Statistical Signal Processing

Lecture 5

chapter 1: parameter estimation: deterministic parameters

- some optimality properties
- Maximum Likelihood estimation, examples
- Fischer Information Matrix
- Cramer-Rao lower bound on the MSE, example
- linear model
- asymptotic (large sample) properties
- recap: estimator properties and estimators
- simplified estimators: BLUE, (W)LS, method of moments

Asymptotic (Large Sample) Properties

- asymptotic: $n \to \infty$
- *asymptotically unbiased* : $\lim_{n\to\infty} b_n(\theta) = 0 \;, \quad \forall\theta \in \Theta$
- Example (mean and variance of Gaussian i.i.d. variables):

$$E[\widehat{\sigma^2}_{ML}|\mu,\sigma^2] \;=\; \frac{n-1}{n}\,\sigma^2$$

$$b_n \;=\; \frac{n-1}{n}\,\sigma^2 - \sigma^2 \;=\; -\frac{\sigma^2}{n} \xrightarrow{n\to\infty} 0$$

$\widehat{\sigma^2}_{ML}$: biased but asymptotically unbiased

- *consistency* : convergence of (a series of random vectors:) $\widehat{\theta}_n \;\to \theta$
  - convergence in probability
  - mean square convergence
  - convergence with probability one
  - convergence in distribution

Consistency

the sequence of estimates $\widehat{\theta}(Y_n)$ is said to be

- *simply* or *weakly consistent* if

$$\lim_{n\to\infty} \Pr_{Y_n|\theta}\left\{||\widehat{\theta}(Y_n) - \theta|| < \epsilon\right\} = 1\,, \quad \forall\epsilon > 0,\ \forall\theta \in \Theta$$

- *mean-square consistent* if

$$\lim_{n\to\infty} \mathrm{MSE}_n = \lim_{n\to\infty} E_{Y_n|\theta}||\widehat{\theta}(Y_n) - \theta||^2 = 0\,, \quad \forall\theta \in \Theta$$

- *strongly consistent* if

$$\Pr_{Y_\infty|\theta}\{\lim_{n\to\infty} \widehat{\theta}(Y_n) = \theta\} = 1\,, \quad \forall\theta \in \Theta$$

- Any of these 3 consistencies implies asymptotic unbiasedness. E.g. for mean-square:

$$\underbrace{E_{Y_n|\theta}||\widehat{\theta}(Y_n) - \theta||^2}_{\text{MSE}} = ||\underbrace{E_{Y_n|\theta}\widehat{\theta}(Y_n) - \theta}_{\text{bias}}||^2 + \underbrace{E_{Y_n|\theta}||\widehat{\theta}(Y_n) - E_{Y_n|\theta}\widehat{\theta}||^2}_{\text{variance}} \to 0$$

$$\Rightarrow \lim_{n\to\infty} E_{Y_n|\theta}\widehat{\theta}(Y_n) = \theta$$

Consistency (2)

- Strong and mean-square consistency do not imply each other in general. Either implies weak consistency (e.g. use the Chebyshev inequality to show that mean-square consistency implies weak consistency), but not conversely. Except when Θ is bounded: then weak consistency implies mean-square consistency.

- example: i.i.d. $y_i \sim \mathcal{N}(\mu, \sigma^2)$, $\theta = \mu$, σ^2 known. $\widehat{\mu}_{ML} = \overline{y}$

$$Var(\widehat{\mu}_{ML}) = \frac{\sigma^2}{n} \stackrel{n\to\infty}{\longrightarrow} 0 \text{ mean-square consistent}$$

- example: i.i.d. $y_i \sim U[\theta - \frac{1}{2}, \theta + \frac{1}{2}]$, $\widehat{\theta}_{ML} = \frac{y_{min} + y_{max}}{2}$

$$\left\{ \begin{array}{lll} y_{min} & \to \theta - \frac{1}{2} & \text{in probability} \\ y_{max} & \to \theta + \frac{1}{2} & \text{in probability} \end{array} \right. \quad \text{weak consistency}$$

$$\widehat{\theta}_{ML} \to \theta \quad \text{in probability}$$

mean-square consistency can also be shown

Asymptotic Normality

- if $\widehat{\theta}_n$ consistent, then $\widetilde{\theta} \to 0$ in some sense
- introduce a magnifying glass: $d_n(\widehat{\theta}_n - \theta)$ where $0 < d_{n-1} \leq d_n \to \infty$
- *convergence in distribution*: weaker than the 3 forms of convergence of sequences of random vectors mentioned before
- if $d_n(\widehat{\theta}_n - \theta) \stackrel{in\,dist.}{\longrightarrow} \xi$, some random vector, then the distribution of ξ useful as a measure for the limiting behavior of $\widehat{\theta}_n$
- usually $\quad d_n = \sqrt{n}$
- $\widehat{\theta}_n$ *consistent asymptotically normal* (CAN) :
  if $\widehat{\theta}_n$ simply consistent and $d_n(\widehat{\theta}_n - \theta) \stackrel{in\,dist.}{\longrightarrow} \mathcal{N}(0, \Xi(\theta))$
  CAN implies asympt. unbiased (which requires that bias $\longrightarrow\ 0$ faster than $\frac{1}{d_n}$),
  Ξ = asymptotic normalized covariance of $\widehat{\theta}_n$
- distinguish $\Xi(\theta)$ from $V(\theta) = \lim_{n\to\infty} d_n^2\, C_{\tilde{\theta}\tilde{\theta}}(\theta)$ which may not even exist for a CAN estimate (if $\widehat{\theta}_n$ is simply but not mean-square consistent). $V(\theta)$ exists for a mean-square consistent $\widehat{\theta}_n$, but is not necessarily $= \Xi(\theta)$.

Asymptotic Optimality of ML

- *asymptotic normalized information matrix* : $J_0(\theta) = \lim_{n\to\infty} \frac{1}{d_n^2} J_n(\theta)$ if it exists

  ($J_0(\theta)$ = asymptotic average information per data sample y_n if $d_n = \sqrt{n}$)

- *best asymptotically normal* (BAN): CAN and $\Xi(\theta) = J_0^{-1}(\theta)$

  also called *asymptotically efficient*

- under some regularity conditions (maximum of the likelihood function unique, y_i given θ i.i.d.,...) the ML estimate is strongly consistent and BAN with $d_n = \sqrt{n}$ ($\Rightarrow$ another use of the CRB). In particular, the ML estimate is
  - asymptotically unbiased
  - asymptotically efficient (i.i.d.: $J_n = nJ_1 \;\Rightarrow\; J_0 = J_1$)
  - asymptotically normal

- example: i.i.d. $y_i \sim \mathcal{N}(\mu, \sigma^2)\,,\ \theta = \mu,\ \sigma^2$ known. $\widehat{\mu}_{ML} = \overline{y}$

$$\widehat{\mu}_{ML} \sim \mathcal{N}(\mu, \frac{\sigma^2}{n}) \;\longrightarrow\; \sqrt{n}(\widehat{\mu}_{ML}-\mu) \sim \mathcal{N}(0, \sigma^2)\;,\; J_n = \frac{n}{\sigma^2} \Rightarrow J_0^{-1} = \sigma^2 = \Xi(\theta)$$

Recap: Properties of Estimators $\widehat{\theta}(Y)$

small sample (finite n):

- *bias*: $b_{\widehat{\theta}}(\theta) = E_{Y|\theta}\widehat{\theta}(Y) - \theta$ $\quad$ ($= 0,\ \forall\theta \in \Theta$: unbiased)
- *error correlation*: $R_{\widetilde{\theta}\widetilde{\theta}} = E_{Y|\theta}\left(\widehat{\theta}(Y) - \theta\right)\left(\widehat{\theta}(Y) - \theta\right)^T$

  Cramer-Rao Bound : $\widehat{\theta}$ unbiased: $R_{\widetilde{\theta}\widetilde{\theta}} = C_{\widetilde{\theta}\widetilde{\theta}} = C_{\widehat{\theta}\widehat{\theta}}$

$$C_{\widetilde{\theta}\widetilde{\theta}} \geq J^{-1}(\theta) \quad , \quad J(\theta) = -E_{Y|\theta}\frac{\partial}{\partial\theta}\left(\frac{\partial \ln f(Y|\theta)}{\partial\theta}\right)^T \quad \text{information matrix}$$

  efficient: $C_{\widetilde{\theta}\widetilde{\theta}} = J^{-1}(\theta)\ ,\ \forall\theta \in \Theta \quad \Rightarrow \quad \widehat{\theta}(Y)$ is UMVUE

large sample ($n \to \infty$):

- *asymptotically unbiased*: $\lim_{n\to\infty} b_{\widehat{\theta}}(\theta) = 0,\ \forall\theta \in \Theta$
- *consistency* (weak, in mean square, strong): $\Rightarrow$ asymptotically unbiased
- *asymptotic normality*:

$$\text{BAN}\left\{\begin{array}{l} \left.\begin{array}{l}\diamond \text{ weakly consistent}\\ \diamond \text{ asymptotically normal}\end{array}\right\}\text{ CAN}\\ \diamond \text{ asymptotically efficient}\end{array}\right.$$

Recap: Estimation Techniques

- *Uniformly Minimum Variance Unbiased Estimator* (UMVUE): complicated (via ”sufficient statistics”)
- *Maximum likelihood* (ML): $\widehat{\theta}_{ML} = \arg\max_{\theta} f(Y|\theta)$

  Qualities:

  ◇ if $\exists$ efficient $\widehat{\theta} = \widehat{\theta}_{eff}$ and $\widehat{\theta}_{ML}$ is obtained from $\dfrac{\partial \ln f(Y|\theta)}{\partial\theta} = 0$
  $\Rightarrow \ \widehat{\theta}_{eff} = \widehat{\theta}_{ML} = \widehat{\theta}_{UMVUE}$

  ◇ $\widehat{\theta}_{ML}$ = BAN

  Problems:

  ◇ what if $f(Y|\theta)$ is unknown?

  ◇ if $f(Y|\theta)$ is not concave (local maxima)

- simplified estimators:

  ◇ *Best Linear Unbiased Estimator* (BLUE) $\rightarrow$ linear model

  ◇ *Method of Moments*

  ◇ *Least-Squares* (LS) $\rightarrow$ linear model

Best Linear Unbiased Estimator (BLUE)

- deterministic analog of LMMSE in the Bayesian case
- *linear*: $\widehat{\theta}(Y) = F\,Y \quad (F : m \times n)$
- *unbiased*: $E_{Y|\theta}\widehat{\theta} = F\,E\,(Y|\theta) = \theta$
- *best = minimum variance*: $\min C_{\tilde{\theta}\tilde{\theta}}$
- remarks:
  - BLUE inferior to UMVUE unless UMVUE is linear
  - generalizations: $X = g(Y) \; : \; \widehat{\theta}(Y) = F\,X = F\,g(Y)$ (linear in X)

    e.g.: linear in Y inappropriate if $\theta \neq 0$ and $E\,(Y|\theta) = 0$

Example of $X = g(Y)$

- $y_i \sim \mathcal{N}(0, \sigma^2)$ i.i.d., $\theta = \sigma^2$, $Y = \begin{bmatrix} y_1 \\ \vdots \\ y_n \end{bmatrix}$

- linear: $\widehat{\sigma^2} = F\,Y \;\Rightarrow\; E_{Y|\sigma^2}\widehat{\sigma^2} = F\,E\left(Y|\sigma^2\right) = 0 \neq \sigma^2$
  no linear unbiased estimator $\widehat{\sigma^2}$ exists

- however, let $x_i = y_i^2$, $X = \begin{bmatrix} y_1^2 \\ \vdots \\ y_n^2 \end{bmatrix}$

- $\widehat{\sigma^2} = F\,X \;\Rightarrow\; E_{Y|\sigma^2}\widehat{\sigma^2} = F\,E\left(X|\sigma^2\right) = \sigma^2\,F\,\mathbf{1} = \sigma^2 \;\Rightarrow\; F\,\mathbf{1} = 1$

- for this problem: $\widehat{\sigma^2}_{UMVUE} = \frac{1}{n}\mathbf{1}^T X = \widehat{\sigma^2}_{BLUE}$ $(F = \frac{1}{n}\mathbf{1}^T)$

BLUE Assumptions

- unbiased: $F\,E\,(Y|\theta) = \theta\ , \quad \forall\theta \in \Theta$

  unbiasedness and the requirement that a large class of linear unbiased estimators (many F satisfying $F\,E\,(Y|\theta) = \theta$) should exist naturally lead to:

- *assumption 1* : $E\,(Y|\theta) \;=\; H\,\theta \quad , \quad$ ($H : n \times m$)

  unbiasedness $\;\rightarrow\;\; F\,H = I_m \quad (\Rightarrow\;\; n \geq m)$

- variance:

$$\begin{aligned} C_{\tilde{\theta}\tilde{\theta}} = C_{\widehat{\theta}\widehat{\theta}} &= E_{Y|\theta}\left(\widehat{\theta} - E_{Y|\theta}\widehat{\theta}\right)\left(\widehat{\theta} - E_{Y|\theta}\widehat{\theta}\right)^T \\ &= E_{Y|\theta}\left(F\,Y - F\,E\,(Y|\theta)\right)\left(F\,Y - F\,E\,(Y|\theta)\right)^T \\ &= F\,E_{Y|\theta}\left(Y - E\,(Y|\theta)\right)\left(Y - E\,(Y|\theta)\right)^T F^T = F\,C_{YY}(\theta)\,F^T \end{aligned}$$

- *assumption 2* : $C_{YY}(\theta) = c(\theta)\,C$

  $c(\theta)$ (> 0, $\forall\theta$) is a scalar function of θ, $C > 0$ is constant w.r.t. θ

BLUE Optimization Problem

- $\min_{\widehat{\theta}:E_{Y|\theta}\widehat{\theta}(Y)=\theta} C_{\widetilde{\theta}\widetilde{\theta}} \quad \rightarrow \quad \min_{F:\,FH=I} F\,C\,F^T$

- introduce matrix square root $B\,(n \times n)$ of $C = C^T > 0\,(n \times n)$: $C = B\,B^T$
  notation: $B = C^{1/2}$, $C^{T/2} = (C^{1/2})^T$, $C = C^{1/2}C^{T/2}$, $C^{-1} = C^{-T/2}C^{-1/2}$

- Consider a vector space of $m \times n$ matrices with matrix inner product $\langle X_1, X_2 \rangle = X_1 X_2^T$. Take $X_1 = H^T C^{-T/2}$, $X_2 = F\,C^{1/2}$. With $FH = I$:

$$\left\langle \begin{bmatrix} X_1 \\ X_2 \end{bmatrix}, \begin{bmatrix} X_1 \\ X_2 \end{bmatrix} \right\rangle = \begin{bmatrix} H^T C^{-T/2} \\ F\,C^{1/2} \end{bmatrix} \begin{bmatrix} H^T C^{-T/2} \\ F\,C^{1/2} \end{bmatrix}^T = \begin{bmatrix} H^T C^{-1} H & I \\ I & FCF^T \end{bmatrix} = \begin{bmatrix} R_{11} & R_{12} \\ R_{21} & R_{22} \end{bmatrix} \geq 0$$

- From the Schur Complements Lemma, $R_{22} \geq R_{21} R_{11}^{-1} R_{12}$ with equality iff $X_2 = R_{21} R_{11}^{-1} X_1$.

- Hence $\min_{F:\,FH=I} F\,C\,F^T = \left(H^T C^{-1} H\right)^{-1}$ for $F = \left(H^T C^{-1} H\right)^{-1} H^T C^{-1}$.

- Or $\widehat{\theta}_{BLUE} = \left(H^T C^{-1} H\right)^{-1} H^T C^{-1} Y = \left(H^T C_{YY}^{-1} H\right)^{-1} H^T C_{YY}^{-1} Y$
  with $C_{\widetilde{\theta}\widetilde{\theta}} = F\,C_{YY}\,F^T = c(\theta)\,F\,C\,F^T = c(\theta)\left(H^T C^{-1} H\right)^{-1} = \left(H^T C_{YY}^{-1} H\right)^{-1}$

BLUE: Example Cont'd and Recap

Example Cont'd:

- $y_i \sim \mathcal{N}(0,\sigma^2)$ i.i.d., $\theta = \sigma^2$, $x_i = y_i^2$, $\widehat{\sigma^2} = F\,X$
- BLUE assumptions OK: $E\left(X|\sigma^2\right) = \mathbf{1}\,\sigma^2 = H\,\theta$, $C_{XX} = 2\sigma^4\,I = c(\theta)\,C$
- $\widehat{\sigma^2}_{BLUE} = \left(H^T C^{-1} H\right)^{-1} H^T C^{-1} X = \frac{1}{n}\mathbf{1}^T X$

  $C_{\widehat{\sigma^2}\widehat{\sigma^2}}(\sigma^2) = \left(H^T C_{XX}^{-1} H\right)^{-1} = \dfrac{2\sigma^4}{n}$
- note: this example is not a linear model!

Recap: BLUE assumptions:

- $\begin{cases} (1)\ E(Y|\theta) = H\,\theta \\ (2)\ C_{YY}(\theta) = c(\theta)\,C \end{cases}$

  Only need to know the first two moments of $f(Y|\theta)$ which need to satisfy these assumptions. The higher-order moments of $f(Y|\theta)$: don't need to know, can be arbitrary functions of θ. So the problem should more or less look like a linear model problem, up to the second-order moments.

BLUE: Linear Model

- $Y = H\,\theta + V\;,\;EV = 0\,,\;EVV^T = C_{VV}$

  (EV and C_{VV} independent of θ, only first two moments of V specified)

- BLUE assumptions satisfied:

  $$\begin{cases} E\,(Y|\theta) \;=\; H\,\theta \\ C_{YY}(\theta) = E_{Y|\theta}\,(Y{-}E\,(Y|\theta))\,(Y{-}E\,(Y|\theta))^T = E_V VV^T = C_{VV} = C\;(c(\theta)=1) \end{cases}$$

- $\widehat{\theta}_{BLUE} = \left(H^TC_{VV}^{-1}H\right)^{-1}H^TC_{VV}^{-1}Y$ with $C_{\tilde{\theta}\tilde{\theta}} = \left(H^TC_{VV}^{-1}H\right)^{-1}$

- If $V \sim \mathcal{N}(0, C_{VV})$ then $\widehat{\theta}_{BLUE} = \widehat{\theta}_{ML} =$ efficient $\Rightarrow = \widehat{\theta}_{UMVUE}$

Method of Moments

Principle:

- m unknown parameters $\theta = \begin{bmatrix} \theta_1 \\ \vdots \\ \theta_m \end{bmatrix}$
- $f(Y|\theta)$ depends on $\theta \quad \Rightarrow \quad$ its moments also
- take m moments $\mu = g(\theta) = \begin{bmatrix} g_1(\theta) \\ \vdots \\ g_m(\theta) \end{bmatrix} = \begin{bmatrix} \mu_1 \\ \vdots \\ \mu_m \end{bmatrix}$

  such that $g(.)$ is invertible, i.e. $\theta = g^{-1}(\mu)$: can determine θ from μ.
- estimate the moments: $\widehat{\mu}$ (e.g. sample moments)
- method of moments: $\widehat{\theta}_{MM} = g^{-1}(\widehat{\mu})$

Method of Moments: Example 1

- $y_i, i = 1, \ldots, n$ i.i.d., $f(y|\theta)$ mixture distribution, θ mixture parameter

$$f(y|\theta) = (1-\theta)\phi_1(y) + \theta\phi_2(y) \;, \quad \phi_k(y) = \frac{1}{\sqrt{2\pi\sigma_k^2}} e^{-\frac{y^2}{2\sigma_k^2}}, k = 1, 2$$

- $\mu = E\left(y^2|\theta\right) = (1-\theta)\,\sigma_1^2 + \theta\,\sigma_2^2 = g(\theta) \;\Rightarrow\; \theta = g^{-1}(\mu) = \dfrac{\mu - \sigma_1^2}{\sigma_2^2 - \sigma_1^2}$

- $\widehat{\theta}_{MM} = g^{-1}(\widehat{\mu}) = \dfrac{\widehat{\mu} - \sigma_1^2}{\sigma_2^2 - \sigma_1^2} \;, \quad \widehat{\mu} = \dfrac{1}{n}\sum_{i=1}^{n} y_i^2$ sample mean squared value

- bias: $E\widehat{\theta} = \frac{1}{\sigma_2^2-\sigma_1^2} E\widehat{\mu} - \frac{\sigma_1^2}{\sigma_2^2-\sigma_1^2} = \frac{1}{\sigma_2^2-\sigma_1^2}\mu - \frac{\sigma_1^2}{\sigma_2^2-\sigma_1^2} = \theta$: unbiased

Method of Moments: Example 1 (cont'd)

- $$Var\left(\widehat{\theta}\right) = Var\left(\frac{1}{\sigma_2^2-\sigma_1^2}\widehat{\mu} - \frac{\sigma_1^2}{\sigma_2^2-\sigma_1^2}\right) = \frac{1}{\left(\sigma_2^2-\sigma_1^2\right)^2}Var\left(\widehat{\mu}\right) = \frac{1}{\left(\sigma_2^2-\sigma_1^2\right)^2}Var\left(\frac{1}{n}\sum_{i=1}^{n} y_i^2\right)$$
  $$= \frac{1}{\left(\sigma_2^2-\sigma_1^2\right)^2}\sum_{i=1}^{n} Var\left(\frac{1}{n}y_i^2\right) = \frac{1}{\left(\sigma_2^2-\sigma_1^2\right)^2}\sum_{i=1}^{n}\frac{1}{n^2}Var\left(y_i^2\right) = \frac{1}{\left(\sigma_2^2-\sigma_1^2\right)^2}\frac{1}{n}Var\left(y^2\right)$$

- $Var(y^2) = Ey^4 - (Ey^2)^2$,
  $$\begin{aligned} f(y|\theta) &= (1-\theta)\,\phi_1(y) &&+\theta\,\phi_2(y) \\ Ey^2 &= (1-\theta)\,\sigma_1^2 &&+\theta\,\sigma_2^2 \\ Ey^4 &= (1-\theta)\,3\sigma_1^4 &&+\theta\,3\sigma_2^4 \end{aligned}$$

- $$\Rightarrow \; Var(\widehat{\theta}_{MM}) = \frac{3(1-\theta)\sigma_1^4 + 3\theta\sigma_2^4 - \left[(1-\theta)\sigma_1^2 + \theta\sigma_2^2\right]^2}{n\,(\sigma_1^2-\sigma_2^2)^2} \;\overset{n\to\infty}{\longrightarrow}\; 0$$

  $\Rightarrow \; \widehat{\theta}_{MM}$ = mean-square consistent

MM Example 2: Sinusoid in White Noise

- $y_k = s_k + v_k = A\,\cos(\omega k + \phi) + v_k\,, \quad k = 1,\ldots,n$

$$Y = \begin{bmatrix} y_1 \\ \vdots \\ y_n \end{bmatrix},\, S = \begin{bmatrix} s_1 \\ \vdots \\ s_n \end{bmatrix},\, V = \begin{bmatrix} v_1 \\ \vdots \\ v_n \end{bmatrix},\, \theta = \begin{bmatrix} A \\ \omega \\ \sigma_v^2 \end{bmatrix},\, \Theta : A > 0, \omega \in [0,\pi], \sigma_v^2 > 0$$

- distributions: $\phi \sim \mathcal{U}[0,2\pi]$ independent of θ, V; $EV = 0$, $EVV^T = \sigma_v^2 I_n$

  randomness: $f(Y,\phi|\theta) = f(\phi|\theta)\, f(Y|\theta,\phi) = f(\phi)\, f_{\mathbf{V}|\sigma_v^2}(Y - S(A,\omega,\phi)|\sigma_v^2)$

  in what follows: only first and second moments of V needed

- mean: $E_{Y,\phi|\theta}\, y_k = AE\cos(\omega k + \phi) + Ev_k = 0$

  covariance sequence:

$$\begin{aligned} r_{yy}(i) &= Ey_k y_{k+i} = A^2 E\cos(\omega k + \phi)\cos(\omega k + \phi + \omega i) \\ &\quad + AE\cos(\omega k + \phi) Ev_{k+i} + AE\cos(\omega k + \phi + \omega i) Ev_k + Ev_k v_{k+i} \\ &= \frac{A^2}{2} E\cos(2\omega k + 2\phi + \omega i) + \frac{A^2}{2} E\cos(\omega i) + \sigma_v^2 \delta_{i0} \\ &= \frac{A^2}{2}\cos(\omega i) + \sigma_v^2 \delta_{i0} \end{aligned}$$

MM Example 2: Sinusoid in White Noise (2)

- moments: $\mu = \begin{bmatrix} r_{yy}(0) \\ r_{yy}(1) \\ r_{yy}(2) \end{bmatrix} = \begin{bmatrix} \frac{A^2}{2} + \sigma_v^2 \\ \frac{A^2}{2}\cos(\omega) \\ \frac{A^2}{2}\cos(2\omega) \end{bmatrix} = g(\theta)$

- $\theta = g^{-1}(\mu)$:

$$\omega = \begin{cases} \arccos\left(\dfrac{r_{yy}(2) + \sqrt{r_{yy}^2(2) + 8r_{yy}^2(1)}}{4\,r_{yy}(1)}\right) & , \; r_{yy}(1) \neq 0 \\ \dfrac{\pi}{2} & , \; r_{yy}(1) = 0 \end{cases}$$

$$A = \begin{cases} \sqrt{\dfrac{2\,r_{yy}(1)}{\cos(\omega)}} & , \; r_{yy}(1) \neq 0 \\ \sqrt{-2\,r_{yy}(2)} & , \; r_{yy}(1) = 0 \end{cases} \quad , \quad \sigma_v^2 = r_{yy}(0) - \frac{A^2}{2}$$

- sample moments $\widehat{\mu}$: $\hat{r}_{yy}(i) = \frac{1}{n}\sum_{k=1}^{n-i} y_k y_{k+i}\,, \; i = 0, 1, 2$

Method of Moments: Properties

- $\widehat{\mu}$ easy to compute, $\widehat{\theta}_{MM} = g^{-1}(\widehat{\mu})$ straightforward if μ chosen well, hence $\widehat{\theta}_{MM}$ easy to determine and easy to implement

- no optimality properties but usually consistent (since $\widehat{\mu}$ consistent)

- if performance of $\widehat{\theta}_{MM}$ not satisfactory, can use $\widehat{\theta}_{MM}$ as initialization in an iterative optimization procedure that finds $\widehat{\theta}_{ML}$