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Clustering Longitudinal Ordinal Data

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Context

- Longitudinal ordinal data $y_{i,j,t}$ whose levels are coded $\{1, \dots, C_j\}$: the observation of the j -th variable for the i -th unit at time t ($i = 1, \dots, N$; $j = 1, \dots, J$ and $t = 1, \dots, T$).
- We want to **cluster units accounting for the temporal behavior**
- \Rightarrow  Idea: rewrite them in a three-way format and use **latent underlying continuous matrix-variate distributions!**
- We organize our data in a random-matrix form such that:

$$Y_i = \begin{pmatrix} y_{i,1,1} & \cdots & y_{i,1,t} & \cdots & y_{i,1,T} \\ \vdots & \cdots & \vdots & \cdots & \vdots \\ y_{i,j,1} & \cdots & y_{i,j,t} & \cdots & y_{i,j,T} \\ \vdots & \cdots & \vdots & \cdots & \vdots \\ y_{i,J,1} & \cdots & y_{i,J,t} & \cdots & y_{i,J,T} \end{pmatrix}$$

From continuous...

- Matrix-variate Normal: $Z \sim \mathcal{MN}_{(J \times T)}(M, \Phi, \Sigma)$, where
 - $M \in \mathbb{R}^{J \times T}$ is the matrix of means
 - $\Phi \in \mathbb{R}^{T \times T}$ is a covariance matrix between the T occasions
 - $\Sigma \in \mathbb{R}^{J \times J}$ is the covariance matrix of the J variables

The matrix-normal probability density function (pdf) is given by

$$\phi^{(J \times T)}(Z|M, \Phi, \Sigma) = (2\pi)^{-\frac{TJ}{2}} |\Phi|^{-\frac{T}{2}} |\Sigma|^{-\frac{T}{2}} \exp \left\{ -\frac{1}{2} \text{tr}[\Sigma^{-1}(Z - M)\Phi^{-1}(Z - M)^\top] \right\}$$

- Mixtures of Matrix-Normals (MMN) were introduced by Viroli [3]

$$f(Y|\boldsymbol{\pi}, \boldsymbol{\Theta}) = \sum_{k=1}^K \pi_k \phi^{(J \times T)}(Z|M_k, \Phi_k, \Sigma_k),$$

where

- K : number of mixture components
- $\boldsymbol{\pi} = \{\pi_k\}_{k=1}^K$: vector of mixing proportions, $\sum_{k=1}^K \pi_k = 1$
- $\boldsymbol{\Theta} = \{\Theta_k\}_{k=1}^K$: set of component-specific parameters $\Theta_k = \{M_k, \Phi_k, \Sigma_k\}$

\Rightarrow **Advantages:** offers a parsimonious and easily interpretable way to include the time dimension in the clustering.

...to ordinal data!

In the **clustMD** framework [2] cross-sectional mixed data are assumed to be all manifestation of underlying multivariate normals, and a Gaussian mixture model operating on the underlying normal variable is used to cluster them.

As for the classical **clustMD**, we can assume that each observed ordinal matrix Y is indeed the manifestation of a latent random matrix Z , which follows a matrix-normal distribution.

$$Z_i = \begin{pmatrix} z_{i,1,1} & \cdots & z_{i,1,t} & \cdots & z_{i,1,T} \\ \vdots & \cdots & \vdots & \cdots & \vdots \\ z_{i,j,1} & \cdots & z_{i,j,t} & \cdots & z_{i,j,T} \\ \vdots & \cdots & \vdots & \cdots & \vdots \\ z_{i,J,1} & \cdots & z_{i,J,t} & \cdots & z_{i,J,T} \end{pmatrix} \longrightarrow Y_i = \begin{pmatrix} y_{i,1,1} & \cdots & y_{i,1,t} & \cdots & y_{i,1,T} \\ \vdots & \cdots & \vdots & \cdots & \vdots \\ y_{i,j,1} & \cdots & y_{i,j,t} & \cdots & y_{i,j,T} \\ \vdots & \cdots & \vdots & \cdots & \vdots \\ y_{i,J,1} & \cdots & y_{i,J,t} & \cdots & y_{i,J,T} \end{pmatrix}$$

To map from Y_i to Z_i , let $\boldsymbol{\gamma}_j$ denote a $C_j + 1$ -dimensional vector of thresholds that partition the space of the underlying latent continuous variable for the j -th ordinal variable. Let the threshold parameters be constrained such that $-\infty = \gamma_{j,0} \leq \gamma_{j,1} \leq \dots \leq \gamma_{j,C_j} = \infty$. If the latent $z_{i,j,t}$ is such that $\gamma_{j,c-1} < z_{i,j,t} < \gamma_{j,c}$ then the observed ordinal response, $y_{i,j,t} = c$.

A key point is of course the choice of the thresholds $\boldsymbol{\gamma} = \{\boldsymbol{\gamma}_j\}_{j=1}^J$. In [1], thresholds are fixed in advance to avoid identifiability and computational complexity issues. Also in [2], for ordinal variables they are fixed such that $\gamma_{j,c} = \varphi^{-1}(\delta_c)$, where δ_c is the proportion of variable J which are less than or equal to level c and φ is the normal cumulative distribution function.

Model

The model relies on the following hypotheses:

- $\boldsymbol{\ell}_i \in \{0, 1\}^K$ is the latent allocation variable such that $\ell_{ik} = 1$ if the i -th unit belongs to the k -th cluster.
- $\mathbf{Y}_i^R = (Y_{i1}^R, \dots, Y_{iR}^R) \in \mathbb{R}^R$ indicate the observed response pattern for the i -th unit.
- $\boldsymbol{\ell}_i \sim \mathcal{M}(1, \boldsymbol{\pi})$, $\boldsymbol{\pi} = (\pi_1, \dots, \pi_K)$
- $\mathbf{Z}_i | \ell_{ik} = 1 \sim \mathcal{MN}_{(J \times T)}(\mathbf{Z}_i | \Theta_k)$, $\Theta_k = \{M_k, \Phi_k, \Sigma_k\}$
- $\mathbf{Y}_i^R | \mathbf{Z}_i, \ell_{ik} = 1 \sim \mathcal{M}(1, \boldsymbol{\xi}_i^R)$, $\boldsymbol{\xi}_i^R = (\mathbf{1}_{\Omega_1}(\mathbf{Z}_i), \dots, \mathbf{1}_{\Omega_R}(\mathbf{Z}_i))$

where \mathcal{M} indicate the multinomial distribution, Ω_r is the portion of the $J \times T$ -space which determines the r -th pattern, and $\mathbf{1}_{\Omega_r}(\mathbf{Z}_i)$ is the indicator function that equals 1 when the elements in \mathbf{Z}_i have values that determine the r -th pattern. Of course, when \mathbf{Z}_i is given, the value of \mathbf{Y}_i is no more random.

We can derive the joint density of $\mathbf{Z}_i, \mathbf{Y}_i^R, \boldsymbol{\ell}_i$ as:

$$f(\mathbf{Y}_i^R, \mathbf{Z}_i, \boldsymbol{\ell}_i) = f(\mathbf{Y}_i^R | \mathbf{Z}_i, \boldsymbol{\ell}_i) f(\mathbf{Z}_i | \boldsymbol{\ell}_i) f(\boldsymbol{\ell}_i),$$

where:

$$f(\boldsymbol{\ell}_i) = \prod_{k=1}^K \pi_k^{\ell_{ik}}, f(\mathbf{Z}_i | \boldsymbol{\ell}_i) = \prod_{k=1}^K [\phi^{(J \times T)}(Z_i | \Theta_k)]^{\ell_{ik}}, f(\mathbf{Y}_i^R | \mathbf{Z}_i, \boldsymbol{\ell}_i) = \prod_{r=1}^R \mathbf{1}_{\Omega_r}(\mathbf{Z}_i)^{Y_{ir}^R}$$

Model Inference

Due to the presence of latent variables, the maximization of the likelihood cannot be done in “close form”, and we must then use an EM algorithm, which maximizes a lower limit of the log-likelihood: the complete log-likelihood. We can write the complete log-likelihood as

$$\log \mathcal{L}_C(\boldsymbol{\pi}, \boldsymbol{\Theta} | \mathbf{Y}^R, \mathbf{Z}, \boldsymbol{\ell}) = \sum_{i=1}^N \left\{ \sum_{r=1}^R Y_{ir} \mathbf{1}_{\Omega_r}(\mathbf{Z}_i) + \sum_{k=1}^K \ell_{ik} \left[\log(\pi_k) - \frac{TJ}{2} \log(2\pi) - \frac{J}{2} \log(|\Phi_k|) - \frac{T}{2} \log(|\Sigma_k|) - \frac{1}{2} \text{tr}[\Sigma_k^{-1}(\mathbf{Z}_i - M_k)\Phi_k^{-1}(\mathbf{Z}_i - M_k)^\top] \right] \right\}$$

- E-step: we compute the expectation of the complete log-likelihood with respect to the latent data Z and the cluster labels $\boldsymbol{\ell}$. For each response pattern r , **we can approximate the value of $\mathbf{Z}_i | \boldsymbol{\ell}_i$ as the expected value of the truncated multivariate normals (using properties of matrix-variate normals)**, given the parameters Θ_k of the assigned cluster.

The latent variable $\boldsymbol{\ell}_i$ can be computed by means of Bayes' theorem as:

$$\mathbb{E}(\ell_{ik}^{(s)} | Y_i^R = r, \boldsymbol{\Theta}^{(s-1)}, \boldsymbol{\pi}^{(s-1)}) = \frac{\pi_k^{(s-1)} \int_{\Omega_r} f(Z | \Theta_k^{(s-1)}) dZ}{\sum_{k=1}^K \pi_k^{(s-1)} \int_{\Omega_r} f(Z | \Theta_k^{(s-1)}) dZ},$$

which would require Monte-Carlo approximation.

- M-step: the parameter updates are given by:

$$\hat{\Sigma}_k^{(s)} = \frac{\sum_{i=1}^N \ell_{ik}^{(s)} (\mathbf{Z}_i - \hat{M}_k^{(s)}) \hat{\Phi}_k^{-1(s-1)} (\mathbf{Z}_i - \hat{M}_k^{(s)})^\top}{T \sum_{i=1}^N \ell_{ik}^{(s)}}, \quad \hat{M}_k^{(s)} = \frac{\sum_{i=1}^N \ell_{ik}^{(s)} \mathbf{Z}_i}{\sum_{i=1}^N \ell_{ik}^{(s)}}$$

$$\hat{\Phi}_k^{(s)} = \frac{\sum_{i=1}^N \ell_{ik}^{(s)} (\mathbf{Z}_i - \hat{M}_k^{(s)}) \hat{\Sigma}_k^{-1(s)} (\mathbf{Z}_i - \hat{M}_k^{(s)})}{J \sum_{i=1}^N \ell_{ik}^{(s)}}, \quad \hat{\boldsymbol{\pi}}_k^{(s)} = \frac{\sum_{i=1}^N \ell_{ik}^{(s)}}{N}$$

The E and M step are iterated until convergence of the log-likelihood.

A longitudinal clustMD?

This is the first step of a broader project, aiming at extending this framework to account for mixed data (continuous, ordinal, nominal, count) in order to cluster mixed longitudinal dataset.

References

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