{k+1} + o\left(\frac{1}{\sqrt{n}}\right) \quad (20)$$

Since in (20)  $W_{k+1}$  is independent with whatever is in the brackets, the sum in (20) is a martingale. Defining

$$v_k = \lambda_n^T [Z_k c_1^T + \dots + Z_{k+m-1} c_m^T] W_{k+1} \quad (21)$$

in order to show the CLT it is enough to show Lindeberg's condition for  $v_k$  and that

$$\frac{1}{n} \sum_{k=1}^n v_k^2 \rightarrow 1 \quad (22)$$

in probability (see Hall [7], page 52). Lindeberg's condition is the following, for any  $\varepsilon > 0$  we want

$$\sum_{k=1}^n \Pr \left\{ \frac{v_k^2}{n} I(|v_k| > \varepsilon \sqrt{n}) \right\} \rightarrow 0$$

where with  $I(A)$  we denote the index function of the set  $A$ . To show Lindeberg's condition we have

$$\Pr \left\{ \frac{v_k^2}{n} I(|v_k| > \varepsilon \sqrt{n}) \right\} \leq \frac{E\{v_k^4\}}{\varepsilon^4 n^2}$$

Applying now Schwarz inequality at (21) yields  $v_k^2 \leq [(\lambda_n^T Z_k)^2 + \dots + (\lambda_n^T Z_{k+m-1})^2] [(c_1^T W_{k+1})^2 + \dots + (c_m^T W_{k+1})^2]$ . And thus using Lemma 2 and the independence of  $W_{k+1}$  with the  $Z_{k+i}$  vectors, we can easily see that  $E\{v_k^4\}$  is uniformly bounded by a constant  $D$ . This yields

$$\sum_{k=1}^n \Pr \left\{ \frac{v_k^2}{n} I(|v_k| > \varepsilon \sqrt{n}) \right\} \leq \frac{D}{\varepsilon^4 n} \rightarrow 0$$

and thus Lindeberg's condition holds. To show (22) since using (17) and (20) we have  $\frac{1}{n} \sum_{k=1}^n E\{v_k^2\} = E\{[\lambda^T \Sigma_n^{-1/2} U_n]^2\} + o\left(\frac{1}{n}\right) = 1 + o\left(\frac{1}{n}\right)$ , it is enough to show that  $\frac{1}{n} \sum_{k=1}^n [v_k^2 - E\{v_k^2\}] \rightarrow 0$  in probability. The random variable  $v_k^2$  is a finite



sum of terms of the form  $b_n y_{k-i} y_{k-j} (c_p^T W_{k+1})(c_q^T W_{k+1})$  where  $b_n$  depends on the vector  $\lambda_n$ . Thus we can write

$$\begin{aligned} \frac{b_n}{n} \sum_{k=1}^n [y_{k-i} y_{k-j} (c_p^T W_{k+1})(c_q^T W_{k+1}) - E\{y_{k-i} y_{k-j}\} E\{(c_p^T W_{k+1})(c_q^T W_{k+1})\}] = \\ \frac{b_n}{n} \sum_{k=1}^n y_{k-i} y_{k-j} [(c_p^T W_{k+1})(c_q^T W_{k+1}) - E\{(c_p^T W_{k+1})(c_q^T W_{k+1})\}] + \\ \frac{b_n}{n} \sum_{k=1}^n E\{(c_p^T W_{k+1})(c_q^T W_{k+1})\} [y_{k-i} y_{k-j} - E\{y_{k-i} y_{k-j}\}] \end{aligned} \quad (23)$$

The first sum in (23) is a martingale. Using similar reasoning as in Theorem 1 we can show that it converges to zero a.s. For the second term in (23) we apply Theorem 1 and thus this term also goes to zero a.s. With these arguments we have shown that (22) is true not only in probability but a.s.

Up to this point we have shown the CLT under  $H_0$ . To prove it also under  $H_1$ , notice that the observation sequence  $\{y_k\}$  is now generated not by the nominal matrix  $F_0$  but by the matrix  $F$  that satisfies  $\|F - F_0\| = o(\frac{1}{\sqrt{n}})$ . Let us call  $\{y_k^0\}$  the observation sequence generated by the nominal matrix  $F_0$  when as input we have exactly the same like the system with  $F$ . If we show for every vector  $\lambda$  with  $\lambda^T \lambda = 1$  that we have

$$E\left\{ \left[ \frac{1}{\sqrt{n}} \sum_{k=1}^n (y_{k+1} - \alpha^T Y_k) \lambda^T Z_k - \frac{1}{\sqrt{n}} \sum_{k=1}^n (y_{k+1}^0 - \alpha_0^T Y_k^0) \lambda^T Z_k^0 \right]^2 \right\} \rightarrow 0 \quad (24)$$

where  $Y_k^0, Z_k^0$  are the  $Y_k, Z_k$  vectors corresponding to the nominal system, then this means that the two terms in (24) have the same asymptotic distribution. Since the second term, as it was shown in the first part of this theorem, is asymptotically Gaussian, the same will be true for the first term as well. To show (24) define

$$\omega_k = (y_{k+1} - \alpha^T Y_k) \lambda^T Z_k - (y_{k+1}^o - \alpha_o^T Y_k^o) \lambda^T Z_k^o$$

The terms  $y_{k+1} - \alpha^T Y_k$  and  $y_{k+1}^o - \alpha_o^T Y_k^o$  are the MA parts of the two observation processes and since they are at most  $m$ -dependent we have that  $E\{\omega_k \omega_{k-j}\} = 0$  for  $j > m$ . Thus (24) is equivalent to

$$\frac{1}{n} E\left\{\left[\sum_{k=1}^n \omega_k\right]^2\right\} = \frac{1}{n} \sum_{k=1}^n E\{\omega_k^2\} + \frac{2}{n} \sum_{j=1}^{m-1} \sum_{k=1}^n E\{\omega_k \omega_{k-j}\} \rightarrow 0 \quad (25)$$

Since  $2|\omega_k \omega_{k-j}| \leq \omega_k^2 + \omega_{k-j}^2$ , in order to prove (25) it is enough to prove that the first term in (25) goes to zero. Notice that we have

$$\omega_k = (y_{k+1} - \alpha^T Y_k) [\lambda^T (Z_k - Z_k^o)] + [(y_{k+1} - \alpha^T Y_k) - (y_{k+1}^o - \alpha_o^T Y_k^o)] \lambda^T Z_k^o$$

The two MA parts are independent of  $Z_k$  and  $Z_k^o$ , thus

$$\begin{aligned} E\{\omega_k^2\} &\leq 2E\{(y_{k+1} - \alpha^T Y_k)^2\} E\{[\lambda^T (Z_k - Z_k^o)]^2\} + \\ &2E\{[\lambda^T Z_k^o]^2\} E\{[(y_{k+1} - \alpha^T Y_k) - (y_{k+1}^o - \alpha_o^T Y_k^o)]^2\} \end{aligned} \quad (26)$$

Using Lemma 5 the quantities  $y_{k+1}$ ,  $Y_k$  and  $Z_k$  differ from the corresponding nominal quantities (in the mean square sense) by an amount  $o(\|F - F_o\|^2) = o(\frac{1}{n})$ . Thus for every  $k \leq n$  we have  $E\{\omega_k^2\} = o(\frac{1}{n})$  which yields  $\frac{1}{n} \sum_{k=1}^n E\{\omega_k^2\} = o(\frac{1}{n})$ . And thus we have shown that (24) is true. What is

left now to prove in order for Theorem 2 to be complete is that the covariance matrix  $\Sigma_n$  is uniformly bounded away from zero. This is shown in Theorem 3.

**Theorem 3.** Let  $F_o, W_k$  satisfy A1 through A5. Let  $\Sigma_n$  be the covariance matrix of  $U_n$  defined in (3) and  $H_n$  the matrix defined in (5), then for large enough  $n$  the matrix  $\Sigma_n$  is uniformly of full rank and uniformly bounded from above and  $H_n$  is a.s. uniformly of full rank.

*Proof.* The proof is based on certain properties that hold under stationarity. From Equation (20) we have that

$$U_n = \frac{1}{\sqrt{n}} \sum_{k=1}^n [Z_k c_1^T + \dots + Z_{k+m-1} c_m^T] W_{k+1} + o\left(\frac{1}{\sqrt{n}}\right)$$

We first prove that  $\Sigma_n$  is full rank. Let  $\lambda$  be a vector with  $\lambda^T \lambda = 1$ . Let us call  $t_k = \lambda^T Z_k$  and  $T_k^T = [t_k \ t_{k+1} \ \dots \ t_{k+m-1}]$ , then

$$\lambda^T U_n = \frac{1}{\sqrt{n}} \sum_{k=1}^n T_k^T \bar{O} W_{k+1} + o\left(\frac{1}{\sqrt{n}}\right)$$

where  $\bar{O}^T = [c_1 \ \dots \ c_m]$ . Using A4 and that  $W_{k+1}$  is independent of  $T_k$ , we have

$$E\{[\lambda^T U_n]^2\} \geq \frac{1}{n} \sum_{k=1}^n E\{[T_k^T \bar{O} s]^2\} + o\left(\frac{1}{n}\right) = \frac{1}{n} \sum_{k=1}^n E\{[T_k^T l]^2\} + o\left(\frac{1}{n}\right) \quad (27)$$

where  $l = \bar{O} s$ . Notice that  $\bar{O}$  comes from the matrix  $O$  defined in A3 by using linear operations. Since  $O$  is assumed full rank so is  $\bar{O}$ , thus the vector  $l$  is nonzero. Let  $l^T = [l_1 \ \dots \ l_m]$  and  $\lambda^T = [\lambda_1 \ \dots \ \lambda_M]$ , then we can see that

$$T_k^T l = [\lambda_1 \ \dots \ \lambda_M] \begin{bmatrix} l_m & l_{m-1} & \dots & l_1 & 0 & \dots & 0 \\ 0 & l_m & \dots & l_2 & l_1 & \dots & 0 \\ \cdot & \cdot & \cdot & \cdot & \cdot & \cdot & \cdot \\ \cdot & \cdot & \cdot & \cdot & \cdot & \cdot & \cdot \\ \cdot & \cdot & \cdot & \cdot & \cdot & \cdot & \cdot \\ 0 & 0 & \dots & 0 & l_m & \dots & l_1 \end{bmatrix} \begin{bmatrix} y_k \\ y_{k-1} \\ \cdot \\ \cdot \\ \cdot \\ y_{k-m-M+2} \end{bmatrix} \quad (28)$$

Denote with  $L$  the matrix in (28). This matrix is of dimensions  $M \times (M+m-1)$ . It is easy to see that  $\lambda^T L$  gives the convolution of the two sequences  $\{\lambda_1, \dots, \lambda_M\}$  and  $\{l_m, \dots, l_1\}$ . It is known that the convolution of two sequences cannot be identically zero unless one of the two sequences is. Since  $l \neq 0$  this means that there is no  $\lambda \neq 0$  such that  $\lambda^T L = 0$  or, that  $L$  is of full rank. Let us call  $\sigma_L > 0$  the smallest singular value of  $LL^T$ . Now define

$$\bar{V}_n = \frac{1}{n} \sum_{k=1}^n E\{[y_k \cdots y_{k-m-M+2}]^T [y_k \cdots y_{k-m-M+2}]\} \quad (29)$$

the average covariance matrix of the random vector in (28). From Lemma 3  $\bar{V}_n$  can be approximated (order  $\frac{1}{n}$ ) by the corresponding covariance of a stationary system with input  $\bar{Q}_n$ . Since from A4 we also have  $\bar{Q}_n \geq ss^T$ , using Lemma 4 this last covariance can be lower bounded by the corresponding covariance  $V$  of a stationary system with input  $ss^T$ . The covariance  $V$  from Lemma 6 is of full rank. Call  $\sigma_V > 0$  the smallest singular value of  $V$ . Going back to (27) and using (28) and (29) we have

$$E\{[\lambda^T U_n]^2\} \geq \lambda^T L \bar{V}_n L^T \lambda \geq \sigma_V \sigma_L + o\left(\frac{1}{n}\right) \quad (30)$$

and thus for large enough  $n$ ,  $\Sigma_n$  is uniformly bounded away from zero. Notice that for the proof of the nonsingularity of  $\Sigma_n$  we did not use A5. To prove that  $\Sigma_n$  uniformly bounded from above we proceed in a similar way but we use the uniform upper bounds of the covariance matrices involved. We omit the details.

To prove that  $H_n$  is a.s. of full rank notice that because of Theorem 1 we have that  $H_n - E\{H_n\} \rightarrow 0$  a.s., thus it is enough to show that  $E\{H_n\}$  is of full rank. Using Lemma 4,  $E\{H_n\}$  can be approximated by the corresponding matrix  $H$  of a stationary system with input  $\bar{Q}_n$ . We will show that this matrix  $H$  is of full rank. We have that  $H = E\{Y_k Z_k^T\}$ , where  $Y_k$  and  $Z_k$  are now generated by a stationary system with input  $\bar{Q}_n$  (assumed constant). We would like to show that there exists real  $\sigma > 0$  such that for any vector  $\lambda$  with  $\lambda^T \lambda = 1$  we have  $\lambda^T H H^T \lambda \geq \sigma$ . Since  $Z_k$  has length  $M \geq m$  we consider only the first  $m$  components of this vector, that is, we will show that  $E\{Y_k Y_{k-m+1}^T\}$  is of full rank. Consider the vector  $E\{\lambda^T Y_k Y_{k-m+1}^T\}$ ; using similar ideas as in [1,10], the  $i$ -th component of this vector (say  $d_i$ ) is given by

$$d_i = \frac{1}{2\pi j} \int (\lambda_1 + \dots + \lambda_m z^{-m+1}) c^T (z^{-1}I - F_0)^{-1} \bar{Q}_n (zI - F_0)^{-1} c z^{m-2+i} \frac{dz}{z} =$$

$$\frac{1}{2\pi j} \int \lambda(z) c^T (z^{-1}I - F_0)^{-1} \bar{Q}_n (zI - F_0)^{-1} c z^{i-1} \frac{dz}{z} \quad i = 1, \dots, m \quad (31)$$

where  $\lambda(z) = \lambda_1 z^{m-1} + \dots + \lambda_m$  and the integration path is the unit circle. The only poles inside the unit circle in (31) are the eigenvalues  $\tau_i$  of  $F_0$ . Thus from (31)

$$d_i = \sum_{k=1}^m \frac{\lambda(\tau_k)}{\tau_k} \mu_k \tau_k^{i-1} \quad i = 1, \dots, m \quad (32)$$

where  $\mu_k = \lim_{z \rightarrow \tau_k} \{(z - \tau_k) c^T (z^{-1}I - F_0)^{-1} \bar{Q}_n (zI - F_0)^{-1} c\}$ . Since the Vandermonde matrix

$$V_r = \begin{bmatrix} 1 & \tau_1 & \dots & \tau_1^{m-1} \\ \vdots & \vdots & \ddots & \vdots \\ \vdots & \vdots & \ddots & \vdots \\ 1 & \tau_m & \dots & \tau_m^{m-1} \end{bmatrix}$$

is of full rank, call  $\sigma_r$  its smallest singular value. Notice that the vector  $[\lambda(\tau_1) \dots \lambda(\tau_m)]$  can be written as  $\bar{\lambda}^T V_r$ , where  $\bar{\lambda}^T = [\lambda_m \dots \lambda_1]$ . Let us call

$$\rho^T = \left[ \frac{\lambda(\tau_1)}{\tau_1} \mu_1 \dots \frac{\lambda(\tau_m)}{\tau_m} \mu_m \right],$$

then using A5, i.e. that  $|\mu_k| \geq \delta$  and also from

A1 that  $|\tau_k| < 1$  we have

$$\sum_{i=1}^m d_i^2 = \rho^T V_r (V_r^T \rho)^* \geq \sigma_r^2 \rho^T \rho^* \geq \sigma_r^2 \delta^2 \sum_{k=1}^m |\lambda(\tau_k)|^2 =$$

$$\sigma_r^2 \delta^2 \bar{\lambda}^T V_r^T V_r \bar{\lambda} \geq \sigma_r^4 \delta^2 \bar{\lambda}^T \lambda$$

where by the superscript  $*$  we mean complex conjugate. And this concludes the proof. Since as we have proved  $\Sigma_n$  is uniformly bounded from above and  $H_n$  is uniformly of full rank we have that  $H_n^T \Sigma_n^{-1} H_n$  is also uniformly bounded away

from zero.

#### IV. CONCLUSION.

We have presented a method for detecting changes in the AR parameters of a nonstationary ARMA process. The detection scheme was derived by using the same idea that is used for the estimation of these parameters i.e. that the MA part of an ARMA process is independent from delayed enough observations. Following a local asymptotic approach the detection of a change in the AR parameters was reduced to the detection of a nonzero mean of a Gaussian random vector. To comment now the assumptions that were made, A1 can be relaxed to include multiple eigenvalues. Assumption A2 is quite strong, it can be probably relaxed to conditions involving only second order moments. A3 seems necessary. For the other two assumptions A4 and A5, it is known that in the stationary case it is required only A5 in order to ensure that  $\Sigma_n$  and  $H_n$  are of full rank. Unfortunately this is not true for nonstationary situations. One can find counter examples where with only A5 the covariance matrix  $\Sigma_n$  is singular. Thus, it seems that an assumption of the form of A4 is also necessary. Finally to say a few things about estimating the covariance matrix  $\Sigma_n$ . Even though  $\hat{\Sigma}_n$  defined in (8) has expectation equal to  $\Sigma_n$  it has the drawback that it is not always positive definite. This can lead to a negative test statistics  $T_n$  (Equ. (7)). For practical purposes we can for example use only the first term of  $\hat{\Sigma}_n$  which is positive definite. This method was successfully used in detecting changes in vibrating modes in linear systems [3].

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## **DETECTING CHANGES IN THE A.R. PARAMETERS OF A NONSTATIONARY A.R.M.A. PROCESS**

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