**Tutorial on EM Algorithm**

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# **Abstract**

Maximum likelihood estimation (MLE) is a popular method for parameter estimation in both applied probability and statistics but MLE cannot solve the problem of incomplete data or hidden data because it is impossible to maximize likelihood function from hidden data. Expectation maximum (EM) algorithm is a powerful mathematical tool for solving this problem if there is a relationship between hidden data and observed data. Such hinting relationship is specified by a mapping from hidden data to observed data or by a joint probability between hidden data and observed data. In other words, the relationship helps us know hidden data by surveying observed data. The essential ideology of EM is to maximize the expectation of likelihood function over observed data based on the hinting relationship instead of maximizing directly the likelihood function of hidden data. Pioneers in EM algorithm proved its convergence. As a result, EM algorithm produces parameter estimators as well as MLE does. This tutorial aims to provide explanations of EM algorithm in order to help researchers comprehend it. Moreover some improvements of EM algorithm are also proposed in the tutorial such as combination of EM and third-order convergence Newton-Raphson process, combination of EM and gradient descent method, and combination of EM and particle swarm optimization (PSO) algorithm.

**Keywords**: expectation maximization, EM, generalized expectation maximization, GEM, EM convergence.

# **1. Introduction**

Literature of expectation maximization (EM) algorithm in this tutorial is mainly extracted from the preeminent article “Maximum Likelihood from Incomplete Data via the EM Algorithm” by Arthur P. Dempster, Nan M. Laird, and Donald B. Rubin (Dempster, Laird, & Rubin, 1977). For convenience, let **DLR** be reference to such three authors.

We begin a review of EM algorithm with some basic concepts. Before discussing main subjects, there are some conventions. For example, if there is no additional explanation, variables are often denoted by letters such as *x*, *y*, *z*, *X*, *Y*, and *Z* whereas values and constants are often denoted by letters such as *a*, *b*, *c*, *A*, *B*, and *C*. Parameters are often denoted as Greek letters such as *α*, *β*, *γ*, Θ, Φ, and Ψ. Uppercase letters often denote vectors and matrices (multivariate quantities) whereas lowercase letters often denote scalars (univariate quantities). Script letters such as and often denote data samples. Bold and uppercase letters such as ***X*** and ***R*** often denote algebraic structures such as spaces, fields, and domains. Moreover, bold and lowercase letters such as ***x***, ***y***, ***z***, ***a***, ***b***, and ***c*** may denote vectors. Bold and uppercase letters such as ***X***, ***Y***, ***Z***, ***A***, ***B***, and ***C*** may denote matrices.

By default, vectors are column vectors although a vector can be column vector or row vector. For example, given two vectors *X* and *Y* and two matrices *A* and *B*:

*X* and *Y* above are column vectors. A row vector is represented as follows:

The number of elements in vector is its dimension. Zero vector is denoted as **0** whose dimension depends on context.

If considering rows and columns, *m*x*n* matrix *A* can be denoted *Am*x*n* or (*aij*)*m*x*n*. Vector is 1-row matrix or 1-column matrix such as *A*1x*n* or *An*x1. Scalar is 1-element vector or 1x1 matrix. A matrix can be considered as a vector whose elements are vectors.

Let (**0**) denote zero matrix whose numbers of rows and columns depend on context. If considering rows and columns, zero matrix can be denoted (0)*m*x*n*.

Matrix *A* is square if *m* = *n*, which can be denoted *An* or (*aij*)*n*. Square matrix *A* is symmetric if *aij* = *aji* for all *i*≠*j*.

Matrix Λ is diagonal if it is square and its elements outside the main diagonal are zero:

Let *I* be identity matrix or unit matrix, as follows:

Note, *I* is diagonal and its diagonal elements are 1. The row (column) number of identity matrix depends on context, but it can be denoted explicitly as *In*.

Vector addition and matrix addition are defined like numerical addition:

(if *n* = *k*)

Vector and matrix can be multiplied with a scalar.

Let superscript “*T*” denote transposition operator for vector and matrix, as follows:

A square matrix *A* is symmetric if and only *AT* = *A*.

Transposition operator is linear with addition operator as follows:

Dot product or scalar product of two vectors can be written with transposition operator, as follows:

However, the product *XYT* results out a symmetric matrix as follows:

The length of module of vector *X* in Euclidean space is:

The notation |.| also denotes absolute value of scalar and determinant of square matrix; for example, we have |–1| = 1 and |*A*| which is determinant of given square matrix *A*. Note, determinant is only defined for square matrix. Let *A* and *B* be two square *n*x*n* matrices, we have:

|*cA*| = *cn*|*A*| where *c* is scalar

|*AT*| = |*A*|

|*AB*| = |*A*||*B*|

If *A* has nonzero determinant (≠0), there exists its inverse denoted *A*–1 such that:

Where *I* is identity matrix. If matrix *A* has its inverse, *A* is called invertible or non-singular. In general, that square matrix *A* is invertible is equivalent to the event that its determinant is nonzero (≠0). There are many documents which guide to calculate inverse of invertible matrix.

Let *A* and *B* be two invertible matrices, we have:

(*AB*)–1 = *B*–1*A*–1

|*A*–1| = |*A*|–1 = 1 / |*A*|

(*AT*)–1 = (*A*–1)*T*

Given invertible matrix *A*, it is called orthogonal matrix if *A*–1 *= AT*, which means *AA*–1 *= A*–1*A* = *AAT* = *ATA* = *I*. Of course, orthogonal matrix is symmetric.

Product (multiplication operator) of two matrices *Am*x*n* and *Bn*x*k* is:

Square matrix *A* is symmetric if *aij* = *aji* for all *i* and *j*. If *A* is symmetric then, *AT* = *A*. If both *A* and *B* are symmetric with the same number of rows and columns then, they are commutative such that *AB* = *BA* with note that the product *AB* and *BA* produces a symmetric matrix. Given invertible matrix *A*, if it is symmetric, its inverse *A*–1 is symmetric too.

Given *N* matrices *Ai* such that their product (multiplication operator) is valid, we have:

Product of matrix and vector is similar to product of matrix and matrix when vector is considered as 1-column matrix or 1-row matrix, which results a vector.

Where .

Where .

Given square matrix *A*, tr(*A*) is trace operator which takes sum of its diagonal elements.

Given invertible matrix *A* (*n* rows and *n* columns), Jordan decomposition theorem (Hardle & Simar, 2013, p. 63) stated that *A* is always decomposed as follows:

Where *U* is orthogonal matrix composed of eigenvectors. Hence, *U* is called eigenvector matrix.

There are *n* column eigenvectors ***u****i* = (*u*11, *u*12,…, *u*1*n*) in *U* and they are mutually orthogonal, ***u****iT****u****j* = 0. Where Λ is diagonal matrix composed of eigenvalues. Hence, Λ is called eigenvalue matrix.

Where *λi* are eigenvalues. When invertible matrix *A* is decomposed according to Jordan decomposition, we call *A* is diagonalized. If *A* can be diagonalized, it is called diagonalizable matrix. Of course, if *A* is invertible, *A* is diagonalizable. There are many documents for matrix diagonalization.

Given two diagonalizable matrices *A* and *B* are equal size (*n*x*n*) then, they are simultaneously diagonalizable (Wikipedia, Commuting matrices, 2017) and hence, there exists an orthogonal eigenvector matrix *U* such that (Wikipedia, Diagonalizable matrix, 2017) (StackExchange, 2013):

Where Γ and Λ are eigenvalue matrices of *A* and *B*, respectively.

Given symmetric matrix *A*, it is positive (negative) definite if and only if *XTAX* > 0 (*XTAX* < 0) for all vector *X*≠**0***T*. It is positive (negative) semi-definite if and only if *XTAX* ≥ 0 (*XTAX* ≤ 0) for all vector *X*. When diagonalizable *A* is diagonalized into *U*Λ*UT*, it is positive (negative) definite if and only if all eigenvalues in Λ are positive (negative). Similarly, it is positive (negative) semi-definite if and only if all eigenvalues in Λ are non-negative (non-positive). If *A* is degraded as a scalar, concepts “positive definite”, “positive semi-definite”, “negative definite”, and “negative semi-definite” becomes concepts “positive”, “non-negative”, “negative”, and “non-positive”, respectively.

Suppose *f*(*X*) is scalar-by-vector function, for instance, *f*: ***R****n* → ***R*** where ***R****n* is *n*-dimensional real vector space. The first-order derivative of *f*(*X*) is gradient vector as follows:

Where is partial first-order derivative of *f* with regard to *xi*. So gradient is row vector. The second-order derivative of *f*(*X*) is called Hessian matrix as follows:

Where

Obviously, Hessian matrix is square matrix. Second-order partial derivatives of *xi* (s) are on diagonal of the Hessian matrix. In general, vector calculus is a complex subject. Here we focus on scalar-by-vector function with some properties. Let *c*, *A*, *B*, and *M* be scalar constant, vector constant, vector constant, and matrix constant, respectively, suppose vector and matrix operators are valid we have:

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If *M* is a square matrix constant, we have:

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Function *f*(*X*) is called *n*th-order analytic function or *n*th-order smooth function if there is existence and continuity of *k*th-order derivatives of *f*(*X*) where *k* = 1, 2,…, *K*. Function *f*(*X*) is called smooth enough function if *K* is large enough. According to Schwarz’s theorem (Wikipedia, Symmetry of second derivatives, 2018), if *f*(*X*) is second-order smooth function then, its Hessian matrix is symmetric.

When *X* is univariate, gradient vector and Hessian matrix are degraded as scalars. Without loss of generality, by default, variable *X* in this research is multivariate as vector if there is no additional explanation.

Given *f*(*X*) being second-order smooth function, *f*(*X*) is convex (strictly convex) in domain ***X*** if and only if its Hessian matrix is semi-positive definite (positive definite) in ***X***.Similarly, *f*(*X*) is concave (strictly concave) in domain ***X*** if and only if its Hessian matrix is semi-negative definite (negative definite) in ***X***. Extreme point, optimized point, optimal point, or optimizer *X*\* is minimum point (minimizer) of convex function and is maximum point (maximizer) of concave function.

Given second-order smooth function *f*(*X*), function *f*(*X*) has stationary point *X*\* if its gradient vector at *X*\* is zero, *Df*(*X*\*) = **0***T*. The stationary point *X*\* is local minimum point if Hessian matrix at *X*\* that is *D*2*f*(*X*\*) is positive definite. Otherwise, the stationary point *X*\* is local maximum point if Hessian matrix at *X*\* that is *D*2*f*(*X*\*) is negative definite. If a stationary point *X*\* is neither minimum point nor maximum point, it is saddle point in which *Df*(*X*\*) = **0***T* and *D*2*f*(*X*\*) = (**0**) where (**0**) denotes zero matrix whose all elements are zero. Finding extreme point (minimum point or maximum point) is optimization problem. Therefore, if *f*(*X*) is second-order smooth function and its gradient vector *Df*(*X*) and Hessian matrix *D*2*f*(*X*) and are both determined, the optimization problem is processed by solving the equation created from setting the gradient *Df*(*X*) to be zero (*Df*(*X*)=**0***T*) and then checking if the Hessian matrix *Df*(*X*\*) is positive definite or negative definite where *X*\* is solution of equation *Df*(*X*)=**0***T*. If such equation cannot be solved due to its complexity, there are some popular methods to solve optimization problem such as Newton-Raphson (Burden & Faires, 2011, pp. 67-71) and gradient descent (Ta, 2014).

A short description of Newton-Raphson method is necessary because it is helpful to solve the equation *Df*(*X*)=**0***T* for optimization problem in practice, especially in case that there is no algebraic formula for solution of such equation. Suppose *f*(*X*) is second-order smooth function, according to first-order Taylor series expansion of *Df*(*X*) at *X*=*X*0 with very small residual, we have:

Because *f*(*X*) is second-order smooth function, *D*2*f*(*X*0) is symmetric matrix according to Schwarz’s theorem (Wikipedia, Symmetry of second derivatives, 2018), which implies:

*D*2*f*(*X*0) = (*D*2*f*(*X*0))*T*

So, we have:

We expect that *Df*(*X*) = **0***T* so that *X* is a solution.

It implies:

This means:

Therefore, Newton-Raphson method starts with an arbitrary value of *X*0 as a solution candidate and then goes through some iterations. Suppose at *k*th iteration, the current value is *Xk* and the next value *Xk*+1 is calculated based on following equation:

The value *Xk* is solution of *Df*(*X*)=**0***T* if *Df*(*Xk*)=**0***T* which means that *Xk*+1=*Xk* after some iterations. At that time *Xk*+1 = *Xk* = *X*\* is the local optimized point (local extreme point). So, the terminated condition of Newton-Raphson method is *Df*(*Xk*)=**0***T*. Note, the *X*\* resulted from Newton-Raphson method is local minimum point (local maximum point) if *f*(*X*) is convex function (concave function) in current domain.

Newton-Raphson method computes second-order derivative *D*2*f*(*X*) but gradient descent method (Ta, 2014) does not. This difference is not significant but a short description of gradient descent method is necessary because it is also an important method to solve the optimization problem in case that solving directly the equation *Df*(*X*)=**0***T* is too complicated. Gradient descent method is also iterative method starting with an arbitrary value of *X*0 as a solution candidate. Suppose at *k*thiteration, the next candidate point *Xk*+1 is computed based on the current *Xk* as follows (Ta, 2014):

The direction ***d****k* is called descending direction, which is the opposite of gradient of *f*(*X*). Hence, we have ***d****k* = –*Df*(*Xk*). The value *tk* is the length of the descending direction ***d****k*. The value *tk* is often selected a minimizer (maximizer) of function *g*(*t*) = *f*(*Xk* + *t****d****k*) for minimization (maximization) where *Xk* and ***d****k* are known at *k*thiteration. Alternately, *tk* is selected by some advanced condition such as Barzilai–Borwein condition (Wikipedia, Gradient descent, 2018). After some iterations, point *Xk* converges to the local optimizer *X\** when ***d****k* = **0***T*. At that time is we have *Xk*+1 = *Xk* = *X*\*. So, the terminated condition of Newton-Raphson method is ***d****k*=**0***T*. Note, the *X*\* resulted from gradient descent method is local minimum point (local maximum point) if *f*(*X*) is convex function (concave function) in current domain.

In the case that the optimization problem has some constraints, Lagrange duality (Jia, 2013) is applied to solve this problem. Given first-order smooth function *f*(*X*) and constraints *gi*(*X*) ≤ 0 and *hj*(*X*) = 0, the optimization problem is stated as follows:

A so-called Lagrange function *la*(*X*, *λ*, *μ*) is established as sum of *f*(*X*) and constraints multiplied by Lagrange multipliers *λ* and *μ*. In case of minimization problem, *la*(*X*, *λ*, *μ*) is

In case of maximization problem, *la*(*X*, *λ*, *μ*) is

Where all *λi* ≥ 0. Note, *λ* = (*λ*1, *λ*2,…, *λm*)*T* and *μ* = (*μ*1, *μ*2,…, *μm*)*T* are called Lagrange multipliers and *la*(*X*, *λ*, *μ*) is function of *X*, *λ*, and *μ*. Thus, optimizing *f*(*X*) with subject to constraints *gi*(*X*) ≤ 0 and *hj*(*X*) = 0 is equivalent to optimize *la*(*X*, *λ*, *μ*), which is the reason that this method is called Lagrange duality. Suppose *la*(*X*, *λ*, *μ*) is also first-order smooth function. In case of minimization problem, the gradient of *la*(*X*, *λ*, *μ*) with regard to *X* is

In case of maximization problem, the gradient of *la*(*X*, *λ*, *μ*) with regard to *X* is

According to KKT condition (Wikipedia, Karush–Kuhn–Tucker conditions, 2014), a local optimized point (local extreme point) *X*\* is solution of the following equation system:

The last equation in the KKT system above is called complementary slackness. The main task of KKT problem is to solve the first equation *Dla*(*X*, *λ*, *μ*) = **0***T*. Again some practical methods such as Newton-Raphson method can be used to solve the equation *Dla*(*X*, *λ*, *μ*) = **0***T*. Alternately, gradient descent method can be used to optimize *la*(*X*, *λ*, *μ*) with constraints specified in KKT system.

Let *P*(.) denote probability,

We need to skim some essential probabilistic rules such as additional rule, multiplication rule, total probability rule, and Bayes’ rule. Given two random events (or random variables) *X* and *Y*, additional rule (Montgomery & Runger, 2003, p. 33) and multiplication rule (Montgomery & Runger, 2003, p. 44) are expressed as follows:

Where notations and denote union operator and intersection operator in set theory (Wikipedia, Set (mathematics), 2014). Your attention please, when *X* and *Y* are numerical variables, notations and also denote operators “or” and “and” in theory logic (Rosen, 2012, pp. 1-12). The probability *P*(*X*, *Y*) is known as joint probability. The probability *P*(*X*|*Y*) is called conditional probability of *X* given *Y*:

Conditional probability is base of Bayes’ rule mentioned later.

If *X* and *Y* are mutually exclusive () then, is often denoted as *X*+*Y* and we have:

(Due to *P*(Ø) = 0)

*X* and *Y* are mutually independent if and only if one of three following conditions is satisfied:

When *X* and *Y* are mutually independent, are often denoted as *XY* and we have:

Given a complete set of mutually exclusive events *X*1, *X*2,…, *Xn* such that

The total probability rule (Montgomery & Runger, 2003, p. 44) is specified as follows:

If *X* and *Y* are continuous variables, the total probability rule is re-written in integral form as follows:

Note, *P*(*Y|X*) and *P*(*X*) are continuous functions known as probability density functions mentioned later. The important Bayes’ rule will also be mentioned later.

A variable *X* is called random variable if it conforms a probabilistic distribution which is specified by a probability density function (PDF) or a cumulative distribution function (CDF) (Montgomery & Runger, 2003, p. 64) (Montgomery & Runger, 2003, p. 102) but CDF and PDF have the same meaning and they share interchangeable property when PDF is derivative of CDF; in other words, CDF is integral of PDF. In practical statistics, PDF is used more common than CDF is used and so, PDF is mentioned over the whole report. When *X* is discrete, PDF is degraded as probability of *X*. Note, notation *P*(.) often denotes probability and it can be used to denote PDF but we prefer to use lower case letters such as *f* and *g* to denote PDF. Given a random variable having PDF *f*, we often state that “such variable has distribution *f* or such variable has density function *f*”. Let *F*(*X*) and *f*(*X*) be CDF and PDF, respectively, equation 1.1 is definition of CDF and PDF.

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|  | (1.1) |

In discrete case, probability at a single point *X*0 is determined as *P*(*X*0) = *f*(*X*0) but in continuous case, probability is determined in an interval [*a*, *b*], (*a*, *b*), [*a*, *b*), or (*a*, *b*] where *a*, *b*, and *X* are real as integral of the PDF in such interval as follows:

Hence, in continuous case, probability at a single point is 0.

Equation 1.1 defines CDF and PDF for univariate random variable and so it is easy to expend it for multivariate variable when *X* is vector. Let *X* = (*x*1, *x*2,…, *xn*)*T* be *n-*dimension random vector, its CDF and PDF are re-defined as follows:

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| Continuous case:  Discrete case: | (1.2) |

Marginal PDF of partial variable *xi* where *xi* is a component of *X* is the integral of *f*(*X*) over all *xj* except *xi*.

Where,

Joint PDF of *xi* and *xj* is defined as the integral of *f*(*X*) over all *xk* except *xi* and *xj*.

Where,

Conditional PDF of *xi* given *xj* is defined as follows:

Indeed, conditional PDF implies conditional probability.

Given random variable *X* and its PDF *f*(*X*), theoretical expectation *E*(*X*) and theoretical variance *V*(*X*) of *X* are:

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|  | (1.3) |

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|  | (1.4) |

The expectation *E*(*X*) of *X* is often called theoretical mean. When *X* is multivariate vector, *E*(*X*) is mean vector and *V*(*X*) is covariance matrix whichis always symmetric. When *X* = (*x*1, *x*2,…, *xn*)*T* is multivariate, *E*(*X*) and *V*(*X*) have following forms:

Therefore, theoretical means and variances of partial variables *xi* can be determined separately. For instance, each *E*(*xi*) is theoretical mean of partial variable *xi* given marginal PDF .

Each *V*(*xi, xj*) is theoretical covariance of partial variables *xi* and *xj* given joint PDF .

Note,

Each *V*(*xi*) on diagonal of *V*(*X*) is theoretical variance of partial variable *xi* given marginal PDF .

Note,

Given two random variables *X* and *Y* along with a joint PDF *f*(*X*, *Y*), theoretical covariance of *X* and *Y* is defined as follows:

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|  | (1.5) |

If the random variables *X* and *Y* are mutually independent given the joint PDF *f*(*X*, *Y*), its covariance is zero, *V*(*X*, *Y*)=0. However, it is not sure that *X* and *Y* are mutually independent if *V*(*X*, *Y*)=0. Note, joint PDF is the PDF having two or more random variables. If *X* and *Y* are multivariate vectors, *V*(*X*, *Y*) is theoretical covariance matrix of *X* and *Y* given the joint PDF *f*(*X*, *Y*). When *X* = (*x*1, *x*2,…, *xm*)*T* and *Y* = (*y*1, *y*2,…, *yn*)*T* are multivariate, *V*(*X*, *Y*) has following form:

Where *V*(*xi, yj*) is covariance of *xi* and *yj*. We have:

Note,

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As usual, *E*(*X*) and *V*(*X*) are often denoted as *μ* and Σ, respectively if they are parameters of PDF. Note, most of PDFs whose parameters are not *E*(*X*) and *V*(*X*). When *X* is univariate, Σ is often denoted as *σ*2 (if it is parameter of PDF). For example, if *X* is univariate and follows normal distribution, its PDF is:

If *X* = (*x*1, *x*2,…, *xn*)*T* is multivariate and follows multinormal (multivariate normal) distribution, its PDF is:

In this case, parameters *μ* and Σ have following forms:

Of course, *μ* and Σ are determined by equation 1.3 and equation 1.4, respectively with note that Σ is symmetric and invertible in case of multinormal distribution. Each *μi* is theoretical mean of partial variable *xi* as usual.

Each *σij* where *i*≠*j* is theoretical covariance of partial variables *xi* and *xj* as usual.

Note,

Each *σii* on diagonal of Σ is theoretical variance of partial variable *xi* as usual.

Note,

Without loss of generality, by default, random variable *X* in this research is multivariate as vector if there is no additional explanation. Followings are some formulas related to theoretical expectation *E*(*X*) and variance *V*(*X*).

Let *a* and *A* be scalar constant and vector constant, respectively, we have:

Given a set of random variables = {*X*1, *X*2,…, *XN*) and *N* scalar constants *ci* (s), we have:

Where *V*(*Xi*, *Xj*) is covariance of *Xi* and *Xj*.

If all *Xi* (s) are mutually independent, then

Note, given joint PDF *f*(*X*1, *X*2,…, *XN*), two random variables *Xi* and *Xj* are mutually independent if *f*(*Xi*, *Xj*) = *f*(*Xi*)*f*(*Xj*) where *f*(*Xi*, *Xj*), *f*(*Xi*), and *f*(*Xj*) are defined as aforementioned integrals of *f*(*X*1, *X*2,…, *XN*). Therefore, if only one PDF *f*(*X*) is defined then, of course *X*1, *X*2,…, and *XN* are mutually independent and moreover, they are identically distributed.

If all *Xi* (s) are identically distributed, which implies that every *Xi* has the same distribution (the same PDF) with the same parameter, then

Note, if all *Xi* (s) are identically distributed, every *Xi* can be represented by the same random variable *X*.

If all *Xi* (s) are mutually independent and identically distributed (iid), then

Because EM algorithm is essentially an advanced version of maximum likelihood estimation (MLE) method, it is necessary to describe MLE in short. Suppose random variable *X* conforms to a distribution specified by the PDF denoted *f*(*X* | Θ) with parameter Θ. For example, if *X* is vector and follows normal distribution then,

Where *μ* and Σ are theoretical mean vector and covariance matrix, respectively with note that Θ = (*μ*, Σ)*T*. The notation |.| denotes determinant of given matrix and the notation Σ–1 denotes inverse of matrix Σ. Note, Σ is invertible and symmetric. Parameter of normal distribution is theoretical mean and theoretical variance,

But parameters of different distributions can be different from such mean and variance. Anyhow theoretical mean and theoretical variance can be calculated based on parameter Θ.

For example, suppose *X* = (*x*1, *x*2,…, *xn*)*T* follows multinomial distribution of *K* trials, its PDF is:

Where *xj* are integers and Θ = (*p*1, *p*2,…, *pn*)*T* is the set of probabilities such that

Note, *xj* is the number of trials generating nominal value *j*. Obviously, the parameter Θ = (*p*1, *p*2,…, *pn*)*T* does not include theoretical mean *E*(*X*) and theoretical variance *V*(*X*) but *E*(*X*) and *V*(*X*) of multinomial distribution can be calculated based on Θ as follows:

When random variable *X* is considered as an observation, a statistic denoted *τ*(*X*) is function of *X*. For example, *τ*(*X*) = *X*, *τ*(*X*) = *aX* + *A* where *a* is scalar constant and *A* is vector constant, and *τ*(*X*) = *XXT* are statistics of *X*. Statistic *τ*(*X*) can be vector-by-vector functions, for example, *τ*(*X*) = (*X*, *XXT*)*T* is a very popular statistic of *X*.

In practice, if *X* is replaced by sample = {*X*1, *X*2,…, *XN*} including *N* observation *Xi*, a statistic is now function of *Xi* (s), for instance, quantities and *S* defined below are statistics:

For multinormal distribution, and *S* are estimates of theoretical mean *μ* and theoretical covariance matrix Σ. They are called sample mean and sample variance, respectively. Essentially, *X* is special case of when has only one observation as = {*X*}.

Statistic *τ*(*X*) is called sufficient statistic if it has all and only information to estimate parameter Θ. For example, sufficient statistic of normal PDF is *τ*(*X*) = (*X*, *XXT*)*T*. In fact, parameter Θ = (*μ*, Σ)*T* of normal PDF, which includes theoretical mean *μ* and theoretical covariance matrix Σ, is totally determined based on all and only *X* and *XXT* (there is no redundant information in *τ*(*X*)) where *X* is observation considered as random variable, as follows:

Similarly, given *X* = (*x*1, *x*2,…, *xn*)*T*, sufficient statistic of multinomial PDF of *K* trials is *τ*(*X*) = (*x*1, *x*2,…, *xn*)*T* due to:

Given a sample containing observations, purpose of point estimation is to estimate unknown parameter Θ based on such sample. The result of estimation process is the estimate as approximation of unknown Θ. Formula to calculate based on sample is called estimator of Θ. As a convention, estimator of Θ is denoted or where *X* is an observation and is sample including many observations. Actually, or is the same to but the notation or implies that is calculated based on observations. For example, given sample = {*X*1, *X*2,…, *XN*} including *N* observations iid *Xi*, estimator of theoretical mean *μ* of normal distribution is:

As usual, estimator of Θ is determined based on sufficient statistics which in turn are functions of observations where observations are considered as random variables. Estimation methods mentioned in this research are MLE, Maximum A Posteriori (MAP), and EM in which MAP and EM are variants of MLE.

According to viewpoint of Bayesian statistics, the parameter Θ is random variable and it conforms some distribution. In some research, Θ represents a hypothesis. Equation 1.6 specifies Bayes’ rule in which *f*(Θ|*ξ*) is called prior PDF (prior distribution) of Θ whereas *f*(Θ|*X*) is called posterior PDF (posterior distribution) of Θ given observation *X*. Note, *ξ* is parameter of the prior *f*(Θ|*ξ*), which is known as second-level parameter. For instance, if the prior *f*(Θ|*ξ*) is multinormal (multivariate normal) PDF, we have *ξ* = (*μ*0, Σ02)*T* which are theoretical mean and theoretical covariance matrix of random variable Θ. Because *ξ* is constant, the prior PDF *f*(Θ|*ξ*) can be denoted *f*(Θ). The posterior PDF *f*(Θ|*X*) ignores *ξ* because *ξ* is constant in *f*(Θ|*X*).

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| --- | --- |
|  | (1.6) |

In Bayes’ rule, the PDF *f*(*X* | Θ) is called likelihood function. If posterior PDF *f*(Θ|*X*) has the same form of prior PDF *f*(Θ|*ξ*), such posterior PDF and prior PDF are called conjugate PDFs (conjugate distributions, conjugate probabilities) and *f*(Θ|*ξ*) is called conjugate prior (Wikipedia, Conjugate prior, 2018) for likelihood function *f*(*X*|Θ). Such pair of *f*(Θ|*ξ*) and *f*(*X*|Θ) is called conjugate pair. For example, if prior PDF *f*(Θ|*ξ*) is beta distribution and likelihood function *P*(*X*|Θ) follows binomial distribution then, posterior PDF *f*(Θ|*X*) is beta distribution and hence, *f*(Θ|*ξ*) and *f*(Θ|*X*) are conjugate distributions. Shortly, whether posterior PDF and prior PDF are conjugate PDFs depends on prior PDF and likelihood function.

There is a special conjugate pair that both prior PDF *f*(Θ|*ξ*) and likelihood function *f*(*X*|Θ) are multinormal, which results that posterior PDF *f*(Θ|*X*) is multinormal. For instance, when *X* = (*x*1, *x*2,…, *xn*)*T*, the likelihood function *f*(*X*|Θ) is multinormal as follows:

Where Θ = (*μ*, Σ)*T* and *μ* = (*μ*1, *μ*2,…, *μn*)*T*. Suppose only *μ* is random variable which follows multinormal distribution with parameter *ξ* = (*μ*0, Σ0)*T* where *μ*0 = (*μ*01, *μ*02,…, *μ*0*n*)*T*. Note, Σ and Σ0 are symmetric and invertible. The prior PDF *f*(Θ|*ξ*) is:

It is proved that the posterior PDF *f*(Θ|*X*)=*f*(*μ*|*X*) distributes normally with theoretical mean *Mμ* and covariance matrix Σ*μ* as follows (Steorts, 2018, p. 13):

Where (Steorts, 2018, p. 13),

The sign “” indicates proportion ■

When *X* is evaluated as observation, let be estimate of Θ. It is calculated as a maximizer of the posterior PDF *f*(Θ|*X*) given *X*. Here data sample has only one observation *X* as = {*X*}, in other words, *X* is special case of here.

Because the prior PDF *f*(Θ|*ξ*) is assumed to be fixed and the value is constant with regard to Θ, we have:

Obviously, MLE method determines as a maximizer of the likelihood function *f*(*X* | Θ) with regard to Θ when *X* is evaluated as observation. It is interesting that the likelihood function *f*(*X*|Θ) is the PDF of *X* with parameter Θ. For convenience, MLE maximizes the natural logarithm of the likelihood function denoted *l*(Θ) instead of maximizing the likelihood function.

|  |  |
| --- | --- |
|  | (1.7) |

Where *l*(Θ) = log(*f*(*X* | Θ)) is called log-likelihood function of Θ. Recall that equation 1.7 implies the optimization problem. Note, *l*(Θ) is function of Θ if *X* is evaluated as observation.

|  |  |
| --- | --- |
|  | (1.8) |

Equation 1.7 is the simple result of MLE for estimating parameter based on observed sample. The notation *l*(Θ|*X*) implies that *l*(Θ) is determined based on *X*. If the log-likelihood function *l*(Θ) is first-order smooth function then, from equation 1.7, the estimate can be solution of the equation created by setting the first-order derivative of *l*(Θ) regarding Θ to be zero, *Dl*(Θ)=**0***T*. If solving such equation is too complex or impossible, some popular methods to solve optimization problem are Newton-Raphson (Burden & Faires, 2011, pp. 67-71), gradient descent (Ta, 2014), and Lagrange duality (Wikipedia, Karush–Kuhn–Tucker conditions, 2014). Note, solving the equation *Dl*(Θ)=**0***T* may be incorrect in some case, for instance, in theory, such that *Dl*()=**0***T* may be a saddle point (not a maximizer).

For example, suppose *X* = (*x*1, *x*2,…, *xn*)*T* is vector and follows multinormal distribution,

Then the log-likelihood function is

Where *μ* and Σ are mean vector and covariance matrix of *f*(*X* | Θ), respectively with note that Θ = (*μ*, Σ)*T*. The notation |.| denotes determinant of given matrix and the notation Σ–1 denotes inverse of matrix Σ. Note, Σ is invertible and symmetric. Because normal PDF is smooth enough function, from equation 1.7, the estimate is solution of the equation created by setting the first-order of *l*(Θ) regarding *μ* and Σ to be zero. The first-order partial derivative of *l*(Θ) with respect to *μ* is (Nguyen, 2015, p. 35):

Setting this partial derivative to be zero, we obtain:

The first-order partial derivative of *l*(Θ) with respect to Σ is:

Due to:

And

Because Bilmes (Bilmes, 1998, p. 5) mentioned:

Where tr(*A*) is trace operator which takes sum of diagonal elements of square matrix, . This implies (Nguyen, 2015, p. 45):

Where Σ is symmetric and invertible matrix. Substituting the estimate into the first-order partial derivative of *l*(Θ) with respect to Σ, we have:

The estimate is the solution of equation formed by setting the first-order partial derivative of *l*(Θ) regarding Σ to zero matrix. Let (**0**) denote zero matrix.

We have:

Finally, MLE results out the estimate for normal distribution given observation *X* as follows:

When then , which implies that the estimate of covariance matrix is arbitrary with constraint that it is symmetric and invertible. This is reasonable because the sample is too small with only one observation *X*. When *X* is replaced by a sample = {*X*1, *X*2,…, *XN*} in which all *Xi* (s) are mutually independent and identically distributed (iid), it is easy to draw the following result by the similar way with equation 1.11.

Here, and are sample mean and sample variance ■

In practice, if *X* is observed as *N* particular observations *X*1, *X*2,…, *XN*. Let = {*X*1, *X*2,…, *XN*} be the observed sample of size *N* in which all *Xi* (s) are iid. Essentially, *X* is special case of when has only one observation as = {*X*}. The Bayes’ rule specified by equation 1.6 is re-written as follows:

However, the meaning of Bayes’ rule does not change. Because all *Xi* (s) are iid, the likelihood function becomes product of partial likelihood functions as follows:

|  |  |
| --- | --- |
|  | (1.9) |

The log-likelihood function of Θ becomes:

|  |  |
| --- | --- |
|  | (1.10) |

The notation *l*(Θ|) implies that *l*(Θ) is determined based on . We have:

|  |  |
| --- | --- |
|  | (1.11) |

Equation 1.11 is the main result of MLE for estimating parameter based on observed sample. If the log-likelihood function *l*(Θ) is first-order smooth function then, from equation 1.11, the estimate can be solution of the equation created by setting the first-order derivative of *l*(Θ) regarding Θ to be zero. If solving such equation is too complex, some popular methods to solve optimization problem are Newton-Raphson (Burden & Faires, 2011, pp. 67-71), gradient descent (Ta, 2014), and Lagrange duality (Wikipedia, Karush–Kuhn–Tucker conditions, 2014).

For example, suppose each *Xi* = (*xi*1, *xi*2,…, *xin*)*T* is vector and follows multinomial distribution of *K* trials,

Where *xik* are integers and Θ = (*p*1, *p*2,…, *pn*)*T* is the set of probabilities such that

Note, *xik* is the number of trials generating nominal value *k*.

Given sample = {*X*1, *X*2,…, *XN*} in which all *Xi* (s) are iid, according to equation 1.10, the log-likelihood function is

Because there is the constraint , we use Lagrange duality method to maximize *l*(Θ). The Lagrange function *la*(Θ, *λ*) is sum of *l*(Θ) and the constraint as follows:

Note, *λ* is called Lagrange multiplier. Of course, *la*(Θ, *λ*) is function of Θ and *λ*. Because multinomial PDF is smooth enough, the estimate is solution of the equation created by setting the first-order of *la*(Θ) regarding *pj* and *λ* to be zero. The first-order partial derivative of *la*(Θ) with respect to *pj* is:

Setting this partial derivative to be zero, we obtain following equation:

Summing this equation over *n* variables *pj*, we obtain:

Due to

We have

Substitute *λ* = *nN* into equation

We get the estimate as follows:

Quality of estimation is measured by mean and variance of the estimate . The mean of is:

|  |  |
| --- | --- |
|  | (1.12) |

The notation implies the formulation to calculate , which is resulted from MLE, MAP, or EM. Hence, is considered as function of *X* in the integral . The is unbiased estimate if . Otherwise, if then, is biased estimate. As usual, unbiased estimate is better than biased estimate. The condition is the criterion to check if an estimate is unbiased, which is applied for all estimation methods.

The variance of is:

|  |  |
| --- | --- |
|  | (1.13) |

The smaller the variance , the better the is.

For example, given multinormal distribution and given sample = {*X*1, *X*2,…, *XN*} where all *Xi* (s) are iid, the estimate from MLE is:

Due to:

Then is unbiased estimate. We also have:

(Due to )

(Due to )

(Let *X* be random variable representing all iid *Xi* (s))

(Due to and the variance )

It is necessary to calculate the variance . In fact, we have:

Therefore, we have:

Hence, we conclude that is biased estimate because of ■

Without loss of generality, suppose parameter Θ is vector, the second-order derivative of the log-likelihood function *l*(Θ) is called likelihood Hessian matrix (Zivot, 2009, p. 7) denoted *S*(Θ).

|  |  |
| --- | --- |
|  | (1.14) |

Suppose Θ = (*θ*1, *θ*2,…, *θr*)*T* where there are *r* partial parameters *θk*, equation 1.14 is expended as follows:

Where,

The notation *l*(Θ|*X*) implies that *l*(Θ) is determined based on *X*, according to equation 1.8. The notation *S*(Θ|*X*) implies *S*(Θ) is calculated based on *X*. If sample replaces *X* then,

|  |  |
| --- | --- |
|  | (1.15) |

Where = {*X*1, *X*2,…, *XN*} be the observed sample of size *N* in which all *Xi* (s) are iid. The notation *l*(Θ|) implies that *l*(Θ) is determined based on , according to equation 1.11. The notation *S*(Θ|) implies *S*(Θ) is calculated based on .

The negative expectation of likelihood Hessian matrix is called information matrix or Fisher information matrix denoted *I*(Θ). Please distinguish information matrix *I*(Θ) from identity matrix *I*.

|  |  |
| --- | --- |
|  | (1.16) |

If *S*(Θ) is calculated by equation 1.14 with observation *X* then, *I*(Θ) becomes:

|  |  |
| --- | --- |
|  | (1.17) |

The notation *l*(Θ|*X*) implies that *l*(Θ) is determined based on *X*, according to equation 1.8. The notation *I*(Θ|*X*) implies *I*(Θ) is calculated based on *X*. Note, is considered as function of *X* in the integral .

If *S*(Θ) is calculated by equation 1.15 with observation sample = {*X*1, *X*2,…, *XN*} in which all *Xi* (s) are iid then, *I*(Θ) becomes:

|  |  |
| --- | --- |
|  | (1.18) |

Where *X* is random variable representing every *Xi*. The notation *I*(Θ|) implies *I*(Θ) is calculated based on . Note, is considered as function of *X* in the integral . Following is proof of equation 1.18.

(The notation *l*(Θ|) implies that *l*(Θ) is determined based on )

(Due to equation 1.8 and iid *Xi* (s))

(Let *X* be random variable representing every *Xi*)

For MLE method, the inverse of estimator information matrix is called Cramer-Rao lower bound denoted .

|  |  |
| --- | --- |
|  | (1.19) |

Where *I*(Θ) is calculated by equation 1.17 or equation 1.18. Any covariance matrix of a MLE estimate has such Cramer-Rao lower bound. Such Cramer-Rao lower bound becomes if and only if is unbiased, (Zivot, 2009, p. 11):

|  |  |
| --- | --- |
|  | (1.20) |

Note, equation 1.19 and equation 1.20 are only valid for MLE method. The sign “≥” implies lower bound. In other words, Cramer-Rao lower bound is variance of the optimal MLE estimate. Moreover, beside the criterion , equation 1.20 can be used as another criterion to check if an estimate is unbiased. However, the criterion is applied for all estimation methods whereas equation 1.20 is only applied for MLE.

Suppose Θ = (*θ*1, *θ*2,…, *θr*)*T* where there are *r* partial parameter *θk*, so the estimate is . Each element on diagonal of the Cramer-Rao lower bound is lower bound of a variance of , denoted . Let be lower bound of , of course we have:

|  |  |
| --- | --- |
|  | (1.21) |

The sign “≥” implies lower bound. Derived from equation 1.18 and equation 1.19, is specified by equation 1.22.

|  |  |
| --- | --- |
|  | (1.22) |

Where *N* is size of sample = {*X*1, *X*2,…, *XN*} in which all *Xi* (s) are iid. If there is only one observation *X* then, *N* = 1. Of course, is information matrix of . If is univariate, is scalar, which called information value.

For example, let = {*X*1, *X*2,…, *XN*} be the observed sample of size *N* with note that all *Xi* (s) are iid, given multinormal PDF as follows:

Where *n* is dimension of vector *X* and Θ = (*μ*, Σ)*T* with note that *μ* is theoretical mean vector and Σ is theoretical covariance matrix. Note, Σ is invertible and symmetric. From previous example, the MLE estimate given is:

Mean and variance of from previous example are:

We knew that is unbiased estimate with criterion . Now we check again if is unbiased estimate with equation 1.21 as another criterion for MLE. Hence, we firstly calculate the lower bound and then compare it with the variance . In fact, according to equation 1.8, the log-likelihood function is:

The partial first-order derivative of *l*(Θ|*X*) with regard to *μ* is (Nguyen, 2015, p. 35):

The partial second-order derivative of *l*(Θ|*X*) with regard to *μ* is (Nguyen, 2015, p. 36):

(Due to Σ is symmetric)

According to equation 1.22, the lower bound is:

Due to , is unbiased estimate according to the criterion specified by equation 1.21.

Mean of from previous example is:

We knew that is biased estimate because . Now we check again if is biased estimate with equation 1.21 as another criterion for MLE. The partial first-order derivative of *l*(Θ|*X*) with regard to Σ is:

Due to:

And

Because Bilmes (Bilmes, 1998, p. 5) mentioned:

Where tr(*A*) is trace operator which takes sum of diagonal elements of matrix .

This implies (Nguyen, 2015, p. 45):

According to equation 1.22, the lower bound is:

(Due to *l*(Θ|*X*) is smooth enough)

(Because Σ–1 and are symmetric matrices)

Where (**0**) is zero matrix. This implies the lower bound is inexistent. Hence, is biased estimate. Even there is no unbiased estimate of variance for normal distribution by MLE ■

MLE ignores prior PDF *f*(Θ|*ξ*) because *f*(Θ|*ξ*) is assumed to be fixed but Maximum A Posteriori (MAP) method (Wikipedia, Maximum a posteriori estimation, 2017) concerns *f*(Θ|*ξ*) in maximization task when is constant with regard to Θ.

Let *f*(*X*, Θ | *ξ*) be the joint PDF of *X* and Θ where Θ is also random variable too. Note, *ξ* is parameter in the prior PDF *f*(Θ|*ξ*). The likelihood function in MAP is also *f*(*X*, Θ | *ξ*).

|  |  |
| --- | --- |
|  | (1.23) |

Theoretical mean and variance of *X* are based on the joint PDF *f*(*X*, Θ | *ξ*) as follows:

|  |  |
| --- | --- |
|  | (1.24) |

|  |  |
| --- | --- |
|  | (1.25) |

Theoretical mean and variance of Θ are based on *f*(Θ|*ξ*) because *f*(Θ|*ξ*) is function of only Θ when *ξ* is constant.

|  |  |
| --- | --- |
|  | (1.26) |

|  |  |
| --- | --- |
|  | (1.27) |

In general, statistics of Θ are still based on *f*(Θ|*ξ*). Given sample = {*X*1, *X*2,…, *XN*} in which all *Xi* (s) are iid, the likelihood function becomes:

|  |  |
| --- | --- |
|  | (1.28) |

The log-likelihood function in MAP is re-defined with observation *X* or sample as follows:

|  |  |
| --- | --- |
|  | (1.29) |

|  |  |
| --- | --- |
|  | (1.30) |

Where *l*(Θ) is specified by equation 1.8 with observation *X* or equation 1.10 with sample . Therefore, the estimate is determined according to MAP as follows:

|  |  |
| --- | --- |
|  | (1.31) |

Good information provided by the prior *f*(Θ|*ξ*) can improve quality of estimation. Essentially, MAP is an improved variant of MLE. Later on, we also recognize that EM algorithm is also a variant of MLE. All of them aim to maximize log-likelihood functions. Likelihood Hessian matrix , information matrix , and Cramer-Rao lower bound , are extended in MAP with the new likelihood function .

|  |
| --- |
|  |

Where *N* is size of sample = {*X*1, *X*2,…, *XN*} in which all *Xi* (s) are iid. If there is only one observation *X* then, *N* = 1.

Mean and variance of the estimate which are used to measure estimation quality are not changed except that the joint PDF *f*(*X*, Θ | *ξ*) is used instead.

|  |  |
| --- | --- |
|  | (1.32) |

|  |  |
| --- | --- |
|  | (1.33) |

The notation implies the formulation to calculate , which is considered as function of *X* and Θ in the integral . Recall the is unbiased estimate if . Otherwise, if then, is biased estimate. Moreover, the smaller the variance , the better the is. Recall that there are two criteria to check if is unbiased estimate. Concretely, is unbiased estimate if one of two following conditions is satisfied:

The criterion is expended for MAP.

It is necessary to have an example for parameter estimation with MAP. Given sample = {*X*1, *X*2,…, *XN*} in which all *Xi* (s) are iid. Each *n*-dimension *Xi* has following multinormal PDF:

Where *μ* and Σ are mean vector and covariance matrix of *f*(*X* | Θ), respectively with note that Θ = (*μ*, Σ)*T*. The notation |.| denotes determinant of given matrix and the notation Σ–1 denotes inverse of matrix Σ. Note, Σ is invertible and symmetric.

In Θ = (*μ*, Σ)*T*, suppose only *μ* distributes normally with parameter *ξ* = (*μ*0, Σ0) where *μ*0 and Σ0 are theoretical mean and covariance matrix of *μ*. Thus, Σ is variable but not random variable. The second-level parameter *ξ* is constant. The prior PDF *f*(Θ|*ξ*) becomes *f*(*μ*|*ξ*), which specified as follows:

Note, *μ*0 is *n*-element vector like *μ* and Σ0 is *n*x*n* matrix like Σ. Of course, Σ0 is also invertible and symmetric. Suppose *μ* = (*μ*1, *μ*2,…, *μn*)*T*, *μ*0 = (*μ*01, *μ*02,…, *μ*0*n*)*T*, and

It is deduced that *μ*0*j* is theoretical mean of *μj* whereas *δij* (*i*≠*j*) is covariance of *μi* and *μj*. Especially, *δii* is variance of *μi*.

Theoretical mean of *X* is:

Theoretical variance of *X* is:

The log-likelihood function in MAP is

Because normal PDF is smooth enough, from equation 1.24, the estimate is solution of the equation created by setting the first-order of (Θ) regarding *μ* and Σ to be zero. Due to (Nguyen, 2015, p. 35):

And (Nguyen, 2015, p. 35)

The first-order partial derivative of (Θ) with respect to *μ* is:

Setting this partial derivative to be zero, we obtain:

Where *I* is identity matrix. Let,

We obtain the following equation to estimate *μ* and Σ:

The first-order partial derivative of *l*(Θ) with respect to Σ is:

Due to:

And

Because Bilmes (Bilmes, 1998, p. 5) mentioned:

Where tr(*A*) is trace operator which takes sum of diagonal elements of square matrix, . This implies (Nguyen, 2015, p. 45):

Where Σ is symmetric and invertible matrix. The estimate is the solution of equation formed by setting the first-order partial derivative of *l*(Θ) regarding Σ to zero matrix. Let (**0**) denote zero matrix.

We have:

MAP results out a system of two equations whose solution is the estimate as follows:

Where *I* is identity matrix and

Because Σ is independent from the prior PDF *f*(*μ* | *μ*0, Σ0), it is estimated by MLE as usual,

The estimate in MAP here is as same as the one in MLE and so, it is biased. Substituting for Σ, we obtain the estimate in MAP:

Note,

Now we check if is unbiased estimate. In fact, we have:

(Due to *E*(*X*) = *μ*0)

Therefore, the estimate is biased because the variable *μ* is not always to equal *μ*0.

Now we try to check again if is unbiased estimate with Cramer-Rao lower bound. The second-order partial derivative of regarding *μ* is:

(Because Σ and Σ0 are symmetric)

Cramer-Rao lower bound of is:

Variance of is:

Because it is difficult to calculate , suppose we fix Σ so that , we have:

(Due to )

The Cramer-Rao lower bound of is re-written as follows:

Obviously, is biased estimate due to . In general, the estimate in MAP is affected by the prior PDF *f*(Θ|*ξ*). Even though it is biased, it can be better than the one resulted from MLE because of valuable information in *f*(Θ|*ξ*). For instance, if fixing Σ, the variance of from MAP is “smaller” (lower bounded) than the one from MLE ■

Now we skim through an introduction of EM algorithm. Suppose there are two spaces ***X*** and ***Y***, in which ***X*** is *hidden space* whereas ***Y*** is *observed space*. We do not know ***X*** but there is a mapping from ***X*** to ***Y*** so that we can survey ***X***by observing ***Y***. The mapping is many-one function *φ*: ***X*** → ***Y*** and we denote *φ*–1(*Y*) = {: *φ*(*X*) = *Y*} as all such that *φ*(*X*) = *Y*. We also denote ***X***(*Y*) = *φ*–1(*Y*). Let *f*(*X* | Θ) be the PDF of random variable and let *g*(*Y* | Θ) be the PDF of random variable . Note, *Y* is also called observation. Equation 1.34 specifies *g*(*Y* | Θ) as integral of *f*(*X* | Θ) over *φ*–1(*Y*).

|  |  |
| --- | --- |
|  | (1.34) |

Where Θ is probabilistic parameter represented as a column vector, Θ= (*θ*1, *θ*2,…, *θr*)*T* in which each *θi* is a particular parameter. If *X* and *Y* are discrete, equation 1.34 is re-written as follows:

According to viewpoint of Bayesian statistics, Θ is also random variable. As a convention, let Ω be the domain of Θ such that and the dimension of Ω is *r*. For example, normal distribution has two particular parameters such as mean *μ* and variance *σ*2 and so we have Θ= (*μ*, *σ*2)*T*. Note that, Θ can degrades into a scalar as Θ = *θ*. The conditional PDF of *X* given *Y*, denoted *k*(*X* | *Y*, Θ), is specified by equation 1.35.

|  |  |
| --- | --- |
|  | (1.35) |

According to DLR (Dempster, Laird, & Rubin, 1977, p. 1), ***X*** is called *complete data* and the term “incomplete data” implies existence of ***X*** and ***Y*** where ***X*** is not observed directly and ***X*** is only known by the many-one mapping *φ*: ***X*** → ***Y***. In general, we only know ***Y***, *f*(*X* | Θ), and *k*(*X* | *Y*, Θ) and so our purpose is to estimate Θ based on such ***Y***, *f*(*X* | Θ), and *k*(*X* | *Y*, Θ). Like MLE approach, EM algorithm also maximizes the likelihood function to estimate Θ but the likelihood function in EM concerns ***Y*** and there are also some different aspects in EM which will be described later. Pioneers in EM algorithm firstly assumed that *f*(*X* | Θ) belongs to exponential family with note that many popular distributions such as normal, multinomial, and Poisson belong to exponential family (please see table 1.1). Although DLR (Dempster, Laird, & Rubin, 1977) proposed a generality of EM algorithm in which *f*(*X* | Θ) distributes arbitrarily, we should concern exponential family a little bit. Exponential family (Wikipedia, Exponential family, 2016) refers to a set of probabilistic distributions whose PDF (s) have the same exponential form according to equation 1.36 (Dempster, Laird, & Rubin, 1977, p. 3):

|  |  |
| --- | --- |
|  | (1.36) |

Where *b*(*X*) is a function of *X*, which is called base measure and *τ*(*X*) is a vector function of *X*, which is sufficient statistic. For example, the sufficient statistic of normal distribution is *τ*(*X*) = (*X*, *XXT*)*T*. Equation 1.36 expresses the canonical form of exponential family. Recall that Ω is the domain of Θ such that . Suppose that Ω is a convex set. If Θ is restricted only to Ω then, *f*(*X* | Θ) specifies a *regular exponential family*. If Θ lies in a curved sub-manifold Ω0 of Ω then, *f*(*X* | Θ) specifies a *curved exponential family*. The *a*(Θ) is *partition function* for variable *X*, which is used for normalization.

As usual, a PDF is known as a popular form but its exponential family form (canonical form of exponential family) specified by equation 1.36 looks unlike popular form although they are the same. Therefore, parameter in popular form is different from parameter in exponential family form.

For example, multinormal distribution with theoretical mean *μ* and covariance matrix Σ of random variable *X* = (*x*1, *x*2,…, *xn*)*T* has PDF in popular form is:

Hence, parameter in popular form is Θ = (*μ*, Σ)*T*. Exponential family form of such PDF is:

Where,

Hence, parameter in exponential family form is Θ = (*θ*1, *θ*2)*T*. Although, *f*(*X* | *θ*1, *θ*2) looks unlike *f*(*X* | *μ*, Σ) but they are the same, *f*(*X* | *θ*1, *θ*2) = *f*(*X* | *μ*, Σ). In fact, we have:

We also have:

(Due to |Σ–1| = |Σ|–1)

Therefore,

(Because Σ is symmetric, *μT*Σ–1*X* = *XT*Σ–1*μ*)

The exponential family form is used to represents all distributions belonging to exponential family as canonical form. Parameter in exponential family form is called exponential family parameter. As a convention, parameter Θ mentioned in EM algorithm is often exponential family parameter if PDF belongs to exponential family and there is no additional information.

Table 1.1 shows some popular distributions belonging to exponential family along with their canonical forms (Wikipedia, Exponential family, 2016). In case of multivariate distributions, dimension of random variable *X* = (*x*1, *x*2,…, *xn*)*T* is *n*.

|  |  |  |  |  |  |
| --- | --- | --- | --- | --- | --- |
| Distribution | Popular  PDF | Exponential  family  parameter  Θ | *τ*(*X*) | *b*(*X*) | *a*(Θ) |
| Multinormal |  |  |  |  |  |
| Multinomial | Where, , , and . |  |  |  | 1 |

**Table 1.1.** Some popular distributions belonging to exponential family

It is necessary to survey some features of exponential family. The first-order derivative of log(*a*(Θ)) is expectation of transposed *τ*(*X*).

The second-order derivative of log(*a*(Θ)) is (Jebara, 2015):

Where,

Hence (Hardle & Simar, 2013, pp. 125-126),

Where *V*(*τ*(*X*) | Θ) is central covariance matrix of *τ*(*X*). Please read the book “Matrix Analysis and Calculus” by Nguyen (Nguyen, 2015) for comprehending derivative of vector and matrix. Let *a*(Θ | *Y*) be a so-called *observed partition function* for observation *Y*.

Similarly, we obtain that the first-order derivative of log(*a*(***θ*** | *Y*)) is expectation of transposed *τ*(*X*) based on *Y*.

If *f*(*X* | Θ) follows exponential family, the conditional density *k*(*X* | *Y*, Θ) is determined as follows:

Indeed, *k*(*X* | *Y*, Θ) is conditional PDF. If *f*(*X* | Θ) follows exponential family then, *k*(*X* | *Y*, Θ) also follows exponential family. In fact, we have:

Note that *k*(*X* | *Y*, Θ) is determined on . Of course, we have:

The first-order derivative of log(*a*(Θ | *Y*)) is:

The second-order derivative of log(*a*(Θ) | *Y*) is:

Where *V*(*τ*(*X*) | *Y*, Θ) is central covariance matrix of *τ*(*X*) given observed *Y*. Table 1.2 is summary of *f*(*X* | Θ), *g*(*Y* | Θ), *k*(*X* | *Y*, Θ), *a*(Θ), log’(*a*(Θ)), log’’(*a*(Θ)), *a*(Θ | *Y*), log’(*a*(Θ | *Y*)), and log’’(*a*(Θ | *Y*)) with exponential family.

|  |
| --- |
|  |

**Table 1.2.** Summary of *f*(*X* | Θ), *g*(*Y* | Θ), *k*(*X* | *Y*, Θ), *a*(Θ), log’(*a*(Θ)), *a*(Θ | *Y*), and log’(*a*(Θ | *Y*)) with exponential family

Simply, EM algorithm is iterative process including many iterations, in which each iteration has expectation step (E-step) and maximization step (M-step). E-step aims to estimate sufficient statistic given current parameter and observed data *Y* whereas M-step aims to re-estimate the parameter based on such sufficient statistic by maximizing likelihood function of *X* related to *Y*. EM algorithm is described in the next section in detail. As an introduction, DLR gave an example for illustrating EM algorithm (Dempster, Laird, & Rubin, 1977, pp. 2-3).

**Example 1.1.** Rao (Rao, 1955) presents observed data *Y* of 197 animals following multinomial distribution with four categories, such as *Y* = (*y*1, *y*2, *y*3, *y*4) = (125, 18, 20, 34). The PDF of *Y* is:

Note, probabilities *py*1, *py*2, *py*3, and *py*4 in *g*(*Y* | *θ*) are 1/2 + *θ*/4, 1/4 – *θ*/4, 1/4 – *θ*/4, and *θ*/4, respectively as parameters. The expectation of any sufficient statistic *yi* with regard to *g*(*Y* | *θ*) is:

Observed data *Y* is associated with hidden data *X* following multinomial distribution with five categories, such as *X* = {*x*1, *x*2, *x*3, *x*4, *x*5} where *y*1 = *x*1 + *x*2, *y*2 = *x*3, *y*3 = *x*4, *y*4 = *x*5. The PDF of *X* is:

Note, probabilities *px*1, *px*2, *px*3, *px*4, and *px*5 in *f*(*X* | *θ*) are 1/2, *θ*/4, 1/4 – *θ*/4, 1/4 – *θ*/4, and *θ*/4, respectively as parameters. The expectation of any sufficient statistic *xi* with regard to *f*(*X* | *θ*) is:

Due to *y*1 = *x*1 + *x*2, *y*2 = *x*3, *y*3 = *x*4, *y*4 = *x*5, the mapping function *φ* between *X* and *Y* is *y*1 = *φ*(*x*1, *x*2) = *x*1 + *x*2. Therefore *g*(*Y* | *θ*) is sum of *f*(*X* | *θ*) over *x*1 and *x*2 such that *x*1 + *x*2 = *y*1 according to equation 1.34. In other words, *g*(*Y* | *θ*) is resulted from summing *f*(*X* | *θ*) over all (*x*1, *x*2) pairs such as (0, 125), (1, 124),…, (125, 0) and then substituting (18, 20, 34) for (*x*3, *x*4, *x*5) because of *y*1 = 125 from observed *Y*.

Rao (Rao, 1955) applied EM algorithm into determining the optimal estimate *θ\**. Note *y*2 = *x*3, *y*3 = *x*4, *y*4 = *x*5 are known and so only sufficient statistics *x*1 and *x*2 are not known. Given the *t*th iteration, sufficient statistics *x*1 and *x*2 are estimated as *x*1(*t*) and *x*2(*t*) based on current parameter *θ*(*t*) and *g*(*Y* | *θ*) in E-step below:

Given *py*1 = 1/2 + *θ*/4, which implies that:

When *y*1 = 125, we have:

This suggests us to select:

According to M-step, the next estimate *θ*(*t*+1) is a maximizer of the log-likelihood function of *X* related to *Y*. This log-likelihood function is:

The first-order derivative of log(*f*(*X* | *θ*) is:

Because *y*2 = *x*3 = 18, *y*3 = *x*4 = 20, *y*4 = *x*5 = 34 and *x*2 is approximated by *x*2(*t*), we have:

As a maximizer of log(*f*(*X* | *θ*), the next estimate *θ*(*t*+1) is solution of the following equation

So we have:

Where,

For example, given the initial *θ*(1) = 0.5, at the first iteration, we have:

After five iterations we gets the optimal estimate *θ\**:

Table 1.3 (Dempster, Laird, & Rubin, 1977, p. 3) lists estimates of *θ* over five iterations (*t* =1, 2, 3, 4, 5) with note that *θ*(1) is initialized arbitrarily and *θ*\* = *θ*(5) = *θ*(6) is determined at the 5th iteration. The third column gives deviation *θ\** and *θ*(*t*) whereas the fourth column gives the ratio of successive deviations. Later on, we will know that such ratio implies convergence rate.

|  |  |  |  |
| --- | --- | --- | --- |
| *t* | *θ*(*t*) | *θ\** – *θ*(*t*) | (*θ\** – *θ*(*t*+1)) /  (*θ\** – *θ*(*t*)) |
| 1 | *θ*(1) = 0.5 | 0.1268 | 0.1465 |
| *θ*(2) = 0.6082 | 0.0186 | 0.1346 |
| 2 | *θ*(2) = 0.6082 | 0.0186 | 0.1346 |
| *θ*(3) = 0.6243 | 0.0025 | 0.1330 |
| 3 | *θ*(3) = 0.6243 | 0.0025 | 0.1330 |
| *θ*(4) = 0.6265 | 0.0003 | 0.1328 |
| 4 | *θ*(4) = 0.6265 | 0.0003 | 0.1328 |
| *θ*(5) = 0.6268 | 0 | 0.1328 |
| 5 | *θ*(5) = 0.6268 | 0 | 0.1328 |
| *θ*(6) = 0.6268 | 0 | 0.1328 |

**Table 1.3.** EM algorithm in simple case

For example, at the first iteration, we have:

# **2. EM algorithm**

Expectation maximization (EM) algorithm has many iterations and each iteration has two steps in which expectation step (E-step) calculates sufficient statistic of hidden data based on observed data and current parameter whereas maximization step (M-step) re-estimates parameter. When DLR proposed EM algorithm (Dempster, Laird, & Rubin, 1977), they firstly concerned that the PDF *f*(*X* | Θ) of hidden space belongs to exponential family. E-step and M-step at the *t*th iteration are described in table 2.1 (Dempster, Laird, & Rubin, 1977, p. 4), in which the current estimate is Θ(*t*), with note that *f*(*X* | Θ) belongs to regular exponential family.

|  |
| --- |
| *E-step*:  We calculate current value *τ*(*t*) of the sufficient statistic *τ*(*X*) from observed *Y* and current parameter Θ(*t*) according to equation 2.6:  *M-step*:  Basing on *τ*(*t*), we determine the next parameter Θ(*t*+1) as solution of equation 2.3:  Note, Θ(*t*+1) will become current parameter at the next iteration ((*t*+1)th iteration). |

**Table 2.1.** E-step and M-step of EM algorithm given regular exponential PDF *f*(*X*|Θ)

EM algorithm stops if two successive estimates are equal, Θ*\** = Θ(*t*) = Θ(*t*+1), at some *t*th iteration. At that time we conclude that Θ*\** is the optimal estimate of EM process. Please see table 1.2 to know how to calculate *E*(*τ*(*X*) | Θ(*t*)) and *E*(*τ*(*X*) | *Y*, Θ(*t*)). As a convention, the estimate of parameter Θ resulted from EM process is denoted Θ\* instead of in order to emphasize that Θ\* is solution of optimization problem.

It is necessary to explain E-step and M-step as well as convergence of EM algorithm. Essentially, the two steps aim to maximize log-likelihood function of Θ, denoted *L*(Θ), with respect to observation *Y*.

Where,

Note that log(.) denotes logarithm function. Therefore, EM algorithm is an extension of maximum likelihood estimation (MLE) method. In fact, let *l*(Θ) be log-likelihood function of Θ with respect to *X*.

|  |  |
| --- | --- |
|  | (2.1) |

By referring to table 1.2, the first-order derivative of *l*(Θ) is:

|  |  |
| --- | --- |
|  | (2.2) |

We set the first-order derivative of *l*(Θ) to be zero with expectation that *l*(Θ) will be maximized. Therefore, the optimal estimate Θ*\** is solution of the following equation which is specified in M-step.

The expression *E*(*τ*(*X*) | Θ) is function of Θ but *τ*(*X*) is still dependent on *X*. Let *τ*(*t*) be value of *τ*(*X*) at the *t*th iteration of EM process, candidate for the best estimate of Θ is solution of equation 2.3 according to M-step.

|  |  |
| --- | --- |
|  | (2.3) |

Where,

Thus, we will calculate *τ*(*t*) by maximizing the log-likelihood function *L*(Θ) given *Y*. Recall that maximizing *L*(Θ) is the ultimate purpose of EM algorithm.

Where,

|  |  |
| --- | --- |
|  | (2.4) |

Due to:

It implies:

Because *f*(*X* | Θ) belongs to exponential family, we have:

The log-likelihood function *L*(Θ) is reduced as follows:

By referring to table 1.2, the first-order derivative of *L*(Θ) is:

|  |  |
| --- | --- |
|  | (2.5) |

We set the first-order derivative of *L*(Θ) to be zero with expectation that *L*(Θ) will be maximized, as follows:

It implies:

Let Θ(*t*) be the current estimate at some *t*th iteration of EM process. Derived from the equality above, the value *τ*(*t*) is calculated as seen in equation 2.6.

|  |  |
| --- | --- |
|  | (2.6) |

Where,

Equation 2.6 specifies the E-step of EM process. After *t* iterations we will obtain Θ*\** = Θ(*t*+1) = Θ(*t*) such that *E*(*τ*(*X*) | *Y*, Θ(*t*)) = *E*(*τ*(*X*) | *Y*, Θ*\**) = *τ*(*t*) = *E*(*τ*(*X*) | Θ*\**) = *E*(*τ*(*X*) | Θ(*t*+1)) when Θ(*t*+1) is solution of equation 2.3 (Dempster, Laird, & Rubin, 1977, p. 5). This means that Θ*\** is the optimal estimate of EM process because Θ*\** is solution of the equation:

Thus, we conclude that Θ*\** is the optimal estimate of EM process.

The EM algorithm shown in table 2.1 is totally exact with assumption that *f*(*X*|Θ) belongs to regular exponential family. If *f*(*X*|Θ) is not regular, the maximal point (maximizer) of the log-likelihood function *l*(Θ) is not always the stationary point Θ\* so that the first-order derivative of *l*(Θ) is zero, *l*’(Θ\*) = 0. However, if *f*(*X*|Θ) belongs to curved exponential family, the M-step of the EM algorithm shown in table 2.1 is modified as follows (Dempster, Laird, & Rubin, 1977, p. 5):

|  |  |
| --- | --- |
|  | (2.7) |

Where *τ*(*t*) is calculated by equation 2.6 in E-step. This means that, in more general manner, the maximizer Θ(*t*+1) will be found by some way. Recall that if Θ lies in a curved sub-manifold Ω0 of Ω where Ω is the domain of Θ then, *f*(*X* | Θ) belongs to curved exponential family.

In general, given exponential family, within simple EM algorithm, E-step aims to calculate the current sufficient statistic *τ*(*t*) that the log-likelihood function *L*(Θ(*t*)) gets maximal with such *τ*(*t*) at current Θ(*t*) given *Y* whereas M-step aims to maximize the log-likelihood function *l*(Θ) given *τ*(*t*), as seem in table 2.2. Note, in table 2.2, *f*(*X*|Θ) belongs to curved exponential family but it is not necessary to be regular.

|  |
| --- |
| *E-step*:  Given observed *Y* and current Θ(*t*), current value *τ*(*t*) of the sufficient statistic *τ*(*X*) is the value that the log-likelihood function *L*(Θ(*t*)) gets maximal with such *τ*(*t*). Concretely, suppose Θ\* be a maximizer of *L*(Θ) given *Y* where *L*(Θ) is specified by equation 2.4.  Suppose Θ\* is formulated as function of *τ*(*X*), for instance, Θ\* = *h*(*τ*(*X*)) with note that Θ\* is not evaluated because *τ*(*X*) is not evaluated. Thus, the equation Θ\* = *h*(*τ*(*X*)) is only symbolic formula. Let *τ*(*t*) be a value of *τ*(*X*) such that Θ(*t*) = *h*(*τ*(*X*)). This means with note that Θ\* is replaced by Θ(*t*). If *h*(*τ*(*X*)) is invertible, *τ*(*t*) = *h*–1(Θ(*t*)).  If the PDF *f*(*X*|Θ) belongs to regular exponential family, *τ*(*t*) is calculated more easily according to equation 2.6, given *Y* and Θ(*t*).  Where,  *M-step*:  Basing on *τ*(*t*), we determine the next parameter Θ(*t*+1) by maximizing the log-likelihood function *l*(Θ) given *τ*(*t*), where *l*(Θ) is specified by equation 2.1. Actually, the sufficient statistic *τ*(*t*) calculated in E-step is substituted for unobserved *τ*(*X*) in *l*(Θ) so that it is possible to maximize *l*(Θ) with subject to Θ.  If the PDF *f*(*X*|Θ) belongs to regular exponential family, Θ(*t*+1) is solution of equation 2.3 given *τ*(*t*).  Where,  If the PDF *f*(*X*|Θ) belongs to curved exponential family, Θ(*t*+1) is determined by equation 2.7 given *τ*(*t*). |

**Table 2.2.** E-step and M-step of EM algorithm given exponential PDF *f*(*X*|Θ)

EM algorithm stops if two successive estimates are equal, Θ*\** = Θ(*t*) = Θ(*t*+1), at some *t*th iteration. At that time, Θ*\** is the optimal estimate of EM process, which is an optimizer of *L*(Θ).

Going back example 1.1, given the *t*th iteration, sufficient statistics *x*1 and *x*2 are estimated as *x*1(*t*) and *x*2(*t*) based on current parameter *θ*(*t*) in E-step according to equation 2.6.

Given *py*1 = 1/2 + *θ*/4, which implies that:

Because the probability of *y*1 is 1/2 + *θ*/4 and *y*1 is sum of *x*1 and *x*2, let be conditional probability of *x*1 given *y*1 and let be conditional probability of *x*2 given *y*1 such that

Where *P*(*x*1, *y*1) and *P*(*x*2, *y*1) are joint probabilities of (*x*1, *y*1) and (*x*2, *y*1), respectively. We can select *P*(*x*1, *y*1) = 1/2 and *P*(*x*2, *y*1) = *θ*/4, which implies:

Such that

Note, we can select alternately as *P*(*x*1, *y*1) = *P*(*x*2, *y*1) = (1/2 + *θ*/4) / 2, for example but fixing *P*(*x*1, *y*1) as 1/2 is better because the next estimate *θ*(*t*+1) known later depends only on *x*2(*t*).

When *y*1 is evaluated as *y*1 = 125, we obtain:

The expectation *y*1(*t*) = *E*(*y*1 | *Y*, *θ*(*t*)) gets value 125 when *y*1 is evaluated as *y*1 = 125 and the probability corresponding to *y*1 gets maximal as 1/2 + *θ*(*t*)/4 = 1.

Essentially, equation 2.3 specifying M-step is result of maximizing the log-likelihood function *l*(Θ). This log-likelihood function is:

The first-order derivative of log(*f*(*X* | *θ*) is:

Because *y*2 = *x*3 = 18, *y*3 = *x*4 = 20, *y*4 = *x*5 = 34 and *x*2 is approximated by *x*2(*t*), we have:

As a maximizer of log(*f*(*X* | *θ*), the next estimate *θ*(*t*+1) is solution of the following equation

So we have:

Where,

For example, given the initial *θ*(1) = 0.5, at the first iteration, we have:

After five iterations we gets the optimal estimate *θ\**:

Table 1.3 (Dempster, Laird, & Rubin, 1977, p. 3) show resulted estimation ■

For further research, DLR gave a preeminent generality of EM algorithm (Dempster, Laird, & Rubin, 1977, pp. 6-11) in which *f*(*X* | Θ) specifies arbitrary distribution. In other words, there is no requirement of exponential family. They define the conditional expectation *Q*(Θ’ | Θ) according to equation 2.8 (Dempster, Laird, & Rubin, 1977, p. 6).

|  |  |
| --- | --- |
|  | (2.8) |

If *X* and *Y* are discrete, equation 2.8 can be re-written as follows:

The two steps of generalized EM (*GEM*) algorithm aim to maximize *Q*(Θ | Θ(*t*)) at some *t*th iteration as seen in table 2.3 (Dempster, Laird, & Rubin, 1977, p. 6).

|  |
| --- |
| *E-step*:  The expectation *Q*(Θ | Θ(*t*)) is determined based on current parameter Θ(*t*), according to equation 2.8. Actually, *Q*(Θ | Θ(*t*)) is formulated as function of Θ.  *M-step*:  The next parameter Θ(*t*+1) is a maximizer of *Q*(Θ | Θ(*t*)) with subject to Θ. Note that Θ(*t*+1) will become current parameter at the next iteration (the (*t*+1)th iteration). |

**Table 2.3.** E-step and M-step of GEM algorithm

DLR proved that GEM algorithm converges at some *t*th iteration. At that time, Θ*\** = Θ(*t*+1) = Θ(*t*) is the optimal estimate of EM process, which is an optimizer of *L*(Θ).

It is deduced from E-step and M-step that *Q*(Θ | Θ(*t*)) is increased after every iteration. How to maximize *Q*(Θ|Θ(*t*)) is the optimization problem which is dependent on applications. For example, the estimate Θ(*t*+1) can be solution of the equation created by setting the first-order derivative of *Q*(Θ|Θ(*t*)) regarding Θ to be zero, *DQ*(Θ|Θ(*t*)) = **0***T*. If solving such equation is too complex or impossible, some popular methods to solve optimization problem are Newton-Raphson (Burden & Faires, 2011, pp. 67-71), gradient descent (Ta, 2014), and Lagrange duality (Wikipedia, Karush–Kuhn–Tucker conditions, 2014). Note, solving the equation *DQ*(Θ|Θ(*t*)) = **0***T* may be incorrect in some case, for instance, in theory, Θ(*t*+1) such that *DQ*(Θ(*t*+1)|Θ(*t*)) = **0***T* may be a saddle point (not a maximizer).

GEM algorithm still aims to maximize the log-likelihood function *L*(Θ) specified by equation 2.4, which is explained here. Following is proof of equation 2.8. Suppose the current parameter is Θ after some iteration. Next we must find out the new estimate Θ\* that maximizes the next log-likelihood function *L*(Θ’).

The next log-likelihood function *L*(Θ’) is re-written as follows:

Due to

By applying Jensen’s inequality (Sean, 2009, pp. 3-4) with concavity of logarithm function

into *L*(Θ’), we have (Sean, 2009, p. 6):

Where,

The lower-bound of *L*(Θ’) is defined as follows:

*lb*(Θ’ | Θ) = *Q*(Θ’ | Θ) – *H*(Θ | Θ)

Of course, we have:

*L*(Θ’) ≥ *lb*(Θ’ | Θ)

Suppose at some *t*th iteration, when the current parameter is Θ(*t*), the lower-bound of *L*(Θ) is re-written:

*lb*(Θ | Θ(*t*)) = *Q*(Θ | Θ(*t*)) – *H*(Θ(*t*) | Θ(*t*))

Of course, we have:

*L*(Θ) ≥ *lb*(Θ | Θ(*t*))

The lower bound *lb*(Θ | Θ(*t*)) has following property (Sean, 2009, p. 7):

*lb*(Θ(*t*) | Θ(*t*)) = *Q*(Θ(*t*) | Θ(*t*)) – *H*(Θ(*t*) | Θ(*t*)) = *L*(Θ(*t*))

Indeed, we have:

Recall that the main purpose of GEM algorithm is to maximize the log-likelihood *L*(Θ) = log(*g*(*Y*|Θ)) with observed data *Y*. However, it is too difficult to maximize log(*g*(*Y* | Θ)) because *g*(*Y* | Θ) is not well-defined when *g*(*Y* | Θ) is integral of *f*(*X* | Θ) given a general mapping function. DLR solved this problem by an iterative process which is an instance of GEM algorithm. The lower-bound (Sean, 2009, pp. 7-8) of *L*(Θ) is maximized over many iterations of the iterative process so that *L*(Θ) is maximized finally. Such lower-bound is determined indirectly by the condition expectation *Q*(Θ | Θ(*t*)) so that maximizing *Q*(Θ | Θ(*t*))is the same to maximizing the lower bound. Suppose Θ(*t*+1) is a maximizer of *Q*(Θ | Θ(*t*)) at *t*th iteration, which is also a maximizer of the lower bound at *t*th iteration.

Note, *H*(Θ(*t*) | Θ(*t*)) is constant with regard to Θ. The lower bound is increased after every iteration. As a result, the maximizer Θ*\** of the final lower-bound after many iterations will be expected as a maximizer of *L*(Θ) in final.

Therefore, the two steps of GEM is interpreted with regard to the lower bound *lb*(Θ | Θ(*t*)) as seen in table 2.4.

|  |
| --- |
| *E-step*:  The lower bound *lb*(Θ | Θ(*t*)) is re-calculated based on *Q*(Θ | Θ(*t*)).  *M-step*:  The next parameter Θ(*t*+1) is a maximizer of *Q*(Θ | Θ(*t*)) which is also a maximizer of *lb*(Θ | Θ(*t*)) because *H*(Θ(*t*) | Θ(*t*)) is constant.  Note that Θ(*t*+1) will become current parameter at the next iteration so that the lower bound is increased in the next iteration. |

**Table 2.4.** An interpretation of GEM with lower bound

Because *Q*(Θ | Θ(*t*)) is defined fixedly in E-step, most variants of EM algorithm focus on how to maximize *Q*(Θ’ | Θ) in M-step more effectively so that EM is faster or more accurate. Figure 2.1 (Borman, 2004, p. 7) shows relationship between the log-likelihood function *L*(Θ)and its lower-bound *lb*(Θ | Θ(*t*)).

Diagram

Description automatically generated

**Figure 2.1.** Relationship between the log-likelihood function and its lower-bound

Now ideology of GEM is explained in detail ■

The next section focuses on convergence of GEM algorithm proved by DLR (Dempster, Laird, & Rubin, 1977, pp. 7-10) but firstly we should discuss some features of *Q*(Θ’ | Θ). In special case of exponential family, *Q*(Θ’ | Θ) is modified by equation 2.9.

|  |  |
| --- | --- |
|  | (2.9) |

Where,

Following is a proof of equation 2.9.

Because *k*(*X* | *Y*, Θ) belongs exponential family, the expectation *E*(*τ*(*X*) | *Y*, Θ) is function of Θ, denoted *τ*Θ. It implies:

If *f*(*X*|Θ) belongs to regular exponential family, *Q*(Θ’ | Θ) gets maximal at the stationary point Θ\* so that the first-order derivative of *Q*(Θ’ | Θ) is zero. By referring to table 1.2, the first-order derivative of *Q*(Θ’ | Θ) with regard to Θ’ is:

Let *τ*(*t*) be the value of *τ*Θ at the *t*th iteration.

The equation above is indeed equation 2.6. The next parameter Θ(*t*+1) is determined at M-step as solution of the following equation.

This implies

The equation above is indeed equation 2.3. If *f*(*X*|Θ) belongs to curved exponential family, Θ(*t*+1) is determined as follows:

The equation above is indeed equation 2.7. Therefore, GEM shown in table 2.3 degrades into EM shown in table 2.1 and table 2.2 if *f*(*X*|Θ) belongs to exponential family. Of course, this recognition is trivial. Example 1.1 is also a good example for GEM when multinomial distribution belongs to exponential family and then we apply equation 2.7 into maximizing *Q*(Θ’ | Θ).

In practice, if *Y* is observed as particular *N* observations *Y*1, *Y*2,…, *YN*. Let = {*Y*1, *Y*2,…, *YN*} be the observed sample of size *N* with note that all *Yi* (s) are mutually independent and identically distributed (iid). Given an observation *Yi*, there is an associated random variable *Xi*. All *Xi* (s) are iid and they are not existent in fact. Each is a random variable like *X*. Of course, the domain of each *Xi* is ***X***. Let = {*X*1, *X*2,…, *XN*} be the set of associated random variables. Because all *Xi* (s) are iid, the joint PDF of is determined as follows:

Because all *Xi* (s) are iid and each *Yi* is associated with *Xi*, the conditional joint PDF of given is determined as follows:

The conditional expectation *Q*(Θ’ | Θ) given samples ***X*** and ***Y*** is determined as follows:

(Suppose *f*(*Xi* | Θ) and *k*(*Xj* | *Yj*, Θ) are analytic functions)

(Suppose *f*(*Xi* | Θ) and *k*(*Xj* | *Yj*, Θ) are analytic functions)

Like taking Riemann integral on , we have:

As a result, the conditional expectation *Q*(Θ’ | Θ) given an observed sample = {*Y*1, *Y*2,…, *YN*} and a set of associated random variables = {*X*1, *X*2,…, *XN*} is specified as follows:

Note, all *Xi* (s) are iid and they are not existent in fact. Because all *Xi* are iid, let *X* be the random variable representing every *Xi* and the equation of *Q*(Θ’ | Θ) is re-written according to equation 2.10.

|  |  |
| --- | --- |
|  | (2.10) |

The similar proof of equation 2.10 in case that *Xi* (s) are discrete is found in (Bilmes, 1998, p. 4). If *X* and all *Yi* (s) are discrete, equation 2.10 can be re-written as follows:

In case that *f*(*X* | Θ) and *k*(*X* | *Yi*, Θ) belong to exponential family, equation 2.10 becomes equation 2.11 with an observed sample = {*Y*1, *Y*2,…, *YN*}.

|  |  |
| --- | --- |
|  | (2.11) |

Where,

Please combine equation 2.9 and equation 2.10 to comprehend how to derive equation 2.11. Note, is dependent on both Θ and *Yi*.

DLR (Dempster, Laird, & Rubin, 1977, p. 1) called ***X*** as *complete data* because the mapping *φ*: ***X*** → ***Y*** is many-one function. There is another case that the complete space ***Z*** consists of hidden space ***X*** and observed space ***Y*** with note that ***X*** and ***Y*** are separated. There is no explicit mapping *φ* from ***X*** and ***Y*** but there exists a PDF of as the joint PDF of and .

In this case, the equation 2.8 is modified with the joint PDF *f*(*X*, *Y* | Θ). The PDF of *Y* becomes:

The PDF *f*(*Y*|Θ) is equivalent to the PDF *g*(*Y*|Θ) mentioned in equation 1.34. Although there is no explicit mapping from ***X*** to ***Y***, the PDF of *Y* above implies an implicit mapping from ***Z*** to ***Y***. The conditional PDF of *X* given *Z* is specified according to Bayes’ rule as follows:

The conditional PDF *f*(*X*|*Y*, Θ) is equivalent to the conditional PDF *k*(*X*|*Y*, Θ) mentioned in equation 1.35. Of course, given *Y*, we always have:

Equation 2.12 specifies the conditional expectation *Q*(Θ’ | Θ) in case that there is no explicit mapping from ***X*** to ***Y*** but there exists the joint PDF of *X* and *Y*.

|  |  |
| --- | --- |
|  | (2.12) |

Where,

Note, ***X*** is separated from ***Y*** and the complete data ***Z*** = (***X***, ***Y***) is composed of ***X*** and ***Y***. For equation 2.12, the existence of the joint PDF *f*(*X*, *Y* | Θ) can be replaced by the existence of the conditional PDF *f*(*Y*|*X*, Θ) and the prior PDF *f*(*X*|Θ) due to:

In applied statistics, equation 2.8 is often replaced by equation 2.12 because specifying the joint PDF *f*(*X*, *Y* | Θ) is more practical than specifying the mapping *φ*: ***X*** → ***Y***. However, equation 2.8 is more general equation 2.12 because the requirement of the joint PDF for equation 2.12 is stricter than the requirement of the explicit mapping for equation 2.8. In case that *X* and *Y* are discrete, equation 2.12 becomes:

In case that *X* and *Y* are discrete, *P*(*X*, *Y* | Θ) is the joint probability of *X* and *Y* whereas *P*(*X* | *Y*, Θ) is the conditional probability of *X* given *Y*.

Equation 2.12 can be proved alternately without knowledge related to complete data (Sean, 2009). This proof is like the proof of equation 2.8. In fact, given hidden space ***X***, observed space ***Y***, and a joint PDF *f*(*X*, *Y* | Θ), the likelihood function *L*(Θ’) is re-defined here as log(*f*(*Y* | Θ’)). The maximizer is:

Suppose the current parameter is Θ after some iteration. Next we must find out the new estimate Θ\* that maximizes the next log-likelihood function *L*(Θ’). Suppose the total probability of observed data can be determined by marginalizing over hidden data:

The expansion of *f*(*Y* | Θ’)is total probability rule. The next log-likelihood function *L*(Θ’) is re-written:

Because hidden *X* is the complete set of mutually exclusive variables, the sum of conditional probabilities of *X* is equal to 1 given *Y* and Θ.

Where,

Applying Jensen’s inequality (Sean, 2009, pp. 3-4) with concavity of logarithm function

into *L*(Θ’), we have (Sean, 2009, p. 6):

Where,

Obviously, the lower-bound of *L*(Θ’) is:

As aforementioned, the lower-bound *lb*(Θ’|Θ) (Sean, 2009, pp. 7-8) is maximized over many iterations of the iterative process so that *L*(Θ’) is maximized finally. Because *H*(Θ|Θ) is constant with regard to Θ’, it is possible to eliminate *H*(Θ|Θ) so that maximizing *Q*(Θ’|Θ) is the same to maximizing the lower bound. In final, when GEM converges Θ(*t*) = Θ(*t*+1) = Θ\*, we have:

We have the proof ■

Mixture model mentioned in subsection 5.1 is a good example for GEM without explicit mapping from ***X*** to ***Y***. Another well-known example is three-coin toss example (Collins & Barzilay, 2005) which applies GEM into estimating parameters of binomial distributions without explicit mapping.

**Example 2.1.** There are three coins named coin 1, coin 2 and coin 3. Each coin has two sides such as head (*H*) side and tail (*T*) side. Let hidden random variable *X* represent coin 1 where *X* is binary (*X* = {*H*, *T*}). Let *θ*1 be probability of coin 1 receiving head side.

*θ*1 = *P*(*X*=*H*)

Of course, we have:

*P*(*X*=*T*) = 1 – *θ*1

Let observed random variable *Y* represent a sequence of tossing coin 2 or coin 3 three times. Such sequence depends on first tossing coin 1. For instance, if coin 1 shows head side (*X*=*H*), the sequence is result of tossing coin 2 three times. Otherwise, if coin 1 shows tail side (*X*=*T*), the sequence is result of tossing coin 3 three times. For example, suppose first tossing coin 1 results *X*=*H* then, a possible result *Y* = *HHT* means that we toss coin 2 three times resulting head, head, and tail from coin 2. Obviously, *X* is hidden and *Y* is observed. In this example, we observe that

*Y*=*HHT*

Suppose *Y* conforms binomial distribution as follows:

Where *θ*2 and *θ*3 are probabilities of coin 2 and coin 3 receiving head side, respectively. Note, *h* is the number of head side from trials of tossing coin 2 (if *X*=*H*) or coin 3 (if *X*=*T*).Similarly, *t* is the number of tail side from trials of tossing coin 2 (if *X*=*H*) or coin 3 (if *X*=*T*). The joint probability *P*(*X*, *Y*) is:

In short, we need to estimate Θ = (*θ*1, *θ*2, *θ*3)*T* from the observation *Y*=*HHT* by discrete version of *Q*(Θ’ | Θ). Given *Y*=*HHT*, we have *h*=2 and *t*=1. Thus, the probability *P*(*Y*|*X*) becomes:

The joint probability *P*(*X*, *Y*) becomes:

The probability of *Y* is calculated as follows:

The conditional probability of *X* given *Y* is determined as follows:

The discrete version of *Q*(Θ’ | Θ) is determined as follows:

Note, *Q*(Θ’|Θ) is function of Θ’ = (*θ*1’, *θ*2’, *θ*3’)*T*. The next parameter Θ(*t*+1) = (*θ*1(*t*+1), *θ*2(*t*+1), *θ*3(*t*+1))*T* is a maximizer of *Q*(Θ’|Θ) with regard to Θ’, which is solution of the equation created by setting the first-order derivative of *Q*(Θ’|Θ) to be zero with note that the current parameter is Θ(*t*) = Θ.

The first-order partial derivative of *Q*(Θ’|Θ) with regard to *θ*1’ is:

Setting this partial derivative to be zero, we obtain:

Therefore, in M-step, given current parameter Θ(*t*) = (*θ*1(*t*), *θ*2(*t*), *θ*3(*t*))*T*, the next partial parameter *θ*1(*t*+1) is calculated as follows:

The first-order partial derivative of *Q*(Θ’|Θ) with regard to *θ*2’ is:

Setting this partial derivative to be zero, we obtain:

Therefore, in M-step, given current parameter Θ(*t*) = (*θ*1(*t*), *θ*2(*t*), *θ*3(*t*))*T*, the next partial parameter *θ*2(*t*+1) is fixed:

The first-order partial derivative of *Q*(Θ’|Θ) with regard to *θ*3’ is:

Setting this partial derivative to be zero, we obtain:

Therefore, in M-step, given current parameter Θ(*t*) = (*θ*1(*t*), *θ*2(*t*), *θ*3(*t*))*T*, the next partial parameter *θ*3(*t*+1) is fixed:

In short, in M-step of some *t*th iteration, given current parameter Θ(*t*) = (*θ*1(*t*), *θ*2(*t*), *θ*3(*t*))*T*, only *θ*1(*t*+1) is updated whereas both *θ*2(*t*+1) and *θ*3(*t*+1) are fixed with observation *Y*=*HHT*.

For instance, let Θ(1) = (*θ*1(1), *θ*2(1), *θ*3(1))*T* be initialized arbitrarily as *θ*1(1) = *θ*2(1) = *θ*3(1) = 0.5, at the first iteration, we obtain:

At the second iteration with current parameter Θ(2) = (*θ*1(2)=0.5, *θ*2(2)=2/3, *θ*3(2)=2/3)*T*, we obtain:

As a result, GEM inside this example converges at the second iteration with final estimate Θ(2) = Θ(3) = Θ\* = (*θ*1\*=0.5, *θ*2\*=2/3, *θ*3\*=2/3)*T* ■

In practice, suppose *Y* is observed as a sample = {*Y*1, *Y*2,…, *YN*} of size *N* with note that all *Yi* (s) are mutually independent and identically distributed (iid). The observed sample is associated with a a hidden set (latent set) = {*X*1, *X*2,…, *XN*} of size *N*. All *Xi* (s) are iid and they are not existent in fact. Let be the random variable representing every *Xi*. Of course, the domain of *X* is ***X***. Equation 2.13 specifies the conditional expectation *Q*(Θ’ | Θ) given such .

|  |  |
| --- | --- |
|  | (2.13) |

Equation 2.13 is a variant of equation 2.10 in case that there is no explicit mapping between *Xi* and *Yi* but there exists the same joint PDF between *Xi* and *Yi*. Please see the proof of equation 2.10 to comprehend how to derive equation 2.13. If both *X* and *Y* are discrete, equation 2.13 becomes:

|  |  |
| --- | --- |
|  | (2.14) |

If *X* is discrete and *Y* is continuous such that *f*(*X*, *Y* | Θ) = *P*(*X*|Θ)*f*(*Y* | *X*, Θ) then, according to the total probability rule, we have:

Note, when only *X* is discrete, its PDF *f*(*X*|Θ) becomes the probability *P*(*X*|Θ). Therefore, equation 2.15 is a variant of equation 2.13, as follows:

|  |  |
| --- | --- |
|  | (2.15) |

Where *P*(*X* | *Yi*, Θ) is determined by Bayes’ rule, as follows:

Equation 2.15 is the base for estimating the probabilistic mixture model by EM algorithm, which will be described later in detail. Some other properties of GEM will be mentioned in next section.

# **3. Properties and convergence of EM algorithm**

Recall that DLR proposed GEM algorithm which aims to maximize the log-likelihood function *L*(Θ) by maximizing *Q*(Θ’ | Θ) over many iterations. This section focuses on mathematical explanation of the convergence of GEM algorithm given by DLR (Dempster, Laird, & Rubin, 1977, pp. 6-9). Recall that we have:

Let *H*(Θ’ | Θ) be another conditional expectation which has strong relationship with *Q*(Θ’ | Θ) (Dempster, Laird, & Rubin, 1977, p. 6).

|  |  |
| --- | --- |
|  | (3.1) |

If there is no explicit mapping from ***X*** to ***Y*** but there exists a joint PDF *f*(*X*, *Y* | Θ) of *X* and *Y*, equation 3.1 can be re-written as follows:

Where,

From equation 2.8 and equation 3.1, we have:

|  |  |
| --- | --- |
|  | (3.2) |

Following is a proof of equation 3.2.

**Lemma3.1** (Dempster, Laird, & Rubin, 1977, p. 6). For any pair (Θ’, Θ) in Ω x Ω,

|  |  |
| --- | --- |
|  | (3.3) |

The equality occurs if and only if *k*(*X* | *Y*, Θ’) = *k*(*X* | *Y*, Θ) almost everywhere ■

Following is a proof of lemma 3.1 as well as equation 3.3. The log-likelihood function *L*(Θ’) is re-written as follows:

Due to

By applying Jensen’s inequality (Sean, 2009, pp. 3-4) with concavity of logarithm function

into *L*(Θ’), we have (Sean, 2009, p. 6):

(Due to *Q*(Θ’|Θ) = *L*(Θ’) + *H*(Θ’|Θ))

It implies:

According to Jensen’s inequality (Sean, 2009, pp. 3-4), the equality *H*(Θ’|Θ) = *H*(Θ|Θ) occurs if and only if *k*(*X* | *Y*, Θ’) is linear or *f*(*X* | Θ’) is constant. In other words, the equality occurs if and only if *k*(*X* | *Y*, Θ’) = *k*(*X* | *Y*, Θ) almost everywhere when *f*(*X* | Θ) is not constant and *k*(*X* | *Y*, Θ’) is a PDF ■

We also have the lower-bound of *L*(Θ’), denoted *lb*(Θ’|Θ) as follows:

*lb*(Θ’|Θ) = *Q*(Θ’|Θ) – *H*(Θ|Θ)

Obviously, we have:

*L*(Θ’) ≥ *lb*(Θ’|Θ)

As aforementioned, the lower-bound *lb*(Θ’|Θ) is maximized over many iterations of the iterative process so that *L*(Θ’) is maximized finally. Such lower-bound is determined indirectly by *Q*(Θ’|Θ) so that maximizing *Q*(Θ’|Θ) with regard to Θ’ is the same to maximizing *lb*(Θ’|Θ) because *H*(Θ|Θ) is constant with regard to Θ’.

Let be a sequence of estimates of Θ resulted from iterations of EM algorithm. Let Θ → *M*(Θ) be the mapping such that each estimation Θ(*t*) → Θ(*t*+1) at any given iteration is defined by equation 3.4 (Dempster, Laird, & Rubin, 1977, p. 7).

|  |  |
| --- | --- |
|  | (3.4) |

**Definition 3.1** (Dempster, Laird, & Rubin, 1977, p. 7). An iterative algorithm with mapping *M*(Θ) is a GEM algorithm if

|  |  |
| --- | --- |
|  | (3.5) |

Of course, specification of GEM shown in table 2.3 satisfies the definition 3.1 because Θ(*t*+1) is a maximizer of *Q*(Θ | Θ(*t*)) with regard to variable Θ in M-step.

**Theorem 3.1** (Dempster, Laird, & Rubin, 1977, p. 7). For every GEM algorithm

|  |  |
| --- | --- |
|  | (3.6) |

Where equality occurs if and only if *Q*(*M*(Θ) | Θ) = *Q*(Θ | Θ) and *k*(*X* | *Y*, *M*(Θ)) = *k*(*X* | *Y*, Θ) almost everywhere ■

Following is the proof of theorem 3.1 (Dempster, Laird, & Rubin, 1977, p. 7):

Because the equality of lemma 3.1 occurs if and only if *k*(*X* | *Y*, Θ’) = *k*(*X* | *Y*, Θ) almost everywhere and the equality of the definition 3.1 is *Q*(*M*(Θ) | Θ) = *Q*(Θ | Θ), we deduce that the equality of theorem 3.1 occurs if and only if *Q*(*M*(Θ) | Θ) = *Q*(Θ | Θ) and *k*(*X* | *Y*, *M*(Θ)) = *k*(*X* | *Y*, Θ) almost everywhere. It is easy to draw corollary 3.1 and corollary 3.2 from definition 3.1 and theorem 3.1.

**Corollary 3.1** (Dempster, Laird, & Rubin, 1977). Suppose for some , *L*(Θ*\**) ≥ *L*(Θ) for all then for every GEM algorithm:

1. *L*(*M*(Θ*\**)) = *L*(Θ*\**)
2. *Q*(*M*(Θ*\**) | Θ*\**) = *Q*(Θ*\** | Θ*\**)
3. *k*(*X* | *Y*, *M*(Θ*\**)) = *k*(*X* | *Y*, Θ*\**) ■

*Proof*. From theorem 3.1 and the assumption of corollary 3.1, we have:

This implies:

As a result,

From theorem 3.1, we also have:

**Corollary 3.2** (Dempster, Laird, & Rubin, 1977). If for some , *L*(Θ*\**) > *L*(Θ) for all such that Θ ≠ Θ\*, then for every GEM algorithm:

*M*(Θ*\**) = Θ*\** ■

*Proof*. From corollary 3.1 and the assumption of corollary 3.2, we have:

If *M*(Θ*\**) ≠ Θ*\**, there is a contradiction *L*(*M*(Θ*\**)) = *L*(Θ*\**) > *L*(*M*(Θ*\**)). Therefore, we have *M*(Θ*\**) = Θ*\** ■

**Theorem 3.2** (Dempster, Laird, & Rubin, 1977, p. 7). Suppose is the sequence of estimates resulted from GEM algorithm such that:

1. The sequence is bounded above, and
2. *Q*(Θ(*t*+1) | Θ(*t*)) – *Q*(Θ(*t*) | Θ(*t*)) ≥ *ξ*(Θ(*t*+1) – Θ(*t*))*T*(Θ(*t*+1) – Θ(*t*)) for some scalar *ξ* > 0 and all *t*.

Then the sequence converges to some Θ*\** in the closure of Ω ■

*Proof*. The sequence is non-decreasing according to theorem 3.1 and is bounded above according to the assumption 1 of theorem 3.2 and hence, the sequence converges to some *L\** < +∞. According to Cauchy criterion (Dinh, Pham, Nguyen, & Ta, 2000, p. 34), for all *ε* > 0, there exists a *t*(*ε*) such that, for all *t* ≥ *t*(*ε*) and all *v* ≥ 1:

By applying equation 3.2 and equation 3.3, for all *i* ≥ 1, we obtain:

(Due to *L*(Θ(*t*+*i*–1)) = *Q*(Θ(*t*+*i*–1) | Θ(*t*+*i*–1)) – *H*(Θ(*t*+*i*–1) | Θ(*t*+*i*–1)) according to equation 3.2)

It implies

By applying *v* times the assumption 2 of theorem 3.2, we obtain:

It means that

Where,

Notation |.| denotes length of vector and so |Θ(*t*+*i*) – Θ(*t+i* –1)| is distance between Θ(*t*+*i*) and Θ(*t*+*i* –1). Applying triangular inequality, for any *ε* > 0, for all *t* ≥ *t*(*ε*) and all *v* ≥ 1, we have:

According to Cauchy criterion, the sequence converges to some Θ*\** in the closure of Ω.

Theorem 3.1 indicates that *L*(Θ) is non-decreasing on every iteration of GEM algorithm and is strictly increasing on any iteration such that *Q*(Θ(*t*+1) | Θ(*t*)) > *Q*(Θ(*t*) | Θ(*t*)). The corollaries 3.1 and 3.2 indicate that the optimal estimate is a fixed point of GEM algorithm. Theorem 3.2 points out convergence condition of GEM algorithm but does not assert the converged point Θ\* is maximizer of *L*(Θ). So, we need mathematical tools of derivative and differential to prove convergence of GEM to a maximizer Θ\*. We assume that *Q*(Θ’ | Θ), *L*(Θ), *H*(Θ’ | Θ), and *M*(Θ) are smooth enough. As a convention for derivatives of bivariate function, let *Dij* denote as the derivative (differential) by taking *i*th-order partial derivative (differential) with regard to first variable and then, taking *j*th-order partial derivative (differential) with regard to second variable. If *i* = 0 (*j* = 0) then, there is no partial derivative with regard to first variable (second variable). For example, following is an example of how to calculate the derivative *D*11*Q*(Θ(*t*) | Θ(*t*+1)).

* Firstly, we determine
* Secondly, we substitute Θ(*t*) and Θ(*t*+1) for such *D*11*Q*(Θ’ | Θ) to obtain *D*11*Q*(Θ(*t*) | Θ(*t*+1)).

Equation 3.1 shows some derivatives (differentials) of *Q*(Θ’ | Θ), *H*(Θ’ | Θ), *L*(Θ), and *M*(Θ).

|  |
| --- |
|  |

**Table 3.1.** Some differentials of *Q*(Θ’ | Θ), *H*(Θ’ | Θ), *L*(Θ), and *M*(Θ)

When Θ’ and Θ are vectors, *D*10(…) is gradient vector and *D*20(…) is Hessian matrix. As a convention, let **0** = (0, 0,…, 0)*T* be zero vector.

**Lemma 3.2** (Dempster, Laird, & Rubin, 1977, p. 8). For all Θ in Ω,

|  |  |
| --- | --- |
|  | (3.7) |

|  |  |
| --- | --- |
|  | (3.8) |

|  |  |
| --- | --- |
|  | (3.9) |

|  |  |
| --- | --- |
|  | (3.10) |

|  |  |
| --- | --- |
|  | (3.11) |

|  |  |
| --- | --- |
|  | (3.12) |

Note, *VN*(.) denotes non-central variance (non-central covariance matrix). Followings are proofs of equation 3.7, equation 3.8, equation 3.9, equation 3.10, equation 3.11, and equation 3.12. In fact, we have:

It implies:

Thus, equation 3.7 is proved.

We also have:

It implies:

We also have:

It implies:

Hence, equation 3.8 and equation 3.9 are proved.

From equation 3.2, we have:

We also have:

It implies:

Thus, equation 3.10 is proved.

We have:

(Hence, equation 3.11 is proved)

It implies:

Due to:

We have:

Therefore, equation 3.12 is proved ■

**Lemma 3.3** (Dempster, Laird, & Rubin, 1977, p. 9). If *f*(*X* | Θ) and *k*(*X* | *Y*, Θ) belong to exponential family, for all Θ in Ω, we have:

|  |  |
| --- | --- |
|  | (3.13) |

|  |  |
| --- | --- |
|  | (3.14) |

|  |  |
| --- | --- |
|  | (3.15) |

|  |  |
| --- | --- |
|  | (3.16) |

*Proof*. If *f*(*X* | Θ’) and *k*(*X* | *Y*, Θ’) belong to exponential family, from table 1.2 we have:

And,

And,

And,

Hence,

Thus, equation 3.13 is proved.

We have:

Thus, equation 3.14 is proved.

We have:

Thus, equation 3.15 is proved.

We have:

Thus, equation 3.16 is proved ■

**Theorem 3.3** (Dempster, Laird, & Rubin, 1977, p. 8). Suppose the sequence is an instance of GEM algorithm such that

Then for all *t*, there exists a Θ0(*t*+1) on the line segment joining Θ(*t*) and Θ(*t*+1) such that

Furthermore, if *D*20*Q*(Θ0(*t*+1) | Θ(*t*)) is negative definite, and the sequence is bounded above then, the sequence converges to some Θ*\** in the closure of Ω ■

Note, if Θ is a scalar parameter, *D*20*Q*(Θ0(*t*+1) | Θ(*t*)) degrades as a scalar and the concept “negative definite” becomes “negative” simply. Following is a proof of theorem 3.3.

*Proof*. Second-order Taylor series expending for *Q*(Θ | Θ(*t*)) at Θ = Θ(*t*+1) to obtain:

Where Θ0(*t*+1) is on the line segment joining Θ and Θ(*t*+1). Let Θ = Θ(*t*), we have:

If *D*20*Q*(Θ(*t*+1) | Θ(*t*)) is negative definite then,

Whereas,

So, for all *t*, there exists some *ξ* > 0 such that

In other words, the assumption 2 of theorem 3.2 is satisfied and hence, the sequence converges to some Θ*\** in the closure of Ω if the sequence is bounded above ■

**Theorem 3.4** (Dempster, Laird, & Rubin, 1977, p. 9). Suppose the sequence is an instance of GEM algorithm such that

1. The sequence converges to Θ*\** in the closure of Ω.
2. *D*10*Q*(Θ(*t*+1) | Θ(*t*)) = **0***T* for all *t*.
3. *D*20*Q*(Θ(*t*+1) | Θ(*t*)) is negative definite for all *t*.

Then *DL*(Θ*\**) = **0***T*, *D*20*Q*(Θ*\** | Θ*\**) is negative definite, and

|  |  |
| --- | --- |
|  | (3.17) |

The notation “–1” denotes inverse of matrix. Note, *DM*(Θ*\**) is differential of *M*(Θ) at Θ = Θ\*, which implies convergence rate of GEM algorithm. Obviously, Θ\* is local maximizer due to *DL*(Θ*\**) = **0***T* and *D*20*Q*(Θ*\** | Θ*\**). Followings are proofs of theorem 3.4.

From equation 3.2, we have:

When *t* approaches +∞ such that Θ(*t*) = Θ(*t*+1) = Θ*\** then, *D*10*H*(Θ*\** | Θ*\**) is zero according to equation 3.7 and so we have:

*DL*(Θ*\**) = **0***T*

Of course, *D*20*Q*(Θ*\** | Θ*\**) is negative definite because *D*20*Q*(Θ(*t*+1) | Θ(*t*)) is negative definite, when *t* approaches +∞ such that Θ(*t*) = Θ(*t*+1) = Θ*\**.

By first-order Taylor series expansion for *D*10*Q*(Θ2 | Θ1) as a function of Θ1 at Θ1 = Θ*\** and as a function of Θ2 at Θ2 = Θ*\**, respectively, we have:

Where *R*1(Θ1) and *R*2(Θ2) are remainders. By summing such two series, we have:

By substituting Θ1 = Θ(*t*) and Θ2 = Θ(*t*+1), we have:

Due to *D*10*Q*(Θ(*t*+1) | Θ(*t*)) = **0***T*, we obtain:

It implies:

Multiplying two sides of the equation above by *D*20*Q*(Θ*\** | Θ(*t*))–1 and letting *M*(Θ(*t*)) = Θ(*t*+1), *M*(Θ\*) = Θ\*, we obtain:

Let *t* approach +∞ such that Θ(*t*) = Θ(*t*+1) = Θ*\**, we obtain *DM*(Θ*\**) as differential of *M*(Θ) at Θ*\** as follows:

|  |  |
| --- | --- |
|  | (3.18) |

Due to, when *t* approaches +∞, we have:

The derivative *D*11*Q*(Θ’ | Θ) is expended as follows:

It implies:

(Due to theorem 3.4)

(Due to equation 3.8)

Therefore, equation 3.18 becomes equation 3.17.

Finally, theorem 3.4 is proved. By combination of theorems 3.2 and 3.4, I propose corollary 3.3 as a convergence criterion to local maximizer of GEM.

**Corollary 3.3.** If an algorithm satisfies three following assumptions:

1. *Q*(*M*(Θ(*t*)) | Θ(*t*)) > *Q*(Θ(*t*) | Θ(*t*)) for all *t*.
2. The sequence is bounded above.
3. *D*10*Q*(Θ*\** | Θ*\**) = **0***T* and *D*20*Q*(Θ*\** | Θ*\**) negative definite with suppose that Θ*\** is the converged point.

Then,

1. Such algorithm is an GEM and converges to a local maximizer Θ*\** of *L*(Θ) such that *DL*(Θ*\**) = **0***T* and *D*2*L*(Θ*\**) negative definite.
2. Equation 3.17 is obtained ■

The assumption 1 of corollary 3.3 implies that the given algorithm is a GEM according to definition 3.1. From such assumption, we also have:

So there exists some *ξ* > 0 such that

In other words, the assumption 2 of theorem 3.2 is satisfied and hence, the sequence converges to some Θ*\** in the closure of Ω when the sequence is bounded above according to the assumption 2 of corollary 3.3. From equation 3.2, we have:

When *t* approaches +∞ such that Θ(*t*) = Θ(*t*+1) = Θ*\** then,

*DL*(Θ*\**) = *D*10*Q*(Θ*\** | Θ*\**) – *D*10*H*(Θ*\** | Θ*\**)

*D*10*H*(Θ*\** | Θ*\**) is zero according to equation 3.7. Hence, along with the assumption 3 of corollary 3.3, we have:

*DL*(Θ*\**) = *D*10*Q*(Θ*\** | Θ*\**) = **0***T*

Due to *DL*(Θ*\**) = 0, we only assert here that the given algorithm converges to Θ*\** as a stationary point of *L*(Θ). Later on, we will prove that Θ*\** is a local maximizer of *L*(Θ) when *Q*(*M*(Θ(*t*)) | Θ(*t*)) > *Q*(Θ(*t*) | Θ(*t*)), *DL*(Θ*\**) = 0, and *D*20*Q*(Θ*\** | Θ*\**) negative definite. Due to *D*10*Q*(Θ*\** | Θ*\**) = **0***T*, we obtain equation 3.17. Please see the proof of equation 3.17 ■

By default, suppose all GEM algorithms satisfy the assumptions 2 and 3 of corollary 3.3. Thus, we only check the assumption 1 to verify whether a given algorithm is a GEM which converges to local maximizer Θ*\**. Note, if the assumption 1 of corollary 3.3 is replaced by “*Q*(*M*(Θ(*t*)) | Θ(*t*)) ≥ *Q*(Θ(*t*) | Θ(*t*)) for all *t*” then, Θ*\** is only asserted to be a stationary point of *L*(Θ) such that *DL*(Θ*\**) = **0***T*. Wu (Wu, 1983) gave a deep research on convergence of GEM in her/his article “On the Convergence Properties of the EM Algorithm”. Please read this article for more details about convergence of GEM.

Because *H*(Θ’ | Θ) and *Q*(Θ’ | Θ) are smooth enough, *D*20*H*(Θ*\** | Θ*\**) and *D*20*Q*(Θ*\** | Θ*\**) are symmetric matrices according to Schwarz’s theorem (Wikipedia, Symmetry of second derivatives, 2018). Thus, *D*20*H*(Θ*\** | Θ*\**) and *D*20*Q*(Θ*\** | Θ*\**) are commutative:

*D*20*H*(Θ*\** | Θ*\**)*D*20*Q*(Θ*\** | Θ*\**) = *D*20*Q*(Θ*\** | Θ*\**)*H*20*Q*(Θ*\** | Θ*\**)

Suppose both *D*20*H*(Θ*\** | Θ*\**) and *D*20*Q*(Θ*\** | Θ*\**) are diagonalizable then, they are simultaneously diagonalizable (Wikipedia, Commuting matrices, 2017). Hence there is an (orthogonal) eigenvector matrix *U* such that (Wikipedia, Diagonalizable matrix, 2017) (StackExchange, 2013):

Where *He\** and *Qe\** are eigenvalue matrices of *D*20*H*(Θ*\** | Θ*\**) and *D*20*Q*(Θ*\** | Θ*\**), respectively, according to equation 3.19 and equation 3.20. Of course, *h*1*\**, *h*2*\**,…, *hr\** are eigenvalues of *D*20*H*(Θ*\** | Θ*\**) whereas *q*1*\**, *q*2*\**,…, *qr\** are eigenvalues of *D*20*Q*(Θ*\** | Θ*\**).

|  |  |
| --- | --- |
|  | (3.19) |

|  |  |
| --- | --- |
|  | (3.20) |

From equation 3.17, *DM*(Θ*\**) is decomposed as seen in equation 3.21.

|  |  |
| --- | --- |
|  | (3.21) |

Let *Me\** be eigenvalue matrix of *DM*(Θ*\**), specified by equation 3.17. As a convention *Me\** is called convergence matrix.

|  |  |
| --- | --- |
|  | (3.22) |

Of course, all *mi\** = *hi\** / *qi\** are eigenvalues of *DM*(Θ*\**) with assumption *qi\** < 0 for all *i*. We will prove that 0 ≤ *mi\** ≤ 1 for all *i* by contradiction. Conversely, suppose we *always* have *mi\** > 1 or *mi\** < 0 for some *i*. When Θ degrades into scalar as Θ = *θ* with note that scalar is 1-element vector, equation 3.17 is re-written as equation 3.23:

|  |  |
| --- | --- |
|  | (3.23) |

From equation 3.23, the next estimate *θ*(*t*+1) approaches*θ\** when *t* → +∞ and so we have:

So equation 3.24 is a variant of equation 3.23 (McLachlan & Krishnan, 1997, p. 120).

|  |  |
| --- | --- |
|  | (3.24) |

Because the sequence is non-decreasing, the sequence is monotonous. This means:

Or

It implies

So we have

However, this contradicts the converse assumption “there always exists *mi\** > 1 or *mi\** < 0 for some *i*”. Therefore, we conclude that 0 ≤ *mi\** ≤ 1 for all *i*. In general, if Θ*\** is stationary point of GEM then*, D*20*Q*(Θ*\** | Θ*\**) and *Qe\** are negative definite, *D*20*H*(Θ*\** | Θ*\**) and *He\** are negative semi-definite, and *DM*(Θ*\**) and *Me\** are positive semi-definite, according to equation 3.25.

|  |  |
| --- | --- |
|  | (3.25) |

As a convention, if GEM algorithm fortunately stops at the first iteration such that Θ(1) = Θ(2) = Θ*\** then, *mi\** = 0 for all *i*.

Suppose Θ(*t*) = (*θ*1(*t*), *θ*2(*t*),…, *θr*(*t*)) at current *t*th iteration and Θ*\** = (*θ*1*\**, *θ*2*\**,…, *θr\**), each *mi\** measures how much the next *θi*(*t*+1) is near to *θi*\*. In other words, the smaller the *mi\** (s) are, the faster the GEM is and so the better the GEM is. This is why DLR (Dempster, Laird, & Rubin, 1977, p. 10) defined that the convergence rate *m\** of GEM is the maximum one among all *mi\**, as seen in equation 3.26. The convergence rate *m\** implies lowest speed.

|  |  |
| --- | --- |
|  | (3.26) |

From equation 3.2 and equation 3.17, we have (Dempster, Laird, & Rubin, 1977, p. 10):

Where *I* is identity matrix:

By the same way to draw convergence matrix *Me\** with note that *D*20*H*(Θ*\** | Θ*\**), *D*20*Q*(Θ*\** | Θ*\**), and *DM*(Θ*\**) are symmetric matrices, we have:

|  |  |
| --- | --- |
|  | (3.27) |

Where *Le\** is eigenvalue matrix of *D*2*L*(Θ*\**). From equation 3.27, each eigenvalue *li\** of *Le\** is proportional to each eigenvalues *qi\** of *Qe\** with ratio 1–*mi\** where *mi\** is an eigenvalue of *Me\**. Equation 3.28 specifies a so-called speed matrix *Se\**:

|  |  |
| --- | --- |
|  | (3.28) |

This implies

From equation 3.25 and equation 3.28, we have 0 ≤ *si\** ≤ 1. Equation 3.29 specifies *Le\** which is eigenvalue matrix of *D*2*L*(Θ*\**).

|  |  |
| --- | --- |
|  | (3.29) |

From equation 3.28, suppose Θ(*t*) = (*θ*1(*t*), *θ*2(*t*),…, *θr*(*t*)) at current *t*th iteration and Θ*\** = (*θ*1*\**, *θ*2*\**,…, *θr\**), each *si\** = 1–*mi\** is really the speed that the next *θi*(*t*+1) moves to *θi*\*. From equation 3.26 and equation 3.28, equation 3.30 specifies the speed *s\** of GEM algorithm.

|  |  |
| --- | --- |
| Where, | (3.30) |

As a convention, if GEM algorithm fortunately stops at the first iteration such that Θ(1) = Θ(2) = Θ*\** then, *s\** = 1.

For example, when Θ degrades into scalar as Θ = *θ*, the fourth column of table 1.3 (Dempster, Laird, & Rubin, 1977, p. 3) gives sequences which approaches *Me\** = *DM*(*θ\**) through many iterations by the following ratio to determine the limit in equation 3.23 with *θ*\* = 0.6268.

In practice, if GEM is run step by step, *θ\** is not known yet at some *t*th iteration when GEM does not converge yet. Hence, equation 3.24 (McLachlan & Krishnan, 1997, p. 120) is used to make approximation of *Me\** = *DM*(*θ\**) with unknown *θ\** and *θ*(*t*) ≠ *θ*(*t*+1).

It is required only two successive iterations because both *θ*(*t*) and *θ*(*t*+1) are determined at *t*th iteration whereas *θ*(*t*+2) is determined at (*t*+1)th iteration. For example, in table 1.3, given *θ*(1) = 0.5, *θ*(2) = 0.6082, and *θ*(3) = 0.6243, at *t* = 1, we have:

Whereas the real *Me\** = *DM*(*θ\**) is 0.1465 shown in the fourth column of table 1.3 at *t* = 1.

We will prove by contradiction that if definition 3.1 is satisfied strictly such that *Q*(*M*(Θ(*t*)) | Θ(*t*)) > *Q*(Θ(*t*) | Θ(*t*)) then, *li\** < 0 for all *i*. Conversely, suppose we *always* have *li\** ≥ 0 for some *i* when *Q*(*M*(Θ(*t*)) | Θ(*t*)) > *Q*(Θ(*t*) | Θ(*t*)). Given Θ degrades into scalar as Θ = *θ* with note that scalar is 1-element vector, when *Q*(*M*(Θ(*t*)) | Θ(*t*)) > *Q*(Θ(*t*) | Θ(*t*)), the sequence is strictly increasing, which in turn causes that the sequence is strictly monotonous. This means:

Or

It implies

So we have

From equation 3.29, we deduce that *D*2*L*(*θ\**) = *Le\** = *Se\*Qe\** < 0 where *Qe\** = *D*20*Q*(*θ\** | *θ\**) < 0. However, this contradicts the converse assumption “there always exists *li\** ≥ 0 for some *i* when *Q*(*M*(Θ(*t*)) | Θ(*t*)) > *Q*(Θ(*t*) | Θ(*t*))”. Therefore, if *Q*(*M*(Θ(*t*)) | Θ(*t*)) > *Q*(Θ(*t*) | Θ(*t*)) then, *li\** < 0 for all *i.* In other words, at that time, *D*2*L*(Θ*\**) = *Le\** is negative definite. Recall that we proved that *DL*(Θ*\**) = 0 for corollary 3.3. Now we have *D*2*L*(Θ*\**) negative definite, which means that Θ*\** is a local maximizer of *L*(Θ*\**) in corollary 3.3. In other words, corollary 3.3 is proved.

Recall that *L*(Θ) is the log-likelihood function of observed *Y* according to equation 2.3.

Both –*D*20*H*(Θ*\** | Θ*\**) and –*D*20*Q*(Θ*\** | Θ*\**) are information matrices (Zivot, 2009, pp. 7-9) specified by equation 3.31.

|  |  |
| --- | --- |
|  | (3.31) |

*IH*(Θ*\**) measures information of *X* about Θ*\** with support of *Y* whereas *IQ*(Θ*\**) measures information of *X* about Θ*\**. In other words, *IH*(Θ*\**) measures observed information whereas *IQ*(Θ*\**) measures hidden information. Let *VH*(Θ*\**) and *VQ*(Θ*\**) be covariance matrices of Θ*\** with regard to *IH*(Θ*\**) and *IQ*(Θ*\**), respectively. They are inverses of *IH*(Θ*\**) and *IQ*(Θ*\**) according to equation 3.32 when Θ*\** is unbiased estimate.

|  |  |
| --- | --- |
|  | (3.32) |

Equation 3.33 is a variant of equation 3.17 to calculate *DM*(Θ*\**) based on information matrices:

|  |  |
| --- | --- |
|  | (3.33) |

If *f*(*X* | Θ), *g*(*Y* | Θ) and *k*(*X* | *Y*, Θ) belong to exponential family, from equation 3.14 and equation 3.16, we have:

Hence, equation 3.34 specifies *DM*(Θ*\**) in case of exponential family.

|  |  |
| --- | --- |
|  | (3.34) |

Equation 3.35 specifies relationships among *VH*(Θ*\**), *VQ*(Θ*\**), *V*(*τ*(*X*) | *Y*, Θ*\**), and *V*(*τ*(*X*) | Θ*\**) in case of exponential family.

|  |  |
| --- | --- |
|  | (3.35) |

# **4. Variants of EM algorithm**

The main purpose of EM algorithm (GEM algorithm) is to maximize the log-likelihood *L*(Θ) = log(*g*(*Y* | Θ)) with observed data *Y* by maximizing the condition expectation *Q*(Θ’ | Θ). Such *Q*(Θ’ | Θ) is defined fixedly in E-step. Therefore, most variants of EM algorithm focus on how to maximize *Q*(Θ’ | Θ) in M-step more effectively so that EM is faster or more accurate.

## **4.1. EM with prior probability**

DLR (Dempster, Laird, & Rubin, 1977, pp. 6, 11) mentioned that the convergence rate *DM*(Θ*\**) specified by equation 3.17 can be improved by adding a prior probability *π*(Θ) in conjugation with *f*(*X* | Θ), *g*(*Y* | Θ) or *k*(*X* | *Y*, Θ) according to maximum a posteriori probability (MAP) method (Wikipedia, Maximum a posteriori estimation, 2017). For example, if *π*(Θ) in conjugation with *g*(*Y* | Θ) then, the posterior probability *π*(Θ | *Y*) is:

Because is constant with regard to Θ, the optimal likelihood-maximization estimate Θ*\** is a maximizer of *g*(*Y* | Θ)*π*(Θ). When π(Θ) is conjugate prior of the posterior probability *π*(Θ | *X*) (or *π*(Θ | *Y*)), both *π*(Θ) and *π*(Θ | *X*) (or *π*(Θ | *Y*)) have the same distributions (Wikipedia, Conjugate prior, 2018); for example, if *π*(Θ) is distributed normally, *π*(Θ | *X*) (or *π*(Θ | *Y*)) is also distributed normally.

For GEM algorithm, the log-likelihood function associated MAP method is specified by equation 4.1.1 with note that *π*(Θ) is non-convex function.

|  |  |
| --- | --- |
|  | (4.1.1) |

It implies from equation 3.2 that

Let,

|  |  |
| --- | --- |
|  | (4.1.2) |

GEM algorithm now aims to maximize *Q*+(Θ’ | Θ) instead of maximizing *Q*(Θ’ | Θ). The proof of convergence for *Q*+(Θ’ | Θ) is not changed in manner but determining the convergence matrix *Me* for *Q*+(Θ’ | Θ) is necessary. Because *H*(Θ’ | Θ) is kept intact whereas *Q*(Θ’ | Θ) is replaced by *Q*+(Θ’ | Θ), we expect that the convergence rate *m\** specified by equation 3.26 is smaller so that the convergence speed *s\** is increased and so GEM algorithm is improved with regard to *Q*+(Θ’ | Θ). Equation 4.1.3 specifies *DM*(Θ*\**) for *Q*+(Θ’ | Θ).

|  |  |
| --- | --- |
|  | (4.1.3) |

Where *Q*+(Θ’ | Θ) is specified by equation 4.1.2 and *D*20*Q*+(Θ’ | Θ) is specified by equation 4.1.4.

|  |  |
| --- | --- |
|  | (4.1.4) |

Where,

Because *Q*(Θ’ | Θ) and *π*(Θ’) are smooth enough, *D*20*Q*(Θ*\** | Θ*\**) and *D*20*L*(*π*(Θ*\**)) are symmetric matrices according to Schwarz’s theorem (Wikipedia, Symmetry of second derivatives, 2018). Thus, *D*20*Q*(Θ*\** | Θ*\**) and *D*20*L*(*π*(Θ*\**)) are commutative:

*D*20*Q*(Θ*\** | Θ*\**)*D*20*L*(*π*(Θ*\**)) = *D*20*L*(*π*(Θ*\**))*D*20*Q*(Θ*\** | Θ*\**)

Suppose both *D*20*Q*(Θ*\** | Θ*\**) and *D*20*L*(*π*(Θ*\**)) are diagonalizable then, they are simultaneously diagonalizable (Wikipedia, Commuting matrices, 2017). Hence there is an (orthogonal) eigenvector matrix *V* such that (Wikipedia, Diagonalizable matrix, 2017) (StackExchange, 2013):

Where *Qe\** and Π*e\** are eigenvalue matrices of *D*20*Q*(Θ*\** | Θ*\**) and *D*20*L*(*π*(Θ*\**)), respectively. Note *Qe\** and its eigenvalues are mentioned in equation 3.20. Because *π*(Θ*\**) is non-convex function, eigenvalues *π*1*\**, *π*2*\**,…, *πr\** of Π*e\** are non-positive.

From equation 4.1.2, *D*20*Q*+(Θ*\** | Θ*\**) is decomposed as below:

So eigenvalue matrix of *D*20*Q*+(Θ*\** | Θ*\**) is (*Qe\** + Π*e\**) and eigenvalues of *D*20*Q*+(Θ*\** | Θ*\**) are *qi\** + *πi\**, as follows:

According to equation 3.19, the eigenvalue matrix of *D*20*H*(Θ*\** | Θ*\**) is *He\** fixed as follows:

Due to *DM*(Θ*\**) = *D*20*H*(Θ*\** | Θ*\**)*D*20*Q*+(Θ*\** | Θ*\**), equation 3.21 is re-calculated:

As a result, the convergence matrix *Me\** which is eigenvalue matrix of *DM*(Θ*\**) is re-calculated by equation 4.1.5.

|  |  |
| --- | --- |
|  | (4.1.5) |

The convergence rate *m\** of GEM is re-defined by equation 4.1.6.

|  |  |
| --- | --- |
|  | (4.1.6) |

Because all *hi\**, *qi\**, and *πi\** are non-positive, we have:

Therefore, by comparing equation 4.1.6 and equation 3.26, we conclude that *m\** is smaller with regard to *Q*+(Θ’ | Θ). In other words, the convergence rate is improved with support of prior probability *π*(Θ). In literature of EM, the combination of GEM and MAP with support of *π*(Θ) results out a so-called MAP-GEM algorithm.

## **4.2. EM with Newton-Raphson method**

In the M-step of GEM algorithm, the next estimate Θ(*t*+1) is a maximizer of *Q*(Θ | Θ(*t*)), which means that Θ(*t*+1) is a solution of equation *D*10*Q*(Θ | Θ(*t*)) = **0***T* where *D*10*Q*(Θ | Θ(*t*)) is the first-order derivative of *Q*(Θ | Θ(*t*)) with regard to variable Θ. Newton-Raphson method (McLachlan & Krishnan, 1997, p. 29) is applied into solving the equation *D*10*Q*(Θ | Θ(*t*)) = **0***T*. As a result, M-step is replaced a so-called Newton step (N-step).

N-step starts with an arbitrary value Θ0 as a solution candidate and also goes through many iterations. Suppose the current parameter is Θ*i*, the next value Θ*i* +1 is calculated based on equation 4.2.1.

|  |  |
| --- | --- |
|  | (4.2.1) |

N-step converges after some *i*th iteration. At that time, Θ*i*+1 is solution of equation *D*10*Q*(Θ | Θ(*t*)) = 0 if Θ*i*+1=Θ*i*. So the next parameter of GEM is Θ(*t*+1) = Θ*i*+1. The equation 4.2.1 is Newton-Raphson process. Recall that *D*10*Q*(Θ | Θ(*t*)) is gradient vector and *D*20*Q*(Θ | Θ(*t*)) is Hessian matrix. Following is a proof of equation 4.2.1.

According to first-order Taylor series expansion of *D*10*Q*(Θ | Θ(*t*)) at Θ = Θ*i* with very small residual, we have:

Because *Q*(Θ | Θ(*t*)) is smooth enough, *D*20*Q*(Θ | Θ(*t*)) is symmetric matrix according to Schwarz’s theorem (Wikipedia, Symmetry of second derivatives, 2018), which implies:

*D*20*Q*(Θ | Θ(*t*)) = (*D*20*Q*(Θ | Θ(*t*)))*T*

So we have:

Let Θ = Θ*i*+1 and we expect that *D*10*Q*(Θ*i*+1 | Θ(*t*)) = **0***T* so that Θ*i*+1 is a solution.

It implies:

This means:

Rai and Matthews (Rai & Matthews, 1993) proposed a so-called EM1 algorithm in which Newton-Raphson process is reduced into one iteration, as seen in table 4.2.1 (Rai & Matthews, 1993, pp. 587-588). Rai and Matthews assumed that *f*(*x*) belongs to exponential family but their EM1 algorithm is really a variant of GEM in general. In other words, there is no requirement of exponential family for EM1.

|  |  |  |
| --- | --- | --- |
| *E-step*:  The expectation *Q*(Θ | Θ(*t*)) is determined based on current Θ(*t*), according to equation 2.8. Actually, *Q*(Θ | Θ(*t*)) is formulated as function of Θ.  *M-step*:  The next parameter Θ(*t*+1) is:   |  |  | | --- | --- | |  | (4.2.2) | |

**Table 4.2.1.** E-step and M-step of EM1 algorithm

Rai and Matthews proved convergence of EM1 algorithm by their proposal of equation 4.2.2. Second-order Taylor series expending for *Q*(Θ | Θ(*t*)) at Θ = Θ(*t*+1) to obtain:

Where Θ0(*t*+1) is on the line segment joining Θ and Θ(*t*+1). Let Θ = Θ(*t*), we have:

By substituting equation 4.2.2 for *Q*(Θ(*t*+1) | Θ(*t*)) – *Q*(Θ(*t*) | Θ(*t*)) with note that *D*20*Q*(Θ | Θ(*t*)) is symmetric matrix, we have:

Let,

Because *Q*(Θ’ | Θ) is smooth enough, *D*20*Q*(Θ(*t*) | Θ(*t*)) and *D*20*Q*(Θ0(*t*+1) | Θ(*t*)) are symmetric matrices according to Schwarz’s theorem (Wikipedia, Symmetry of second derivatives, 2018). Thus, *D*20*Q*(Θ(*t*) | Θ(*t*)) and *D*20*Q*(Θ0(*t*+1) | Θ(*t*)) are commutative:

*D*20*Q*(Θ(*t*) | Θ(*t*))*D*20*Q*(Θ0(*t*+1) | Θ(*t*)) = *D*20*Q*(Θ0(*t*+1) | Θ(*t*))*D*20*Q*(Θ(*t*) | Θ(*t*))

Suppose both *D*20*Q*(Θ(*t*) | Θ(*t*)) and *D*20*Q*(Θ0(*t*+1) | Θ(*t*)) are diagonalizable then, they are simultaneously diagonalizable (Wikipedia, Commuting matrices, 2017). Hence there is an (orthogonal) eigenvector matrix *V* such that (Wikipedia, Diagonalizable matrix, 2017) (StackExchange, 2013):

Where *Qe*(*t*) and *Qe*(*t*+1) are eigenvalue matrices of *D*20*Q*(Θ(*t*) | Θ(*t*)) and *D*20*Q*(Θ0(*t*+1) | Θ(*t*)), respectively. Matrix *A* is decomposed as below:

(Because *Qe*(*t*) and *Qe*(*t*+1) are commutative)

Hence, eigenvalue matrix of *A* is also *Qe*(*t*+1). Suppose *D*20*Q*(Θ0(*t*+1) | Θ(*t*)) is negative definite, *A* is negative definite too. We have:

Because *D*20*Q*(Θ(*t*) | Θ(*t*)) is negative definite, we have:

Because *A* is negative definite, we have:

As a result, we have:

Hence, EM1 surely converges to a local maximizer Θ*\** according to corollary 3.3 with assumption that *D*20*Q*(Θ0(*t*+1) | Θ(*t*)) and *D*20*Q*(Θ(*t*) | Θ(*t*)) are negative definite for all *t* where Θ0(*t*+1) is a point on the line segment joining Θ and Θ(*t*+1).

Rai and Matthews made experiment on their EM1 algorithm (Rai & Matthews, 1993, p. 590). As a result, EM1 algorithm saved a lot of computations in M-step. In fact, by comparing GEM (table 2.3) and EM1 (table 4.2.1), we conclude that EM1 increases *Q*(Θ | Θ(*t*)) after each iteration whereas GEM maximizes *Q*(Θ | Θ(*t*)) after each iteration. However, EM1 will maximizes *Q*(Θ | Θ(*t*)) at the last iteration when it converges. EM1 gains this excellent and interesting result because of Newton-Raphson process specified by equation 4.2.2.

Because equation 3.17 is not changed with regard to EM1, the convergence matrix of EM1 is not changed.

Therefore, EM1 does not improve convergence rate in theory as MAP-GEM algorithm does but EM1 algorithm really speeds up GEM process in practice because it saves computational cost in M-step.

In equation 4.2.2, the second-order derivative *D*20*Q*(Θ(*t*) | Θ(*t*)) is re-computed at every iteration for each Θ(t). If *D*20*Q*(Θ(*t*) | Θ(*t*)) is complicated, it can be fixed by *D*20*Q*(Θ(1) | Θ(1)) over all iterations where Θ(1) is arbitrarily initialized for EM process so as to save computational cost. In other words, equation 4.2.2 is replaced by equation 4.2.3 (Ta, 2014).

|  |  |
| --- | --- |
|  | (4.2.3) |

In equation 4.2.3, only *D*10*Q*(Θ(*t*) | Θ(*t*)) is re-computed at every iteration whereas *D*20*Q*(Θ(1) | Θ(1)) is fixed. Equation 4.2.3 implies a pseudo Newton-Raphson process which still converges to a local maximizer Θ*\** but it is slower than Newton-Raphson process specified by equation 4.2.2 (Ta, 2014).

Newton-Raphson process specified by equation 4.2.2 has second-order convergence. I propose to use equation 4.2.4 for speeding up EM1 algorithm. In other words, equation 4.2.2 is replaced by equation 4.2.4 (Ta, 2014), in which Newton-Raphson process is improved with third-order convergence. Note, equation 4.2.4 is common in literature of Newton-Raphson process.

|  |  |
| --- | --- |
| Where, | (4.2.4) |

The convergence of equation 4.2.4 is same as the convergence of equation 4.2.2. Following is a proof of equation 4.2.4 by Ta (Ta, 2014).

Without loss of generality, suppose Θ is scalar such that Θ = *θ*, let

Let *r*(*θ*) represents improved Newton-Raphson process.

Suppose *ω*(*θ*) has first derivative and we will find *ω*(*θ*). According to Ta (Ta, 2014), the first-order derivative of *η*(*θ*) is:

According to Ta (Ta, 2014), the second-order derivative of *η*(*θ*) is:

If is solution of equation *q*(*θ*) = 0, Ta (Ta, 2014) gave:

In order to achieve , Ta (Ta, 2014) selected:

According to Ta (Ta, 2014), Newton-Raphson process is improved as follows:

This means:

As a result, equation 4.2.4 is a generality of the equation above when Θ is vector.

I propose to apply gradient descent method (Ta, 2014) into M-step of GEM so that Newton-Raphson process is replaced by gradient descent process with expectation that descending direction which is the opposite of gradient vector *D*10*Q*(Θ | Θ(*t*)) speeds up convergence of GEM. Table 4.2.2 specifies GEM associated with gradient descent method, which is called GD-GEM algorithm.

|  |  |  |  |  |
| --- | --- | --- | --- | --- |
| *E-step*:  The expectation *Q*(Θ | Θ(*t*)) is determined based on current Θ(*t*), according to equation 2.8. Actually, *Q*(Θ | Θ(*t*)) is formulated as function of Θ.  *M-step*:  The next parameter Θ(*t*+1) is:   |  |  | | --- | --- | |  | (4.2.5) |   Where *γ*(*t*) > 0 is length of the descending direction. As usual, *γ*(*t*) is selected such that   |  |  | | --- | --- | |  | (4.2.6) |   Where, |

**Table 4.2.1.** E-step and M-step of GD-GEM algorithm

Note, gradient descent method is used to solve minimization problem but its use for solving maximization problem is the same. Second-order Taylor series expending for *Q*(Θ | Θ(*t*)) at Θ = Θ(*t*+1) to obtain:

Where Θ0(*t*+1) is on the line segment joining Θ and Θ(*t*+1). Let Θ = Θ(*t*), we have:

By substituting equation 4.2.5 for *Q*(Θ(*t*+1) | Θ(*t*)) – *Q*(Θ(*t*+1) | Θ(*t*)), we have:

Due to:

As a result, we have:

Hence, GD-GEM surely converges to a local maximizer Θ*\** according to corollary 3.3 with assumption that *D*20*Q*(Θ0(*t*+1) | Θ(*t*)) is negative definite where Θ0(*t*+1) is a point on the line segment joining Θ and Θ(*t*+1).

It is not easy to solve the maximization problem with regard to *γ* according to equation 4.2.6. So if *Q*(Θ | Θ(*t*)) satisfies Wolfe conditions (Wikipedia, Wolfe conditions, 2017) and concavity and *D*10*Q*(Θ | Θ(*t*)) is Lipschitz continuous (Wikipedia, Lipschitz continuity, 2018) then, equation 4.2.6 is replaced by equation 4.2.7 (Wikipedia, Gradient descent, 2018).

|  |  |
| --- | --- |
|  | (4.2.7) |

Where |.| denotes length or module of vector.

## **4.3. EM with Aitken acceleration**

According to Lansky and Casella (Lansky & Casella, 1992), GEM converges faster by combination of GEM and Aitken acceleration. Without loss of generality, suppose Θ is scalar such that Θ = *θ*, the sequence is monotonous. From equation 3.23

We have the following approximate with *t* large enough (Lambers, 2009, p. 1):

We establish the following equation from the above approximation, as follows (Lambers, 2009, p. 1):

Hence, *θ\** is approximated by (Lambers, 2009, p. 1)

We construct Aitken sequence such that (Wikipedia, Aitken's delta-squared process, 2017)

|  |  |
| --- | --- |
|  | (4.3.1) |

Where Δ is forward difference operator,

And

When Θ is vector as Θ = (*θ*1, *θ*2,…, *θr*)*T*, Aitken sequence is defined by applying equation 4.3.1 into its components *θi* (s) according to equation 4.3.2:

|  |  |
| --- | --- |
|  | (4.3.2) |

Where,

According theorem of Aitken acceleration, Aitken sequence approaches Θ*\** faster than the sequence with note that the sequence is instance of GEM.

Essentially, the combination of GEM and Aitken acceleration is to replace the sequence by Aitken sequence as seen in table 4.3.1.

|  |
| --- |
| *E-step*:  The expectation *Q*(Θ | Θ(*t*)) is determined based on current Θ(*t*), according to equation 2.8. Actually, *Q*(Θ | Θ(*t*)) is formulated as function of Θ. Note that *t* = 1, 2, 3,… and Θ(0) = Θ(1).  *M-step*:  Let Θ(*t*+1) = (*θ*1(*t*+1), *θ*2(*t*),…, *θr*(*t*+1))*T* be a maximizer of *Q*(Θ | Θ(*t*)). Note Θ(*t*+1) will become current parameter at the next iteration ((*t*+1)th iteration).  Aitken parameter is calculated according to equation 4.3.2.  If then, the algorithm stops and we have . |

**Table 4.3.1.** E-step and M-step of GEM algorithm combined with Aitken acceleration

Because Aitken sequence converges to Θ*\** faster than the sequence does, the convergence of GEM is improved with support of Aitken acceleration method.

In equation 4.3.2, parametric components *θi* (s) converges separately. Guo, Li, and Xu (Guo, Li, & Xu, 2017) assumed such components converges together with the same rate. So they replaced equation 4.3.2 by equation 4.3.3 (Guo, Li, & Xu, 2017, p. 176) for Aitken sequence .

|  |  |
| --- | --- |
|  | (4.3.3) |

## **4.4. ECM algorithm**

Because M-step of GEM is complicated, Meng and Rubin (Meng & Rubin, 1993) proposed a so-called Expectation Conditional Expectation (ECM) algorithm in which M-step is replaced by several computationally simpler Conditional Maximization (CM) steps. Each CM-step maximizes *Q*(Θ | Θ(*t*)) on given constraint. ECM is very useful in the case that maximization of *Q*(Θ | Θ(*t*)) with constraints is simpler than maximization of *Q*(Θ | Θ(*t*)) without constraints as usual.

Suppose the parameter Θ is partitioned into *S* sub-parameters Θ = {Θ1, Θ2,…, Θ*S*} and there are *S* pre-selected vector function *gs*(Θ):

|  |  |
| --- | --- |
|  | (4.4.1) |

Each function *gs*(Θ) represents a constraint. Support there is a sufficient enough number of derivatives of each *gs*(Θ). In ECM algorithm (Meng & Rubin, 1993, p. 268), M-step is replaced by a sequence of CM-steps. Each CM-step maximizes *Q*(Θ | Θ(*t*)) over Θ but with some function *gs*(Θ) fixed at its previous value. Concretely, there are *S* CM-steps and every *s*th CM-step finds Θ(*t*+*s*/*S*) that maximizes *Q*(Θ | Θ(*t*)) over Θ subject to the constraint *gs*(Θ) = *gs*(Θ(*t*+(*s*–1)/*S*)). The next parameter Θ(*t*+1) is the output of the final CM-step such that Θ(*t*+1) = Θ(*t*+*s*/*S*). Table 4.4.1 (Meng & Rubin, 1993, p. 272) shows E-step and CM-steps of ECM algorithm.

|  |  |  |  |  |
| --- | --- | --- | --- | --- |
| *E-step*:  As usual, *Q*(Θ | Θ(*t*)) is determined based on current Θ(*t*) according to equation 2.8. Actually, *Q*(Θ | Θ(*t*)) is formulated as function of Θ.  *CM-steps*:  There are *S* CM-steps. In every *s*th CM step (*s* =1, 2,…, *S*), finding   |  |  | | --- | --- | |  | (4.4.2) |   The next parameter Θ(*t*+1) is the output of the final CM-step (*S*th CM-step):   |  |  | | --- | --- | |  | (4.4.3) |   Note, Θ(*t*+1) will become current parameter at the next iteration ((*t*+1)th iteration). |

**Table 4.3.1.** E-step and CM-steps of ECM algorithm

ECM algorithm stops at some *t*th iteration such that Θ(*t*) = Θ(*t*+1) = Θ\*. CM-steps depend on how to define pre-selected functions in *G*. For example, if *gs*(Θ) consists all sub-parameters except Θ*s* then, the *s*th CM-step maximizes *Q*(Θ | Θ(*t*)) with regard to Θ*s* whereas other sub-parameters are fixed. If *gs*(Θ) consists only Θ*s* then, the *s*th CM-step maximizes *Q*(Θ | Θ(*t*)) with regard to all sub-parameters except Θ*s*. Note, definition of ECM algorithm is specified by equation 4.4.2 and equation 4.4.3

From equation 4.4.2 and equation 4.4.3, we have:

Hence, the convergence of ECM is asserted according to corollary 3.3. However, Meng and Rubin (Meng & Rubin, 1993, pp. 274-276) provided some conditions for convergence of ECM to a maximizer of *L*(Θ).

# **5. Applications of EM**

## **5.1. Mixture model**

As usual, let ***X*** be the hidden or latent space and let ***Y*** be the observed space. Especially, the random variable *X* in ***X*** represents latent class or latent component of random variable *Y* in ***Y***. Suppose *X* is discrete and ranges in ***X*** = {1, 2,…, *K*}. The so-called probabilisticfinite *mixture model* is represented by the PDF of *Y*, as seen in equation 5.1.1.

|  |  |
| --- | --- |
|  | (5.1.1) |

Where,

Note, *Y* can be discrete or continuous. Recall that the ultimate purpose of EM algorithm is to maximize *f*(*Y*|Θ) with subject to Θ. Each *fX*(*Y*|*θX*) is called the *X*th partial PDF of *Y* whose partial parameter is *θX*. Each *fX*(*Y*|*θX*) is also called the *X*th observational PDF of *Y*. It is really the conditional PDF of *Y* given *X*, as seen in equation 5.1.2.

|  |  |
| --- | --- |
|  | (5.1.2) |

From equation 5.1.1, the mixture model *f*(*Y*|Θ) is the mean of *K* partial PDFs. The variable *X* implies which partial PDF “generates” *Y* (Bilmes, 1998, p. 5).

Each *αX* is called mixture coefficient. It is really the probability of discrete *X*, as seen in equation 5.1.3. However, in mixture model, each *αX* is also considered as parameter, which is belongs to the compound parameter Θ.

|  |  |
| --- | --- |
|  | (5.1.3) |

Note,

The joint probabilistic distribution of *X* and *Y*, which implies the implicit mapping between ***X*** and ***Y***, is product of the mixture coefficient *αX* and the *X*th PDF of *Y*, as seen in equation 5.1.4.

|  |  |
| --- | --- |
|  | (5.1.4) |

This implies:

|  |  |
| --- | --- |
|  | (5.1.5) |

Equation 5.1.6 specifies the conditional probability of *X* given *Y*. Please pay attention to this important probability.

|  |  |
| --- | --- |
|  | (5.1.6) |

Following is the proof of equation 5.1.6. According to Bayes’ rule, we have:

Applying equation 5.1.3 and equation 5.1.4, we have:

In other words, equation 5.1.6 is established■

Now GEM algorithm is applied into mixture model for estimating the parameter Θ. Derived from equation 2.12, the conditional expectation *Q*(Θ’|Θ) of mixture model becomes:

|  |  |
| --- | --- |
|  | (5.1.7) |

In practice, suppose *Y* is observed as a sample = {*Y*1, *Y*2,…, *YN*} of size *N* in which all *Yi* (s) are mutually independent and identically distributed (iid). The observed sample is associated with a a hidden set (latent set) = {*X*1, *X*2,…, *XN*} of size *N*. All *Xi* (s) are iid and they are not existent in fact. Let be the random variable representing every *Xi*. Of course, the domain of *X* is ***X***. Derived from equation 2.15, equation 5.1.8 specifies *Q*(Θ’|Θ) given such .

|  |  |
| --- | --- |
|  | (5.1.8) |

Equation 5.1.8 is the general case of equation 5.1.7. At the *t*th iteration of GEM, given current parameter Θ(*t*) = (*α*1(*t*), *α*2(*t*),…, *αK*(*t*), *θ*1(*t*), *θ*2(*t*),…, *θK*(*t*))*T*, the conditional expectation specified by equation 5.1.8 is written as follows:

Thus, the unknown of *Q*(Θ|Θ(*t*)) is Θ = (*α*1, *α*2,…, *αK*, *θ*1, *θ*2,…, *θK*)*T*. Because *X* is discrete and ranges in {1, 2,…, *K*}, the conditional expectation *Q*(Θ|Θ(*t*)) is re-written as equation 5.1.9 for convenience.

|  |  |
| --- | --- |
|  | (5.1.9) |

Where the conditional probability *P*(*k* | *Y*, Θ(*t*)) is determined by equation 5.1.10 which is indeed equation 5.1.6.

|  |  |
| --- | --- |
|  | (5.1.10) |

At M-step of the current *t*th iteration, *Q*(Θ|Θ(*t*)) specified by equation 5.1.9 is maximized with subject to Θ. How to maximize *Q*(Θ|Θ(*t*)) with subject to Θ is dependent on types of partial PDFs *fk*(*Yi*|*θk*).

Because there is the constraint , we use Lagrange duality method to maximize to maximize *Q*(Θ|Θ(*t*)). The Lagrange function *la*(Θ, *λ* | Θ(*t*)) is sum of *Q*(Θ|Θ(*t*)) and the constraint , which is specified by equation 5.1.11.

|  |  |
| --- | --- |
|  | (5.1.11) |

Note, *λ* ≥ 0 is called Lagrange multiplier. Of course, *la*(Θ, *λ* | Θ(*t*)) is function of Θ and *λ*. The next parameters *αk*(*t*+1) that maximizes *Q*(Θ|Θ(*t*)) is solution of the equation formed by setting the first-order partial derivative of Lagrange function regarding *αk* and *λ* to be zero with suppose that the Lagrange function is first-order smooth function.

This implies:

|  |  |
| --- | --- |
|  | (5.1.12) |

Summing equation 5.1.12 over *K* classes {1, 2,…, *K*}, we have (Bilmes, 1998, p. 5):

Substituting *λ = N* into equation 5.1.12, the next parameters *αk*(*t*+1) is totally determined by equation 5.1.13.

|  |  |
| --- | --- |
|  | (5.1.13) |

Note, the conditional probability *P*(*k* | *Yi*, Θ(*t*)) is determined by equation 5.1.10.

When parameters *αk*(*t*+1) and *λ* are determined, the Lagrange function *la*(Θ, *λ* | Θ(*t*)) is now function of parameters *θk* as *la*(*θk*|*θk*(*t*)). The next parameters *θk*(*t*+1) is solution of the equation formed by setting the first-order partial derivative of Lagrange function regarding *θk* to be zero with suppose that the Lagrange function is first-order smooth function.

Thus, the next parameters *θk*(*t*+1) is solution of the equation 5.1.14.

|  |  |
| --- | --- |
|  | (5.1.14) |

The two steps of GEM algorithm for constructing mixture model at some *t*th iteration are shown in table 5.1.1. Note, suppose the Lagrange function is first-order smooth function.

|  |
| --- |
| *E-step*:  The conditional probability *P*(*k* | *Yi*, Θ(*t*)) is calculated based on current parameter Θ(*t*) = (*α*1(*t*), *α*2(*t*),…, *αK*(*t*), *θ*1(*t*), *θ*2(*t*),…, *θK*(*t*))*T*, according to equation 5.1.10.  *M-step*:  The next parameter Θ(*t*+1) = (*α*1(*t*+1), *α*2(*t*+1),…, *αK*(*t*+1), *θ*1(*t*+1), *θ*2(*t*+1),…, *θK*(*t*+1))*T*, which is a maximizer of *Q*(Θ | Θ(*t*)) with subject to Θ, is calculated by equation 5.1.13 and equation 5.1.14. Note, *θk*(*t*+1) is solution of the equation 5.1.14. |

**Table 5.1.1.** E-step and M-step of GEM algorithm for constructing mixture model regarding first-order smooth Lagrange function

GEM algorithm converges at some *t*th iteration. At that time, Θ*\** = Θ(*t*+1) = Θ(*t*) is the optimal estimate of mixture model regarding first-order smooth Lagrange function.

Suppose that each PDF *fk*(*Yi*|*θk*) ) belongs to regular exponential family and then, solving equation 5.1.4 is easier as follows:

(Due to *fk*(*Yi*|*θk*) ) belongs to exponential family)

(Due to log’(*a*(*θk*)) = (*E*(*τ*(*Y*|*θk*)))*T*, please see table 1.2)

In general, the next parameters *θk*(*t*+1) is solution of the equation 5.1.15 within regular exponential family.

|  |  |
| --- | --- |
|  | (5.1.15) |

Where *Y* is the random variable representing all *Yi* (s) and,

The two steps of GEM algorithm for constructing mixture model at some *t*th iteration are shown in table 5.1.2 with suppose that each partial PDF *fX*(*Y*|*θX*) is assumed to belong regular exponential family.

|  |
| --- |
| *E-step*:  The conditional probability *P*(*k* | *Yi*, Θ(*t*)) is calculated based on current parameter Θ(*t*) = (*α*1(*t*), *α*2(*t*),…, *αK*(*t*), *θ*1(*t*), *θ*2(*t*),…, *θK*(*t*))*T*, according to equation 5.1.10.  *M-step*:  The next parameter Θ(*t*+1) = (*α*1(*t*+1), *α*2(*t*+1),…, *αK*(*t*+1), *θ*1(*t*+1), *θ*2(*t*+1),…, *θK*(*t*+1))*T*, which is a maximizer of *Q*(Θ | Θ(*t*)) with subject to Θ, is calculated by equation 5.1.13 and equation 5.1.15. Note, *θk*(*t*+1) is solution of the equation 5.1.15. |

**Table 5.1.2.** E-step and M-step of GEM algorithm for constructing mixture model regarding regular exponential family

GEM algorithm converges at some *t*th iteration. At that time, Θ*\** = Θ(*t*+1) = Θ(*t*) is the optimal estimate of mixture model regarding regular exponential family.

There is a special case that each *fk*(*Yi*|*θk*) is normal distribution, which is popular in domain of mixture model, with note that normal distribution belongs to regular exponential family. Thus, let *Y* be random variable representing all *Yi*. Without loss of generality, suppose *Y* is vector so that each *fk*(*Y*|*θk*) is multinormal distribution. Recall that each *fk*(*Y*|*θk*) is called the *k*th partial PDF of *Y* or the *k*th observational PDF of *Y*. In this case, the mixture model is called *normal mixture model* (Gaussian mixture model) and it is easy to solve equation 5.1.14 or equation 5.1.15 for *θk*. Suppose random variable *Y* is vector of size *n*.

|  |  |
| --- | --- |
|  | (5.1.16) |

Where *μk* and Σ*k* are mean vector and covariance matrix of *fk*(*Y*|*θk*), respectively. The notation |.| denotes determinant of given matrix and the notation Σ*k*–1 denotes inverse of matrix Σ*k*. Note, Σ*k* is invertible and symmetric. Now we find other parameters *θk*(*t*+1) = (*μk*(*t*+1), Σ*k*(*t*+1))*T* by solving directly equation 5.1.14 or equation 5.1.15. Recall that each *Yi* conforms to multinormal distribution, according to equation 5.1.16.

Where *μk* and Σ*k* are mean and covariance matrix of *fk*(*Yi*|*θk*), respectively. The Lagrange function is re-written as follows:

Where *p* is the dimension of *Yi*; in other words, *p* is the dimension of space ***Y***.

The first-order partial derivative of Lagrange function with respect to *μk* is (Nguyen, 2015, p. 35):

The next parameter *μk*(*t*+1) that maximizes *Q*(Θ|Θ(*t*)) is solution of the equation formed by setting the first-order partial derivative of Lagrange function with regard to *μk* to be **0***T*. Note that **0** = (0, 0,…, 0)*T* is zero vector.

This implies equation 5.1.17 to specify the next parameter *μk*(*t*+1).

|  |  |
| --- | --- |
|  | (5.1.17) |

Note, the conditional probability *P*(*k* | *Yi*, Θ(*t*)) is determined by equation 5.1.10.

The first-order partial derivative of Lagrange function with respect to Σ*k* is:

Due to:

And

Because Bilmes (Bilmes, 1998, p. 5) mentioned:

Where tr(*A*) is trace operator which takes sum of diagonal elements of matrix .

This implies (Nguyen, 2015, p. 45):

Where Σ*k* is symmetric and invertible matrix. Substituting the next parameter *μk*(*t*+1) specified by equation 5.1.16 into the first-order partial derivative of Lagrange function with respect to Σ*k*, we have:

The next parameter Σ*k*(*t*+1) that maximizes *Q*(Θ|Θ(*t*)) is the solution of equation formed by setting the first-order partial derivative of Lagrange function regarding Σ*k* to zero matrix. Let (**0**) denote zero matrix.

We have:

This implies equation 5.1.18 to specify the next parameter Σ*k*(*t*+1).

|  |  |
| --- | --- |
|  | (5.1.18) |

Note, the conditional probability *P*(*k* | *Yi*, Θ(*t*)) is determined by equation 5.1.10 and the next parameter *μk*(*t*+1) is specified by equation 5.1.17.

As a result, the solution *θk*(*t*+1) = (*μk*(*t*+1), Σ*k*(*t*+1))*T* of equation 5.1.14 or equation 5.1.15 is specified by equation 5.1.17 and equation 5.1.18 when each *fk*(*Y*|*θk*) is multinormal distribution within normal mixture model. The two steps of GEM algorithm for constructing normal mixture model at some *t*th iteration are refined in table 5.1.3 (Bilmes, 1998, p. 7).

|  |
| --- |
| *E-step*:  The conditional probability *P*(*k* | *Yi*, Θ(*t*)) is calculated based on current parameter Θ(*t*) = (*α*1(*t*), *α*2(*t*),…, *αK*(*t*), *θ*1(*t*), *θ*2(*t*),…, *θK*(*t*))*T*, according to equation 5.1.10. Note, in normal mixture model, each observational PDF *fk*(*Y*|*θk*) is (multivariate) normal distribution with mean vector *μk* and covariance matrix Σ*k* such that *θk* = (*μk*, Σ*k*)*T*.  *M-step*:  The next parameter Θ(*t*+1) = (*α*1(*t*+1), *α*2(*t*+1),…, *αK*(*t*+1), *θ*1(*t*+1), *θ*2(*t*+1),…, *θK*(*t*+1))*T*, which is a maximizer of *Q*(Θ | Θ(*t*)) with subject to Θ, is calculated by equation 5.1.13, equation 5.1.17, and equation 5.1.18 with current parameter Θ(*t*). |

**Table 5.1.3.** E-step and M-step of GEM algorithm for constructing normal mixture model

GEM algorithm converges at some *t*th iteration. At that time, Θ*\** = Θ(*t*+1) = Θ(*t*) is the optimal estimate of normal mixture model.

An interesting application of finite mixture model is soft clustering. Traditional clustering methods assign a fixed cluster to every data point in sample, which means that every data point belongs exactly to one cluster. There are some popular (hard) clustering methods such as *K*-means and *K*-medoids (Han & Kamber, 2006, pp. 451-457). Soft clustering is more flexible when every data point belongs to more than one cluster and the degree of assignment is represented by a probability. Concretely, GEM algorithm for normal mixture model described in table 5.1.3 is applied into soft clustering. Given sample = {*Y*1, *Y*2,…, *YN*} of size *N* in which all *Yi* (s) are iid and each *Yi* is also called a data point, soft clustering partitions into *K* clusters and each cluster *k* is considered as hidden variable (*X* = 1, 2,…, *K*) and is represented by the aforementioned normal PDF *fk*(*Y*|*θk*)

Where *θk* = (*μk*, Σ*k*)*T* includes mean vector *μk* and covariance matrix Σ*k* of *fk*(*Y*|*θk*), respectively. Especially, *μk* is considered as centroid of cluster *k*. Given cluster *k*, the degree of assignment that a data point *Y* belonging to cluster *k* is specified by such *fk*(*Y*|*θk*). Therefore, GEM algorithm for normal mixture model is used to learn Θ = (*α*1, *α*2,…, *αK*, *θ*1, *θ*2,…, *θK*)*T*. The parameter *αk* indicates degree of popularity of cluster *k*, which can be considered as capacity or size of cluster *k*. It can be also considered as coverage ratio of cluster *k*. The higher the *αk* is, the larger the cluster *k* is. Essentially, soft clustering is to estimate *αk* and *θk* by GEM. Suppose after GEM results out the best estimate Θ\* = (*α*1\*, *α*2\*,…, *αK*\*, *θ*1\*, *θ*2\*,…, *θK*\*)*T*, it is required to determine to which cluster a new data point *Y* is more likely to belong. We calculate *K* joint probabilities *p*1 = *α*1\**f*1(*Y*|*θ*1\*), *p*2 = *α*2\**f*2(*Y*|*θ*2\*),…, and *pK* = *αK*\**fK*(*Y*|*θK*\*). Indeed, each *pk* is the joint probability of *Y* and cluster *k* that come together. Suppose some *pj* is maximum then, *Y* is more likely to belong cluster *j*.

Of course, the probability of each data point *Y* within soft clustering for *K* clusters is

But this probability *f*(*Y*|Θ) is not important. The most important task of GEM for soft clustering is to compute the estimate Θ\* = (*α*1\*, *α*2\*,…, *αK*\*, *θ*1\*, *θ*2\*,…, *θK*\*)*T* from sample in order to determine clusters because each cluster *k* is represented by a pair {*αk*\*, *θk*\*}.

**Example 5.1.1.** Given sample = {*Y*1, *Y*2, *Y*3, *Y*4}, we apply GEM for soft clustering into *K*=2 clusters.

|  |  |  |
| --- | --- | --- |
|  | *y*1 | *y*2 |
| *Y*1 | 0 | 0 |
| *Y*2 | 0 | 1 |
| *Y*3 | 2 | 0 |
| *Y*4 | 2 | 1 |

Of course, we have *Y*1 = (*y*11=0, *y*12=0)*T*, *Y*2 = (*y*21=0, *y*22=1)*T*, *Y*3 = (*y*31=2, *y*32=0)*T*, and *Y*4 = (*y*41=2, *y*42=1)*T*. The parameter Θ = (*α*1, *α*2, *θ*1, *θ*2)*T* is initialized as follows:

Note, it is easy to calculate normal PDF *fk*(*Y*|*θk*) with known *θk* = (*μk*, Σ*k*)*T*.

At the 1st iteration, E-step we have:

At the 1st iteration, M-step we have:

At the 2nd iteration, E-step we have:

At the 2nd iteration, M-step we have:

Therefore, GEM stops at the 2nd iteration with the estimate Θ(2) = Θ(3) = Θ\* = (*α*1\*, *α*2\*, *θ*1\*, *θ*2\*)*T*.

Given new data point *Y* = (0.5, 0.5)*T*, it is required to determine to which cluster *Y* is more likely to belong. We calculate *K* joint probabilities as follows:

Due to some *p*1=*p*2, the likelihood that *Y* belongs to such two clusters is equal ■

Every observation in ordinary sample is univariate or multivariate but there is a case that ordinary sample becomes dyadic sample related to two sets of objects, which causes some modifications of mixture model. *Dyadic data* which is also called co-occurrence data (COD) contains co-occurrent events of objects. It is necessary to obtain statistical models to represent dyadic data and fortunately, finite mixture model is the one. Recall that EM is applied to learn mixture model. Here we focus on EM and mixture model for dyadic data or COD.

Given two finite sets = {*x*1, *x*2,…, *xN*) and = {*y*1, *y*2,…, *yM*) with note that *xi* (s) and *yj* (s) represent -objects and -objects, respectively; exactly, they are names of objects. The numbers of -objects and -objects are =*N* and =*M*, respectively. For example, in information retrieval, *xi* (s) are documents and *yj* (s) are keywords. Hence, *xi* and *yj* are not evaluated as numbers. An observational pair (*xi*, *yj*) is called a *co-occurrence* of *xi* and *yj*. Dyadic data or COD contains these co-occurrences with note that a co-occurrence (*xi*, *yj*) can exist more than one time. So, each co-occurrence (*xi*, *yj*) is indexed by an index *r*. As a result, each co-occurrence is denoted by the triple (*xi*, *yj*, *r*) and we have (Hofmann & Puzicha, 1998, p. 1):

|  |  |
| --- | --- |
|  | (5.1.19) |

Where,

Of course, the size of is . As a convention, *xi*(*r*) and *yj*(*r*) indicate that -object and -object at the *r*th co-occurrence are *xi* and *yj*, respectively. Thus, the triplet (*xi*, *yj*, *r*) can be denoted as (*xi*(*r*), *yj*(*r*), *r*). For example, suppose = {*x*1, *x*2, *x*3) and = {*y*1, *y*2), and dyadic data of 4 co-occurrences, = {(*x*1, *y*1, 1), (*x*1, *y*1, 2), (*x*1, *y*2, 3), (*x*1, *y*1, 4)}, we observe that *x*1 and *y*1 occur together three times at *r*=1, *r*=2, and *r*=4 where as *x*1 and *y*2 occur together one time at *r*=3. In the first co-occurrence (*x*1, *y*1, 1), the notation *x*1(1) indicate that the -object at this co-occurrence is *x*1. In the third co-occurrence (*x*1, *y*2, 3), the notation *y*2(3) indicate that the -object at this co-occurrence is *y*2.

If each co-occurrence of *xi* and *yj* is associated with a value *z* (Hofmann, Puzicha, & Jordan, Learning from Dyadic Data, 1998, p. 1), the triple (*xi*, *yj*, *r*) becomes the quadruplet (*xi*, *yj*, *z*, *r*) which is called *valued co-occurrence* of *xi* and *yj*. The value *z* is called associative value or co-occurrent value. If *z* is value of a variable *Z* then, *Z* is called associative variable or co-occurrent variable. As a result, the sample is called *valued dyadic data* or valued COD. Note, *Z* can be univariate or multivariate (vector).

|  |  |
| --- | --- |
|  | (5.1.20) |

Where,

As a convention, *Z*(*r*) or *z*(*r*) indicates that the associative value at *r*th co-occurrence is *Z*=*z*. Thus, the quadruplet (*xi*, *yj*, *Z*, *r*) can be denoted as (*xi*(*r*), *yj*(*r*), *Z*(*r*), *r*). For example, suppose = {*x*1, *x*2, *x*3) and = {*y*1, *y*2), and valued dyadic sample of 4 co-occurrences, = {(*x*1, *y*1, 6, 1), (*x*1, *y*1, 8, 2), (*x*1, *y*2, 7, 3), (*x*1, *y*1, 9, 4)}, we observe that *x*1 and *y*1 occur together three times at *r*=1, *r*=2, and *r*=4 where as *x*1 and *y*2 occur together one time at *r*=3. Moreover, at *r*=1, *r*=2, *r*=3, and *r*=4, associative values are *Z*(1)=6, *Z*(2)=7, *Z*(3)=8, and *Z*(4)=9, respectively. Valued dyadic data is special case of dyadic data. As a convention, dyadic data is default if there is no additional information.

Given fixed *xk*, let be the -partitioned subset of which contains co-occurrences whose -objects are fixed at *xk* (Hofmann & Puzicha, Statistical Models for Co-occurrence Data, 1998, p. 1). Note, can be empty. The size of is .

|  |  |
| --- | --- |
|  | (5.1.21) |

Dyadic data is partitioned into subsets .

Given fixed *yl*, let be the -partitioned subset of which contains co-occurrences whose -objects are fixed at *yl*. Note, can be empty. The size of is .

|  |  |
| --- | --- |
|  | (5.1.22) |

Dyadic data is partitioned into subsets .

Given fixed *xk* and fixed *yl*, let be the subset of the which contains co-occurrences whose -objects and -objects are fixed at *xk* and *yl*. Note, can be empty. The size of is .

|  |  |
| --- | --- |
|  | (5.1.23) |

Let *n*(*xi*) and *n*(*yj*) denote the number of *xi* and the number of *yj*, respectively.

|  |  |
| --- | --- |
|  | (5.1.24) |

Let *n*(*xi*, *yj*) denote the number of *xi* and *yj*.

|  |  |
| --- | --- |
|  | (5.1.25) |

Let *n*(*xi*|*yj*) and *n*(*yj*|*xi*) denote the frequency of *xi* given *yj* and the frequency of *yj* given *xi*, respectively.

|  |  |
| --- | --- |
|  | (5.1.26) |

For example, suppose = {*x*1, *x*2, *x*3) and = {*y*1, *y*2), and dyadic data of 4 co-occurrences, = {(*x*1, *y*1, 1), (*x*1, *y*1, 2), (*x*1, *y*2, 3), (*x*1, *y*1, 4)}, we have = {(*x*1, *y*1, 1), (*x*1, *y*1, 2), (*x*1, *y*2, 3), (*x*1, *y*1, 4)}, = = Ø, = {(*x*1, *y*1, 1), (*x*1, *y*1, 2), (*x*1, *y*1, 4)}, = {(*x*1, *y*2, 3)}, = = {(*x*1, *y*1, 1), (*x*1, *y*1, 2), (*x*1, *y*1, 4)}, = {(*x*1, *y*2, 3)}, = = = = Ø, *n*(*x*1) = 1, *n*(*x*2) = *n*(*x*3) = 0, *n*(*y*1) = 3, *n*(*y*2) = 1, *n*(*x*1, *y*1) = 3, *n*(*x*1, *y*2) = 1, *n*(*x*2, *y*1) = *n*(*x*2, *y*2) = *n*(*x*3, *y*1) = *n*(*x*3, *y*2) = 0, *n*(*x*1 | *y*1) = 1, *n*(*x*1 | *y*2) = 1, *n*(*x*2 | *y*1) = *n*(*x*2 | *y*2) = *n*(*x*3 | *y*1) = *n*(*x*3 | *y*2) = 0, *n*(*y*1 | *x*1) = 3/4, *n*(*y*2 | *x*1) = 1/4.

Suppose each co-occurrence (*xi*, *yj*) belongs to a latent variable *C* and *C* has *K* values *ck* (s). These values *ck* (s) are called classes or aspects and thus, mixture model for dyadic data is also called aspect model or latent class model which aims to discover the latent variable *C*. Without loss of generality, let *ck* = *k* where *k* = 1, 2,…, *K*. The random variable *C* has discrete distribution such that every value has an associated probability *αk*. Of course, there are *K* probabilities *αk* (s). There are three kinds of dyadic mixture model for dyadic data such as symmetric mixture model (SMM), asymmetric mixture model (AMM), and product-space mixture model (PMM). These models were introduced by Hofmann and Puzicha (Hofmann & Puzicha, Statistical Models for Co-occurrence Data, 1998).

The mixture model of dyadic data is called symmetric mixture model (SMM) if *αk* (s) are independent from both *xi* and *yj*. SMM is defined as follows (Hofmann & Puzicha, Statistical Models for Co-occurrence Data, 1998, p. 2):

|  |  |
| --- | --- |
|  | (5.1.27) |

Where *αk* is the probability of aspect *k*. Note, *P*(.) denote probability.

The is the probability of *xi* given aspect *k*.

The is the probability of *yj* given aspect *k*.

This implies that *xi* and *yj* are mutually independent in SMM.

The joint probability of *xi*, *yj*, and *k* is:

The parameter of SMM is Θ = (*αk*, *pi*|*k*, *qj*|*k*)*T* in which there are *K*( + + 1) partial parameters *αk*, *pi*|*k*, and *qj*|*k*. Note,

By applying GEM, given dyadic sample , at the *t*th iteration of GEM, given current parameter Θ(*t*) = (*αk*(*t*), *pi*|*k*(*t*), *qj*|*k*(*t*))*T*, the conditional expectation *Q*(Θ|Θ(*t*)) is (Nguyen, Learning Dyadic Data and Predicting Unaccomplished Co-occurrent Values by Mixture Model, 2020, p. 5):

|  |  |
| --- | --- |
|  | (5.1.28) |

Where,

|  |  |
| --- | --- |
|  | (5.1.29) |

Note, *n*(*xi*, *yj*) is the number of co-occurrences (*xi*, *yj*) in , which is specified by equation 5.1.25. Please refer to equation 5.1.6 and equation 5.1.10 to comprehend equation 5.1.29. Because there are three constraints

We use Lagrange duality method to maximize to maximize *Q*(Θ|Θ(*t*)). The Lagrange function *la*(Θ, *λ* | Θ(*t*)) is sum of *Q*(Θ|Θ(*t*)) and these constraints, as follows (Nguyen, Learning Dyadic Data and Predicting Unaccomplished Co-occurrent Values by Mixture Model, 2020, p. 5):

Note, *λ* = (*λ*1, *λ*2, *λ*3)*T* where *λ*1≥0, *λ*2≥0, and *λ*3≥0 are called Lagrange multipliers. Of course, *la*(Θ, *λ* | Θ(*t*)) is function of Θ and *λ*. The next parameters Θ(*t*+1) that maximizes *Q*(Θ|Θ(*t*)) at M-step of some *t*th iteration is solution of the equation formed by setting the first-order partial derivatives of Lagrange function regarding Θ and *λ* to be zero.

The first-order partial derivative of Lagrange function regarding *αk* is:

Setting this partial derivative to be zero, we obtain:

Summing the equation above over *K* aspects {1, 2,…, *K*}, we have:

This means the next parameters *αk*(*t*+1) is:

|  |  |
| --- | --- |
|  | (5.1.30) |

The first-order partial derivative of Lagrange function regarding *pi*|*k* is:

Setting this partial derivative to be zero, we obtain:

Summing the equation above over , we have:

This means the next parameters *pi*|*k*(*t*+1) is:

|  |  |
| --- | --- |
|  | (5.1.31) |

Similarly, the next parameters *qj*|*k*(*t*+1) is:

|  |  |
| --- | --- |
|  | (5.1.32) |

The two steps of GEM algorithm for SMM at some *t*th iteration are shown in table 5.1.4.

|  |
| --- |
| *E-step*:  The conditional probability *P*(*k* | *xi*, *yj*, Θ(*t*)) is calculated based on current parameter Θ(*t*) = (*αk*(*t*), *pi*|*k*(*t*), *qj*|*k*(*t*))*T*, according to equation 5.1.29.  *M-step*:  The next parameter Θ(*t*+1) = (*αk*(*t*+1), *pi*|*k*(*t*+1), *qj*|*k*(*t*+1))*T*, which is a maximizer of *Q*(Θ | Θ(*t*)) with subject to Θ, is calculated by equation 5.1.30, equation 5.1.31, and equation 5.1.32. |

**Table 5.1.4.** E-step and M-step of GEM algorithm for SMM

GEM algorithm converges at some *t*th iteration. At that time, Θ*\** = Θ(*t*+1) = Θ(*t*) is the SMM itself. When SMM is applied into soft clustering, dyadic data is clustered according to blocks and each *αk* is coverage ratio of cluster *k* (aspect *k*).

The mixture model of dyadic data is called asymmetric mixture model (AMM) if *αk* (s) are only independent from *xi* or from *yj*. Without loss of generality, given *αk* (s) are only independent from *yj* (of course, it is dependent on *xi*), AMM is defined as follows (Hofmann & Puzicha, Statistical Models for Co-occurrence Data, 1998, p. 3):

|  |  |
| --- | --- |
|  | (5.1.33) |

The *αk*|*i* is the probability of aspect *k* given *xi*.

Where *pi* is the probability of *xi*.

The *qj*|*k* is the conditional probability of *yj* given aspect *k*. Suppose *yj* is dependent from *xi* given *k*, we have:

Note, *qj*|*i* is the conditional probability of *yj* given *xi*, which is defined as follows:

The joint probability of *xi*, *yj*, and *k* is:

The parameter of AMM is Θ = (*αk*|*i*, *pi*, *qj*|*k*)*T* in which there are *K*( + ) + partial parameters *αk*|*i*, *pi*, and *qj*|*k*. Note,

By applying GEM, given dyadic sample , at the *t*th iteration of GEM, given current parameter Θ(*t*) = (*αk*(*t*), *pi*|*k*(*t*), *qj*|*k*(*t*))*T*, the conditional expectation *Q*(Θ|Θ(*t*)) is:

|  |  |
| --- | --- |
|  | (5.1.34) |

Where,

|  |  |
| --- | --- |
|  | (5.1.35) |

Please refer to equation 5.1.6 and equation 5.1.10 to comprehend equation 5.1.35. Because there are three constraints

We use Lagrange duality method to maximize to maximize *Q*(Θ|Θ(*t*)). The Lagrange function *la*(Θ, *λ* | Θ(*t*)) is sum of *Q*(Θ|Θ(*t*)) and these constraints, as follows:

Note, *λ* = (*λ*1, *λ*2, *λ*3)*T* where *λ*1≥0, *λ*2≥0, and *λ*3≥0 are called Lagrange multipliers. Of course, *la*(Θ, *λ* | Θ(*t*)) is function of Θ and *λ*. The next parameters Θ(*t*+1) that maximizes *Q*(Θ|Θ(*t*)) at M-step of some *t*th iteration is solution of the equation formed by setting the first-order partial derivatives of Lagrange function regarding Θ and *λ* to be zero.

The first-order partial derivative of Lagrange function regarding *αk*|*i* is:

Setting this partial derivative to be zero, we obtain:

Summing the equation above over *K* aspects {1, 2,…, *K*}, we have:

This means the next parameters *αk*|*i*(*t*+1) is:

|  |  |
| --- | --- |
|  | (5.1.36) |

The first-order partial derivative of Lagrange function regarding *pi* is:

Setting this partial derivative to be zero, we obtain:

Summing the equation above over , we have:

This means the next parameters *pi*(*t*+1) is:

|  |  |
| --- | --- |
|  | (5.1.37) |

The first-order partial derivative of Lagrange function regarding *qj*|*k* is:

Setting this partial derivative to be zero, we obtain:

Summing the equation above over , we have:

This means the next parameters *qj*|*k*(*t*+1) is:

|  |  |
| --- | --- |
|  | (5.1.38) |

The two steps of GEM algorithm for AMM at some *t*th iteration are shown in table 5.1.5.

|  |
| --- |
| *E-step*:  The conditional probability *P*(*k* | *xi*, *yj*, Θ(*t*)) is calculated based on current parameter Θ(*t*) = (*αk*|*i*(*t*), *pi*(*t*), *qj*|*k*(*t*))*T*, according to equation 5.1.35.  *M-step*:  The next parameter Θ(*t*+1) = (*αk*|*i*(*t*+1), *pi*(*t*+1), *qj*|*k*(*t*+1))*T*, which is a maximizer of *Q*(Θ | Θ(*t*)) with subject to Θ, is calculated by equation 5.1.36, equation 5.1.37, and equation 5.1.38. |

**Table 5.1.5.** E-step and M-step of GEM algorithm for AMM

GEM algorithm converges at some *t*th iteration. At that time, Θ*\** = Θ(*t*+1) = Θ(*t*) is the AMM itself. When AMM is applied into soft clustering, dyadic data is clustered vertically (horizontally) and each *αk*|*i* is coverage ratio of cluster *k* (aspect *k*) according to *xi*. Soft clustering with AMM is also called one-side clustering.

Product-space mixture model (PMM) is derived from SMM with a minor change that the aspect set {1, 2,…, *K*} is Cartesian product of -aspect set {1, 2,…, } and -aspect set {1, 2,…, }. In other words, the aspect space is still symmetric but is checked (stripped) according to two directions and .

|  |  |
| --- | --- |
|  | (5.1.39) |

For every *k* belongs to {1, 2,…, *K*}, there always exists a respective pair: and . However, for each or each , there are many respective *k* (Nguyen, Learning Dyadic Data and Predicting Unaccomplished Co-occurrent Values by Mixture Model, 2020, p. 10).

|  |  |
| --- | --- |
|  | (5.1.40) |

The sign “” denotes correspondence. For example, given aspect set {1, 2, 3, 4, 5, 6}, -aspect set {*a*, *b*, *c*} and -aspect set {*A*, *B*}, we have a set of six correspondences: 1{*a*, *A*}, 2{*a*, *B*}, 3{*a*, *C*}, 4{*b*, *A*}, 5{*b*, *B*}, 6{*b*, *C*}. Given *a* {*a*, *b*, *c*}, we have three correspondences among *a* and aspect set {1, 2, 3, 4, 5, 6} such as *a*1, *a*2, and *a*3.

PMM is defined as follows (Hofmann & Puzicha, Statistical Models for Co-occurrence Data, 1998, p. 4):

|  |  |
| --- | --- |
|  | (5.1.41) |

As usual, *αk* is the probability of aspect *ck* but is the probability of *xi* given of *k* and is the probability of *yj* given of *k*.

The joint probability of *xi*, *yj*, and *k* is:

The parameter of PMM is Θ = (*αk*, , )*T* in which there are *K* + + partial parameters *αk*, , and . Note,

Learning PMM is like learning SMM and so it is not necessary to duplicate the expansion of *Q*(Θ|Θ(*t*)). The two steps of GEM algorithm for PMM at some *t*th iteration are shown in table 5.1.6.

|  |  |  |  |  |  |  |  |  |  |  |  |  |
| --- | --- | --- | --- | --- | --- | --- | --- | --- | --- | --- | --- | --- |
| *E-step*:  The conditional probabilities *P*(*k* | *xi*, *yj*, Θ(*t*)), *P*( | *xi*, *yj*, Θ(*t*)), and *P*( | *xi*, *yj*, Θ(*t*)) are calculated based on current parameter Θ(*t*) = , according to equation 5.1.42, equation 5.1.43, and equation 5.1.44 (Nguyen, Learning Dyadic Data and Predicting Unaccomplished Co-occurrent Values by Mixture Model, 2020, p. 10).   |  |  | | --- | --- | |  | (5.1.42) | |  | (5.1.43) | |  | (5.1.44) |   Please refer to equation 5.1.6 and equation 5.1.10 to comprehend equation 5.1.42.  *M-step*:  The next parameter Θ(*t*+1) = , which is the maximizer of *Q*(Θ | Θ(*t*)) with subject to Θ, is calculated by equation 5.1.45, equation 5.1.46, and equation 5.1.47.   |  |  | | --- | --- | |  | (5.1.45) | |  | (5.1.46) | |  | (5.1.47) | |

**Table 5.1.6.** E-step and M-step of GEM algorithm for PMM

GEM algorithm converges at some *t*th iteration. At that time, Θ*\** = Θ(*t*+1) = Θ(*t*) is the PMM itself. When PMM is applied into soft clustering, dyadic data is clustered in checked (stripped) and each *αk* is coverage ratio of cluster *k* (aspect *k*) but such cluster *k* corresponds to a pair of cluster and cluster . Soft clustering with PMM is also called two-side clustering.

When is valued dyadic data in which every co-occurrence (*xi*, *yj*) is associated with value *z* from random variable *Z* then, SMM is reformed as follows (Nguyen, Learning Dyadic Data and Predicting Unaccomplished Co-occurrent Values by Mixture Model, 2020, pp. 11-12):

|  |  |
| --- | --- |
|  | (5.1.48) |

AMM is reformed as follows:

|  |  |
| --- | --- |
|  | (5.1.49) |

PMM is reformed as follows:

|  |  |
| --- | --- |
|  | (5.1.50) |

Where *fk*(*Z*|*φk*) is the *k*th PDF of *Z* corresponding to the aspect *k*, in which *φk* is parameter of *fk*(*Z*|*φk*). Of course, the parameter Θ now must include all *φk*. It is possible to consider that

Moreover, *Z* is only dependent on *k*.

Note, suppose *xi* and *yj* (as well as *yj* given *xi*) are independent from *Z* given aspect *k*, which is the hint to reform these models.

For example, within SMM, the joint PDF of *xi*, *yj*, *Z*, and *k* is:

Within AMM, the joint PDF of *xi*, *yj*, *Z*, and *k* is:

Within PMM, the joint PDF of *xi*, *yj*, *Z*, and *k* is:

Here it is only necessary to estimate *φk* because how to estimate other partial parameters was aforementioned. By reforming the conditional expectation *Q*(Θ|Θ(*t*)), it is easy to find out that the next parameter *φk*(*t*+1) is solution of following equation:

|  |  |
| --- | --- |
|  | (5.1.51) |

Where *P*(*k* | *xi*(*r*), *yj*(*r*), Θ(*t*)) is specified by equation 5.1.29, equation 5.1.35, and equation 5.1.42 for SMM, AMM, and PMM, respectively. Especially, if *fk*(*Z*|*φk*) distributed normally, the next parameter *φk*(*t*+1) = (*μk*(*t*+1), Σ*k*(*t*+1))*T* containing mean *μk*(*t*+1) and covariance matrix Σ*k*(*t*+1) is calculated as follows:

|  |  |
| --- | --- |
|  | (5.1.52) |

Where *P*(*k* | *xi*(*r*), *yj*(*r*), Θ(*t*)) is specified by equation 5.1.29, equation 5.1.35, and equation 5.1.42 for SMM, AMM, and PMM, respectively. Please refer to equation 5.1.17 and equation 5.1.18 to comprehend equation 5.1.52.

**Example 5.1.2.** Suppose = {*x*1, *x*2} and = {*y*1}, and valued dyadic sample of 2 co-occurrences, = {(*x*1, *y*1, 1, 1), (*x*1, *y*1, 9, 2) }, we will learn SMM given by GEM shown in table 5.1.4. Let *Z* be associative variable which distributes normally with mean – variance *φk* = (*μk*, *σk*2)*T* and is learned by equation 5.1.52. Obviously, we have *Z*(1)=1, *Z*(2)=9, *n*(*x*1, *y*1) = 2, and *n*(*x*2, *y*1) = 0. Suppose the number of aspects is *K*=2. The parameter Θ = (*αk*, *pi*|*k*, *qj*|*k*, *φk*)*T* of SMM is initialized as follows:

|  |
| --- |
|  |

At the 1st iteration, E-step, we have:

At the 1st iteration, M-step, we have:

At the 2nd iteration, E-step, we have:

Note, because the probabilities *P*(*k*=1 | *x*2, *y*1, Θ(2)) and *P*(*k*=2 | *x*2, *y*1, Θ(2)) are arbitrary (0/0), they are assigned to be 0.5.

At the 2nd iteration, M-step, we have:

Therefore, GEM stops at the 2nd iteration with the estimate Θ(2) = Θ(3) = Θ\* = (*αk*\*, *pi*|*k*\*, *qj*|*k*\*, *φk*\*)*T*.

|  |
| --- |
|  |

Similarly, it is easy to learn AMM and PMM ■

## **5.2. Handling missing data**

The goal of MLE, MAP, and EM is to estimate statistical based on sample. Whereas MLE and MAP require complete data, EM accepts hidden data or incomplete data. Therefore, EM is appropriate to handle missing data which contains missing values. Indeed, estimating parameter with missing data is very natural for EM but it is necessary to have a new viewpoint in which missing data is considered as hidden data (*X*). Moreover, the GEM version with joint probability (without mapping function, please see equation 2.12 and equation 2.13) is used and some changes are required. Before describing how to apply EM into handling missing data, we should skim some concepts related to missing data.

Let *X* = (*x*1, *x*2,…, *xn*)*T* be *n*-dimension random variable whose *n* elements are partial random variables *xj* (s). Suppose *X* is composed of two parts such as observed part *Xobs* and missing part *Xmis* such that *X* = {*Xobs*, *Xmis*}. Note, *Xobs* and *Xmis* are considered as random variables.

|  |  |
| --- | --- |
|  | (5.2.1) |

When *X* is observed, *Xobs* and *Xmis* are determined. For example, given *X* = (*x*1, *x*2, *x*3, *x*4)*T*, when *X* is observed as *X* = (*x*1=1, *x*2=?, *x*3=4, *x*4=?, *x*5=9)*T* where question mask “?” denotes missing value, *Xobs* and *Xmis* are determined as *Xobs* = (*x*1=1, *x*3=4, *x*5=9)*T* and *Xmis* = (*x*2=?, *x*4=?)*T*. When *X* is observed as *X* = (*x*1=?, *x*2=3, *x*3=4, *x*4=?, *x*5=?)*T* then, *Xobs* and *Xmis* are determined as *Xobs* = (*x*2=3, *x*3=4)*T* and *Xmis* = (*x*1=?, *x*4=?, *x*5=?)*T*. Let *M* be a set of indices that *xj* (s) are missing when *X* is observed. *M* is called missing index set.

|  |  |
| --- | --- |
|  | (5.2.2) |

Suppose

|  |  |
| --- | --- |
|  | (5.2.3) |

Where,

Let is complementary set of the set *M* given the set {1, 2,…., *n*}.is called existent index set.

|  |  |
| --- | --- |
|  | (5.2.4) |

*M* or can be empty. They are mutual because can be defined based on *M* and vice versa.

Suppose

|  |  |
| --- | --- |
|  | (5.2.5) |

Where,

We have:

|  |  |
| --- | --- |
|  | (5.2.6) |

|  |  |
| --- | --- |
|  | (5.2.7) |

Obviously, dimension of *Xmis* is |*M*| and dimension of *Xobs* is = *n*–|*M*|. Note, when composing *X* from *Xobs* and *Xmis* as *X* = {*Xobs*, *Xmis*}, it is required a right re-arrangement of elements in both *Xobs* and *Xmis*.

Let *Z* = (*z*1, *z*2,…, *zn*)*T* be *n*-dimension random variable whose each element *zj* is binary random variable indicating if *xj* is missing. Random variable *Z* is also called missingness variable.

|  |  |
| --- | --- |
|  | (5.2.8) |

For example, given *X* = (*x*1, *x*2, *x*3, *x*4)*T*, when *X* is observed as *X* = (*x*1=1, *x*2=?, *x*3=4, *x*4=?, *x*5=9)*T*, we have *Xobs* = (*x*1=1, *x*3=4, *x*5=9)*T*, *Xmis* = (*x*2=?, *x*4=?)*T*, and *Z* = (*z*1=0, *z*2=1, *z*3=0, *z*4=1, *z*5=0)*T*.

Generally, when *X* is replaced by a sample = {*X*1, *X*2,…, *XN*} whose *Xi* (s) are iid, let = {*Z*1, *Z*2,…, *ZN*} be a set of missingness variables associated with . All *Zi* (s) are iid too. and can be represented as matrices. Given *Xi*, its associative quantities are *Zi*, *Mi*, and . Let *X* = {*Xobs*, *Xmis*} be random variable representing every *Xi*. Let *Z* be random variable representing every *Zi*. As a convention, *Xobs*(*i*) and *Xmis*(*i*) refer to *Xobs* part and *Xmis* part of *Xi*. We have:

|  |  |
| --- | --- |
|  | (5.2.9) |

For example, given sample of size 4, = {*X*1, *X*2, *X*3, *X*4} in which *X*1 = (*x*11=1, *x*12=?, *x*13=3, *x*14=?)*T*, *X*2 = (*x*21=?, *x*22=2, *x*23=?, *x*24=4)*T*, *X*3 = (*x*31=1, *x*32=2, *x*33=?, *x*34=?)*T*, and *X*4 = (*x*41=?, *x*42=?, *x*43=3, *x*44=4)*T* are iid. Therefore, we also have *Z*1 = (*z*11=0, *z*12=1, *z*13=0, *z*14=1)*T*, *Z*2 = (*z*21=1, *z*22=0, *z*23=1, *z*24=0)*T*, *Z*3 = (*z*31=0, *z*32=0, *z*33=1, *z*34=1)*T*, and *Z*4 = (*z*41=1, *z*42=1, *z*43=0, *z*44=0)*T.* All *Zi* (s) are iid too.

|  |  |  |  |  |  |  |  |  |  |  |
| --- | --- | --- | --- | --- | --- | --- | --- | --- | --- | --- |
|  | *x*1 | *x*2 | *x*3 | *x*4 |  |  | *z*1 | *z*2 | *z*3 | *z*4 |
| *X*1 | 1 | ? | 3 | ? |  | *Z*1 | 0 | 1 | 0 | 1 |
| *X*2 | ? | 2 | ? | 4 |  | *Z*2 | 1 | 0 | 1 | 0 |
| *X*3 | 1 | 2 | ? | ? |  | *Z*3 | 0 | 0 | 1 | 1 |
| *X*4 | ? | ? | 3 | 4 |  | *Z*4 | 1 | 1 | 0 | 0 |

Of course, we have *Xobs*(1) = (*x*11=1, *x*13=3)*T*, *Xmis*(1) = (*x*12=?, *x*14=?)*T*, *Xobs*(2) = (*x*22=2, *x*24=4)*T*, *Xmis*(2) = (*x*21=?, *x*23=?)*T*, *Xobs*(3) = (*x*31=1, *x*32=2)*T*, *Xmis*(3) = (*x*33=?, *x*34=?)*T*, *Xobs*(4) = (*x*43=3, *x*44=4)*T*, and *Xmis*(4) = (*x*41=?, *x*42=?)*T*. We also have *M*1 = {*m*11=2, *m*12=4}, = {=1, =3}, *M*2 = {*m*21=1, *m*22=3}, = {=2, =4}, *M*3 = {*m*31=3, *m*32=4}, = {=1, =2}, *M*4 = {*m*41=1, *m*42=2}, and = {=3, =4}.

Both *X* and *Z* are associated with their own PDFs, as follows:

|  |  |
| --- | --- |
|  | (5.2.10) |

Where Θ and Φ are parameters of PDFs of *X* = {*Xobs*, *Xmis*} and *Z*, respectively. The goal of handling missing data is to estimate Θ and Φ given *X*. Sufficient statistic of *X* = {*Xobs*, *Xmis*} is composed of sufficient statistic of *Xobs* and sufficient statistic of *Xmis*.

|  |  |
| --- | --- |
|  | (5.2.11) |

How to compose *τ*(*X*) from *τ*(*Xobs*) and *τ*(*Xmis*) is dependent on distribution type of the PDF *f*(*X*|Θ).

The joint PDF of *X* and *Z* is main object of handling missing data, which is defined as follows:

|  |  |
| --- | --- |
|  | (5.2.12) |

The PDF of *Xobs* is defined as integral of *f*(*X*|Θ) over *Xmis*:

|  |  |
| --- | --- |
|  | (5.2.13) |

The PDF of *Xmis* is conditional PDF of *Xmis* given *Xobs* is:

|  |  |
| --- | --- |
|  | (5.2.14) |

The notation Θ*M* implies that the parameter Θ*M* of the PDF *f*(*Xmis* | *Xobs*, Θ*M*) is derived from the parameter Θ of the PDF *f*(*X*|Θ), which is function of Θ and *Xobs* asΘ*M* = *u*(Θ, *Xobs*). Thus, Θ*M* is not a new parameter and it is dependent on distribution type.

|  |  |
| --- | --- |
|  | (5.2.15) |

How to determine *u*(Θ, *Xobs*) is dependent on distribution type of the PDF *f*(*X*|Θ).

There are three types of missing data, which depends on relationship between *Xobs*, *Xmis*, and *Z* (Josse, Jiang, Sportisse, & Robin, 2018):

* Missing data (*X* or ) is Missing Completely At Random (MCAR) if the probability of *Z* is independent from on both *Xobs* and *Xmis* such that *f*(Z | *Xobs*, *Xmis*, Φ) = *f*(Z | Φ).
* Missing data (*X* or ) is Missing At Random (MAR) if the probability of *Z* depends on only *Xobs* such that *f*(Z | *Xobs*, *Xmis*, Φ) = *f*(Z | *Xobs*, Φ).
* Missing data (*X* or ) is Missing Not At Random (MNAR) in all other cases, *f*(Z | *Xobs*, *Xmis*, Φ) = *f*(Z | *Xobs*, *Xmis*, Φ).

There are two main approaches for handling missing data (Josse, Jiang, Sportisse, & Robin, 2018):

* Using some statistical models such as EM to estimate parameter with missing data.
* Inputting plausible values for missing values to obtain some complete samples (copies) from the missing data. Later on, every complete sample is used to produce an estimate of parameter by some estimation methods, for example, MLE and MAP. Finally, all estimates are synthesized to produce the best estimate.

Here we focus on the first approach with EM to estimate parameter with missing data. Without loss of generality, given sample = {*X*1, *X*2,…, *XN*} in which all *Xi* (s) are iid, by applying equation 2.13 for GEM with the joint PDF *f*(*Xobs*, *Xmis*, *Z* | Θ, Φ), we consider {*Xobs*, *Z*} as observed part and *Xmis* as hidden part. Let *X* = {*Xobs*, *Xmis*} be random variable representing all *Xi* (s). Let *Xobs*(*i*) denote observed part *Xobs* of *Xi* and let *Zi* be missingness variable corresponding to *Xi*, by following equation 2.13, the expectation *Q*(Θ’, Φ’ | Θ, Φ) becomes:

In short, *Q*(Θ’, Φ’ | Θ, Φ) is specified as follows:

|  |  |
| --- | --- |
|  | (5.2.16) |

Where,

Note, unknowns of *Q*(Θ’, Φ’ | Θ, Φ) are Θ’ and Φ’. Because it is not easy to maximize *Q*(Θ’, Φ’ | Θ, Φ) with regard to Θ’ and Φ’, we assume that the PDF *f*(*X*|Θ) belongs to exponential family.

|  |  |
| --- | --- |
|  | (5.2.17) |

Note,

It is easy to deduce that

|  |  |
| --- | --- |
|  | (5.2.18) |

Therefore,

We have:

Therefore, equation 5.2.19 specifies *Q*1(Θ’|Θ) given *f*(*X*|Θ) belongs to exponential family.

|  |  |
| --- | --- |
|  | (5.2.19) |

Where,

|  |  |
| --- | --- |
|  | (5.2.20) |

|  |  |
| --- | --- |
|  | (5.2.21) |

At M-step of some *t*th iteration, the next parameter Θ(*t*+1) is solution of the equation created by setting the first-order derivative of *Q*1(Θ’|Θ) to be zero. The first-order derivative of *Q*1(Θ’|Θ) is:

By referring table 1.2, we have:

Where,

Thus, the next parameter Θ(*t*+1) is solution of the following equation:

This implies the next parameter Θ(*t*+1) is solution of the following equation:

As a result, at E-step of some *t*th iteration, given current parameter Θ(*t*), the sufficient statistic of *X* is calculated as follows:

|  |  |
| --- | --- |
|  | (5.2.22) |

Where,

Equation 5.2.22 is variant of equation 5.2.11 when *f*(*X*|Θ) belongs to exponential family but how to compose *τ*(*X*) from *τ*(*Xobs*) and *τ*(*Xmis*) is not determined exactly yet.

As a result, at M-step of some *t*th iteration, given *τ*(*t*) and Θ(*t*), the next parameter Θ(*t*+1) is a solution of the following equation:

|  |  |
| --- | --- |
|  | (5.2.23) |

Moreover, at M-step of some *t*th iteration, the next parameter Φ(*t*+1) is a maximizer of *Q*2(Φ | Θ(*t*)) given Θ(*t*) as follows:

|  |  |
| --- | --- |
|  | (5.2.24) |

Where,

|  |  |
| --- | --- |
|  | (5.2.25) |

How to maximize *Q*2(Φ | Θ(*t*)) depends on distribution type of *Zi* which is also formulation of the PDF *f*(Z | *Xobs*, *Xmis*, Φ). For some reasons, such as accelerating estimation speed or ignoring missingness variable *Z* then, the next parameter Φ(*t*+1) will not be estimated.

In general, the two steps of GEM algorithm for handling missing data at some *t*th iteration are summarized in table 5.2.1 with assumption that the PDF of missing data *f*(*X*|Θ) belongs to exponential family.

|  |
| --- |
| *E-step*:  Given current parameter Θ(*t*), the sufficient statistic *τ*(*t*) is calculated according to equation 5.2.22.  Where,  *M-step*:  Given *τ*(*t*) and Θ(*t*), the next parameter Θ(*t*+1) is a solution of equation 5.2.23.  Given Θ(*t*), the next parameter Φ(*t*+1) is a maximizer of *Q*2(Φ | Θ(*t*)) according to equation 5.2.24.  Where, |

**Table 5.2.1.** E-step and M-step of GEM algorithm for handling missing data given exponential PDF

GEM algorithm converges at some *t*th iteration. At that time, Θ*\** = Θ(*t*+1) = Θ(*t*) and Φ*\** = Φ(*t*+1) = Φ(*t*) are optimal estimates. If missingness variable *Z* is ignored for some reasons, parameter Φ is not estimated. Because *Xmis* is a part of *X* and *f*(*Xmis* | *Xobs*, Θ*M*) is derived directly from *f*(*X*|Θ), in practice we can stop GEM after its first iteration was done, which is reasonable enough to handle missing data.

An interesting application of handling missing data is to fill in or predict missing values. For instance, suppose the estimate resulted from GEM is Θ*\**, missing values represented by *τ*(*Xmis*) are fulfilled by expectation of *τ*(*Xmis*) as follows:

|  |  |
| --- | --- |
|  | (5.2.26) |

Where,

Now we survey a popular case that sample = {*X*1, *X*2,…, *XN*} whose *Xi* (s) are iid is MCAR data and *f*(*X*|Θ) is multinormal PDF whereas missingness variable *Z* follows binomial distribution of *n* trials. Let *X* = {*Xobs*, *Xmis*} be random variable representing every *Xi*. Suppose dimension of *X* is *n*. Let *Z* be random variable representing every *Zi*. According to equation 5.2.9, recall that

The PDF of *X* is:

|  |  |
| --- | --- |
|  | (5.2.27) |

Therefore,

The PDF of *Z* is:

|  |  |
| --- | --- |
|  | (5.2.28) |

Therefore,

Where Θ = (*μ*, Σ)*T* and Φ = *p*. Note, given the PDF *f*(*X* | Θ), *µ* is mean and Σ is covariance matrix whose each element *σij* is covariance of *xi* and *xj*.

|  |  |
| --- | --- |
|  | (5.2.29) |

Suppose the probability of missingness at every partial random variable *xj* is *p* and it is independent from *Xobs* and *Xmis*. The quantity *c*(*Z*) is the number of *zj* (s) in *Z* that equal 1. For example, if *Z* = (1, 0, 1, 0)*T* then, *c*(*Z*) = 2. The most important task here is to define equation 5.2.11 and equation 5.2.15 in order to compose *τ*(*X*) from *τ*(*Xobs*), *τ*(*Xmis*) and to extract Θ*M* from Θ when *f*(*X*|Θ) distributes normally.

The conditional PDF of *Xmis* given *Xobs* is also multinormal PDF.

|  |  |
| --- | --- |
|  | (5.2.30) |

Therefore,

Where . We denote

Because only depends on within normal PDF whereas depends on *Xobs*(*i*). Determining the function = *u*(Θ, *Xobs*(*i*)) is now necessary to extract the parameter from Θ given *Xobs*(*i*) when *f*(*Xi*|Θ) is normal distribution.

Let Θ*mis* = (*μmis*, Σ*mis*)*T* be parameter of marginal PDF of *Xmis*, we have:

|  |  |
| --- | --- |
|  | (5.2.31) |

Therefore,

Where,

|  |  |
| --- | --- |
|  | (5.2.32) |

Obviously, Θ*mis*(*i*) is extracted from Θ given indicator *Mi*. Note, is covariance of and .

Let Θ*obs* = (*μobs*, Σ*obs*)*T* be parameter of marginal PDF of *Xobs*, we have:

|  |  |
| --- | --- |
|  | (5.2.33) |

Therefore,

Where,

|  |  |
| --- | --- |
|  | (5.2.34) |

Obviously, Θ*obs*(*i*) is extracted from Θ given indicator or *Mi*. Note, is covariance of and .

We have:

Therefore, it is easy to form the parameter from Θ*mis*(*i*) = (*μmis*(*i*), Σ*mis*(*i*))*T* and Θ*obs*(*i*) = (*μobs*(*i*), Σ*obs*(*i*))*T* as follows (Hardle & Simar, 2013, pp. 156-157):

|  |  |
| --- | --- |
|  | (5.2.35) |

Where from Θ*mis*(*i*) = (*μmis*(*i*), Σ*mis*(*i*))*T* and Θ*obs*(*i*) = (*μobs*(*i*), Σ*obs*(*i*))*T* are specified by equation 5.2.32 and equation 5.2.34. Moreover the *k*x*l* matrix which implies correlation between *Xmis* and *Xobs* is defined as follows (*k* = |*Mi*| and *l* = ):

|  |  |
| --- | --- |
|  | (5.2.36) |

Note, is covariance of and . The *l*x*k* matrix which implies correlation between *Xobs* and *Xmis* is defined as follows:

|  |  |
| --- | --- |
|  | (5.2.37) |

Note, is covariance of and . Therefore, equation 5.2.35 to extract from Θ given *Xobs*(*i*) is an instance of equation 5.2.15. For convenience let,

|  |  |
| --- | --- |
|  | (5.2.38) |

Equation 5.2.38 is result of equation 5.2.35. Given then, is estimated partial mean of and is estimated partial covariance of and given the conditional PDF *f*(*Xmis* | ).

At E-step of some *t*th iteration, given current parameter Θ(*t*), the sufficient statistic of *X* is calculated according to equation 5.2.22. Let,

It is necessary to calculate the sufficient with normal PDF *f*(*Xi*|Θ), which means that we need to define what *τ*1(*t*) and *τ*2(*t*) are. The sufficient statistic of *Xobs*(*i*) is:

The sufficient statistic of *Xmis*(*i*) is:

We also have:

Due to

Where and are and at current iteration, respectively. By referring to equation 5.2.38, we have

And

Where,

Therefore, *τ*1(*t*) is vector and *τ*2(*t*) is matrix and then, the sufficient statistic of *X* at E-step of some *t*th iteration, given current parameter Θ(*t*) is defined as follows:

|  |  |
| --- | --- |
|  | (5.2.39) |

Each is calculated as follows:

|  |  |
| --- | --- |
|  | (5.2.40) |

Please see equation 5.2.35 and equation 5.2.38 to know . Each is calculated as follows:

|  |  |
| --- | --- |
|  | (5.2.41) |

Equation 5.2.39 is an instance of equation 5.2.11, which compose *τ*(*X*) from *τ*(*Xobs*) and *τ*(*Xmis*) when *f*(*X*|Θ) distributes normally. Following is the proof of equation 5.2.41.

If and then, the partial statistic *xiuxiv* is kept intact because *xiu* and *xiv* are in *Xobs* are constant with regard to *f*(*Xmis* | ) If and then, the partial statistic *xiuxiv* is replaced by the expectation as follows:

If and then, the partial statistic *xiuxiv* is replaced by the expectation as follows:

If and then, the partial statistic *xiuxiv* is replaced by the expectation as follows:

At M-step of some *t*th iteration, given *τ*(*t*) and Θ(*t*), the next parameter Θ(*t*+1) = (*μ*(*t*+1), Σ(*t*+1))*T* is a solution of equation 5.2.23.

Due to

Equation 5.2.23 becomes:

Which means that

|  |  |
| --- | --- |
|  | (5.2.42) |

Please see equation 5.2.40 and equation 5.2.41 to know and .

Moreover, at M-step of some *t*th iteration, the next parameter Φ(*t*+1) = *p*(*t*+1) is a maximizer of *Q*2(Φ | Θ(*t*)) given Θ(*t*) according to equation 5.2.24.

Because the PDF of *Zi* is:

The *Q*2(Φ|Θ(*t*)) becomes:

The next parameter Φ(*t*+1) = *p*(*t*+1) is solution of the equation created by setting the first-order derivative of *Q*2(Φ|Θ(*t*)) to be zero, which means that:

It is easy to deduce that the next parameter *p*(*t*+1) is:

|  |  |
| --- | --- |
|  | (5.2.43) |

In general, given sample = {*X*1, *X*2,…, *XN*} whose *Xi* (s) are iid is MCAR data and *f*(*X*|Θ) is multinormal PDF whereas missingness variable *Z* follows binomial distribution of *n* trials, GEM for handling missing data is summarized in table 5.2.2.

|  |
| --- |
| *E-step*:  Given current parameter Θ(*t*) = (*μ*(*t*), Σ(*t*))*T*, the sufficient statistic *τ*(*t*) is calculated according to equation 5.2.39, equation 5.2.40, and equation 5.2.41.  Where and are specified in equation 5.2.35 and equation 5.2.38.  *M-step*:  Given *τ*(*t*) and Θ(*t*), the next parameter Θ(*t*+1) = (*μ*(*t*+1), Σ(*t*+1))*T* is specified by equation 5.2.42.  Given Θ(*t*), the next parameter Φ(*t*+1) = *p*(*t*+1) is specified by equation 5.2.43.  Where *c*(*Zi*) is the number of *zij* (s) in *Zi* that equal 1. |

**Table 5.2.2.** E-step and M-step of GEM algorithm for handling missing data given normal PDF

As aforementioned, an interesting application of handling missing data is to fill in or predict missing values. For instance, suppose the estimate resulted from GEM is Θ*\** = (*μ*\*, Σ\*)*T*, missing part is replaced by as follows:

|  |  |
| --- | --- |
|  | (5.2.44) |

Note, which is extracted from *μ*\* is estimated mean of the conditional PDF of *Xmis* (given *Xobs*) according to equation 5.2.35. Moreover, is estimated partial mean of given the conditional PDF *f*(*Xmis* | ), please see equation 5.2.38 for more details about . As aforementioned, in practice we can stop GEM after its first iteration was done, which is reasonable enough to handle missing data.

It is necessary to have an example for illustrating how to handle missing data with multinormal PDF.

**Example 5.2.1.** Given sample of size two, = {*X*1, *X*2 } in which *X*1 = (*x*11=1, *x*12=?, *x*13=3, *x*14=?)*T* and *X*2 = (*x*21=?, *x*22=2, *x*23=?, *x*24=4)*T* are iid. Therefore, we also have *Z*1 = (*z*11=0, *z*12=1, *z*13=0, *z*14=1)*T* and *Z*2 = (*z*21=1, *z*22=0, *z*23=1, *z*24=0)*T.* All *Zi* (s) are iid too.

|  |  |  |  |  |  |  |  |  |  |  |
| --- | --- | --- | --- | --- | --- | --- | --- | --- | --- | --- |
|  | *x*1 | *x*2 | *x*3 | *x*4 |  |  | *z*1 | *z*2 | *z*3 | *z*4 |
| *X*1 | 1 | ? | 3 | ? |  | *Z*1 | 0 | 1 | 0 | 1 |
| *X*2 | ? | 2 | ? | 4 |  | *Z*2 | 1 | 0 | 1 | 0 |

Of course, we have *Xobs*(1) = (*x*11=1, *x*13=3)*T*, *Xmis*(1) = (*x*12=?, *x*14=?)*T*, *Xobs*(2) = (*x*22=2, *x*24=4)*T* and *Xmis*(2) = (*x*21=?, *x*23=?)*T*. We also have *M*1 = {*m*11=2, *m*12=4}, = {=1, =3}, *M*2 = {*m*21=1, *m*22=3}, and = {=2, =4}. Let *X* and *Z* be random variables representing every *Xi* and every *Zi*, respectively. Suppose *f*(*X*|Θ) is multinormal PDF and missingness variable *Z* follows binomial distribution of 4 trials according to equation 5.2.26 and equation 5.2.27. Dimension of *X* is 4. We will estimate Θ = (*μ*, Σ)*T* and Φ = *p* based on .

The parameters *μ* and Σ are initialized arbitrarily as zero vector and identity vector whereas *p* is initialized 0.5 as follows:

At 1st iteration, E-step, we have:

At 1st iteration, M-step, we have:

At 2nd iteration, E-step, we have:

At 2nd iteration, M-step, we have:

Because the sample is too small for GEM to converge to an exact maximizer with small enough number of iterations, we can stop GEM at the second iteration with Θ(3) = Θ\* = (*μ*\*, Σ\*)*T* and Φ(3) = Φ\* = *p*\* when difference between Θ(2) and Θ(3) is insignificant.

As aforementioned, because *Xmis* is a part of *X* and *f*(*Xmis* | Θ*M*) is derived directly from *f*(*X*|Θ), in practice we can stop GEM after its first iteration was done, which is reasonable enough to handle missing data.

As aforementioned, an interesting application of handling missing data is to fill in or predict missing values. For instance, the missing part *Xmis*(1) of *X*1 = (*x*11=1, *x*12=?, *x*13=3, *x*14=?)*T* is fulfilled by according to equation 5.2.44 as follows:

Now we survey another interesting case that sample = {*X*1, *X*2,…, *XN*} whose *Xi* (s) are iid is MCAR data and *f*(*X*|Θ) is multinomial PDF of *K* trials. We ignore missingness variable *Z* here because it is included in the case of multinormal PDF. Let *X* = {*Xobs*, *Xmis*} be random variable representing every *Xi*. Suppose dimension of *X* is *n*. According to equation 5.2.9, recall that

The PDF of *X* is:

|  |  |
| --- | --- |
|  | (5.2.45) |

Where *xj* are integers and Θ = (*p*1, *p*2,…, *pn*)*T* is the set of probabilities such that

Note, *xj* is the number of trials generating nominal value *j*. Therefore,

Where,

The most important task here is to define equation 5.2.11 and equation 5.2.15 in order to compose *τ*(*X*) from *τ*(*Xobs*), *τ*(*Xmis*) and to extract Θ*M* from Θ when *f*(*X*|Θ) is multinomial PDF.

Let Θ*mis* be parameter of marginal PDF of *Xmis*, we have:

|  |  |
| --- | --- |
|  | (5.2.46) |

Therefore,

Where,

|  |  |
| --- | --- |
|  | (5.2.47) |

Obviously, Θ*mis*(*i*) is extracted from Θ given indicator *Mi*.

Let Θ*obs* be parameter of marginal PDF of *Xobs*, we have:

|  |  |
| --- | --- |
|  | (5.2.48) |

Therefore,

Where,

|  |  |
| --- | --- |
|  | (5.2.49) |

Obviously, Θ*obs*(*i*) is extracted from Θ given indicator or *Mi*.

The conditional PDF of *Xmis* given *Xobs* is calculated based on the PDF of *X* and the marginal PDF of *Xobs* as follows:

This implies that the conditional PDF of *Xmis* given *Xobs* is multinomial PDF of *K* trials.

|  |  |
| --- | --- |
|  | (5.2.50) |

Therefore,

Where

Obviously, the parameter of the conditional PDF is:

|  |  |
| --- | --- |
|  | (5.2.51) |

Therefore, equation 5.2.51 to extract from Θ given *Xobs*(*i*) is an instance of equation 5.2.15. It is easy to check that

At E-step of some *t*th iteration, given current parameter Θ(*t*) = (*p*1(*t*), *p*2(*t*),…, *pn*(*t*))*T*, the sufficient statistic of *X* is calculated according to equation 5.2.22. Let,

The sufficient statistic of *Xobs*(*i*) is:

The sufficient statistic of *Xmis*(*i*) with regard to is:

We also have:

Therefore, the sufficient statistic of *X* at E-step of some *t*th iteration given current parameter Θ(*t*) = (*p*1(*t*), *p*2(*t*),…, *pn*(*t*))*T* is defined as follows:

|  |  |
| --- | --- |
|  | (5.2.52) |

Equation 5.2.52 is an instance of equation 5.2.11, which composes *τ*(*X*) from *τ*(*Xobs*) and *τ*(*Xmis*) when *f*(*X*|Θ) is multinomial PDF.

At M-step of some *t*th iteration, we need to maximize *Q*1(Θ’|Θ) with following constraint

According to equation 5.2.19, we have:

Where quantities *b*(*Xobs*(*i*), *Xmis*) and *a*(Θ’) belongs to the PDF *f*(*X*|Θ) of *X*. Because there is the constraint , we use Lagrange duality method to maximize *Q*1(Θ’|Θ). The Lagrange function *la*(Θ’, *λ* | Θ) is sum of *Q*1(Θ’|Θ) and the constraint , as follows:

Where Θ’ = (*p*1’, *p*2’,…, *pn*’)*T*. Note, *λ* ≥ 0 is called Lagrange multiplier. Of course, *la*(Θ’, *λ* | Θ) is function of Θ’ and *λ*. The next parameter Θ(*t*+1) that maximizes *Q*1(Θ’|Θ) is solution of the equation formed by setting the first-order derivative of Lagrange function regarding Θ’ and *λ* to be zero.

The first-order partial derivative of *la*(Θ’, *λ* | Θ) with regard to Θ’ is:

By referring table 1.2, we have:

Thus,

The first-order partial derivative of *la*(Θ’, *λ* | Θ) with regard to *λ* is:

Therefore, at M-step of some *t*th iteration, given current parameter Θ(*t*) = (*p*1(*t*), *p*2(*t*),…, *pn*(*t*))*T*, the next parameter Θ(*t*+1) is solution of the following equation:

This implies:

Where,

Due to

We obtain *n* equations *Kpj* = –*λ*/*N* + and 1 constraint . Therefore, we have:

Summing *n* equations above, we have:

Suppose every missing value is estimated by such that:

We obtain:

This implies

Such that

Therefore, at M-step of some *t*th iteration, given current parameter Θ(*t*) = (*p*1(*t*), *p*2(*t*),…, *pn*(*t*))*T*, the next parameter Θ(*t*+1) is specified by following equation.

|  |  |
| --- | --- |
|  | (5.2.53) |

In general, given sample = {*X*1, *X*2,…, *XN*} whose *Xi* (s) are iid is MCAR data and *f*(*X*|Θ) is multinomial PDF of *K* trials, GEM for handling missing data is summarized in table 5.2.3.

|  |
| --- |
| *M-step*:  Given *τ*(*t*) and Θ(*t*) = (*p*1(*t*), *p*2(*t*),…, *pn*(*t*))*T*, the next parameter Θ(*t*+1) is specified by equation 5.2.53. |

**Table 5.2.3.** E-step and M-step of GEM algorithm for handling missing data given multinomial PDF

In table 5.2.3, E-step is implied in how to perform M-step. As aforementioned, in practice we can stop GEM after its first iteration was done, which is reasonable enough to handle missing data.

**Example 5.2.2.** Given sample of size two, = {*X*1, *X*2 } in which *X*1 = (*x*11=1, *x*12=?, *x*13=3, *x*14=?)*T* and *X*2 = (*x*21=?, *x*22=2, *x*23=?, *x*24=4)*T* are iid.

|  |  |  |  |  |
| --- | --- | --- | --- | --- |
|  | *x*1 | *x*2 | *x*3 | *x*4 |
| *X*1 | 1 | ? | 3 | ? |
| *X*2 | ? | 2 | ? | 4 |

Of course, we have *Xobs*(1) = (*x*11=1, *x*13=3)*T*, *Xmis*(1) = (*x*12=?, *x*14=?)*T*, *Xobs*(2) = (*x*22=2, *x*24=4)*T* and *Xmis*(2) = (*x*21=?, *x*23=?)*T*. We also have *M*1 = {*m*11=2, *m*12=4}, = {=1, =3}, *M*2 = {*m*21=1, *m*22=3}, and = {=2, =4}. Let *X* be random variable representing every *Xi*. Suppose *f*(*X*|Θ) is multinomial PDF of 10 trials. We will estimate Θ = (*p*1, *p*2, *p*3, *p*4)*T*. The parameters *p*1, *p*2, *p*3, and *p*2 are initialized arbitrarily as 0.25 as follows:

At 1st iteration, M-step, we have:

We stop GEM after the first iteration was done, which results the estimate Θ(2) = Θ\* = (*p*1\*, *p*2\*, *p*3\*, *p*4\*)*T* as follows:

In general, GEM is a powerful tool to handle missing data, which is not so difficult except that how to extract the parameter Θ*M* of the conditional PDF *f*(*Xmis* | *Xobs*, Θ*M*) from the whole parameter Θ of the PDF *f*(*X*|Θ) is most important with note that only *f*(*X*|Θ) is defined firstly and then *f*(*Xmis* | *Xobs*, Θ*M*) is derived from *f*(*X*|Θ). Therefore, equation 5.2.15 is cornerstone of this method. Note, equation 5.2.35 and 5.2.51 are instances of equation 5.2.15 when *f*(*X*|Θ) is multinormal PDF or multinomial PDF.

## **5.3. Learning hidden Markov model**

Simple ideology about EM algorithm was kindled from learning hidden Markov model (HMM) by iterative improvement process but I think that no one had systemized and generalized HMM learning to derive EM until DLR developed EM independently. After EM was popularized, it was conversely used to make clear and explain how to learn HMM. There are many real-world phenomena (so-called states) that we would like to model in order to explain our observations. Often, given sequence of observations symbols, there is demand of discovering real states. For example, there are some states of weather: *sunny*, *cloudy*, *rainy* (Fosler-Lussier, 1998, p. 1). Suppose you are in the room and do not know the weather outside but you are notified observations such as wind speed, atmospheric pressure, humidity, and temperature from someone else. Basing on these observations, it is possible for you to forecast the weather by using hidden Markov model (HMM). Before discussing about HMM, we should glance over the definition of Markov model (MM). First, MM is the statistical model which is used to model the stochastic process. MM is defined as below (Schmolze, 2001):

* Given a finite set of state *S*={*s*1, *s*2,…, *sn*} whose cardinality is *n*. Let ∏ be the *initial state distribution* where *πi*∏ represents the probability that the stochastic process begins in state *si*. In other words *πi* is the initial probability of state *si*, where .
* The stochastic process which is modeled gets only one state from *S* at all time points. This stochastic process is defined as a finite vector *X*=(*x*1, *x*2,…, *xT*) whose element *xt* is a state at time point *t*. The process *X* is called *state stochastic process* and *xt* *S* equals some state *si* *S*. Note that *X* is also called *state sequence*. Time point can be in terms of second, minute, hour, day, month, year, etc. It is easy to infer that the initial probability *πi* = *P*(*x*1*=si*) where *x*1 is the first state of the stochastic process. The state stochastic process *X* must meet fully the *Markov property*, namely, given previous state *xt*–1 of process *X*, the conditional probability of current state *xt* is only dependent on the previous state *xt*–1, not relevant to any further past state (*xt*–2, *xt*–3,…, *x*1). In other words, *P*(*xt | xt*–1, *xt*–2, *xt*–3,…, *x*1) = *P*(*xt | xt*–1) with note that *P*(.) also denotes probability in this research. Such process is called first-order Markov process.
* At each time point, the process changes to the next state based on the *transition probability distribution* *aij*, which depends only on the previous state. So *aij* is the probability that the stochastic process changes current state *si* to next state *sj*. It means that *aij* = *P*(*xt=sj* | *xt–*1*=si*) = *P*(*xt+*1*=sj* | *xt=si*). The probability of transitioning from any given state to some next state is 1*,* we have. All transition probabilities *aij* (s) constitute the *transition probability matrix* *A*. Note that *A* is *n* by *n* matrix because there are *n* distinct states. It is easy to infer that matrix *A* represents state stochastic process *X*. It is possible to understand that the initial probability matrix ∏ is degradation case of matrix *A*.

Briefly, MM is the triple 〈*S*, *A*,∏〉. In typical MM, states are observed directly by users and transition probabilities (*A* and ∏) are unique parameters. Otherwise, hidden Markov model (HMM) is similar to MM except that the underlying states become hidden from observer, they are hidden parameters. HMM adds more output parameters which are called observations. Each state (hidden parameter) has the conditional probability distribution upon such observations. HMM is responsible for discovering hidden parameters (states) from output parameters (observations), given the stochastic process. The HMM has further properties as below (Schmolze, 2001):

* Suppose there is a finite set of possible observations Φ = {*φ*1, *φ*2,…, *φm*} whose cardinality is *m*. There is the second stochastic process which produces *observations* correlating with hidden states. This process is called *observable stochastic process*, which is defined as a finite vector *O* = (*o*1, *o*2,…, *oT*) whose element *ot* is an observation at time point *t*. Note that *ot* Φ equals some *φk*. The process *O* is often known as *observation sequence*.
* There is a probability distribution of producing a given observation in each state. Let *bi*(*k*) be the probability of observation *φk* when the state stochastic process is in state *si*. It means that *bi*(*k*) = *bi*(*ot=φk*) = *P*(*ot=φk* | *xt=si*). The sum of probabilities of all observations which observed in a certain state is 1, we have. All probabilities of observations *bi*(*k*) constitute the *observation probability matrix* *B*. It is convenient for us to use notation *bik* instead of notation *bi*(*k*). Note that *B* is *n* by *m* matrix because there are *n* distinct states and *m* distinct observations. While matrix *A* represents state stochastic process *X*, matrix *B* represents observable stochastic process *O*.

Thus, HMM is the 5-tuple ∆ = 〈*S*, Φ, *A*, *B*,∏〉. Note that components *S*,Φ, *A*, *B*,and∏ are often called parameters of HMM in which *A*, *B*,and∏ are essential parameters. Going back weather example, suppose you need to predict how weather tomorrow is: *sunny*, *cloudy* or *rainy* since you know only observations about the humidity: *dry*, *dryish*, *damp*, *soggy*. The HMM is totally determined based on its parameters *S*,Φ, *A*, *B*,and∏ according to weather example. We have *S* = {*s*1=*sunny*, *s*2=*cloudy*, *s*3=*rainy*}, Φ = {*φ*1=*dry*, *φ*2=*dryish*, *φ*3=*damp*, *φ*4=*soggy*}. Transition probability matrix *A* is shown in table 5.3.1.

|  |  |  |  |  |
| --- | --- | --- | --- | --- |
|  | | Weather current day  (Time point *t*) | | |
| *sunny* | *cloudy* | *rainy* |
| Weather previous day  (Time point *t* –1) | *sunny* | *a*11=0.50 | *a*12=0.25 | *a*13=0.25 |
| *cloudy* | *a*21=0.30 | *a*22=0.40 | *a*23=0.30 |
| *rainy* | *a*31=0.25 | *a*32=0.25 | *a*33=0.50 |

**Table 5.3.1.** Transition probability matrix A

From table 5.3.1, we have *a*11+*a*12+*a*13=1, *a*21+*a*22+*a*23=1, *a*31+*a*32+*a*33=1. Initial state distribution specified as uniform distribution is shown in table 5.3.2.

|  |  |  |
| --- | --- | --- |
| *sunny* | *cloudy* | *rainy* |
| *π*1=0.33 | *π*2=0.33 | *π*3=0.33 |

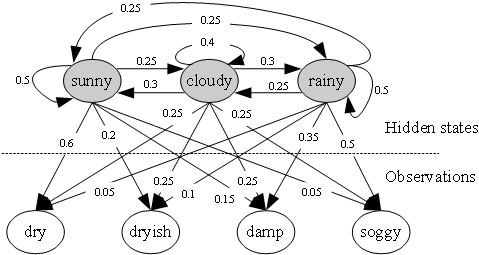
**Table 5.3.2.** Uniform initial state distribution ∏

From table 5.3.2, we have *π*1+*π*2+*π*3=1. Observation probability matrix *B* is shown in table 5.3.3.

|  |  |  |  |  |  |
| --- | --- | --- | --- | --- | --- |
|  | | Humidity | | | |
| *dry* | *dryish* | *damp* | *soggy* |
| Weather | *sunny* | *b*11=0.60 | *b*12=0.20 | *b*13=0.15 | *b*14=0.05 |
| *cloudy* | *b*21=0.25 | *b*22=0.25 | *b*23=0.25 | *b*24=0.25 |
| *rainy* | *b*31=0.05 | *b*32=0.10 | *b*33=0.35 | *b*34=0.50 |

**Table 5.3.3.** Observation probability matrix *B*

From table 5.3.3, we have *b*11+*b*12+*b*13+*b*14=1, *b*21+*b*22+*b*23+*b*24=1, *b*31+*b*32+*b*33+*b*34=1. The whole weather HMM is depicted in figure 5.3.1.



**Figure 5.3.1.** HMM of weather forecast (hidden states are shaded)

There are three problems of HMM (Schmolze, 2001) (Rabiner, 1989, pp. 262-266):

* Given HMM ∆ and an observation sequence *O* = {*o*1, *o*2,…, *oT*} where *ot* Φ, how to calculate the probability *P*(*O*|*∆*) of this observation sequence. Such probability *P*(*O*|*∆*) indicates how much the HMM ∆ affects on sequence *O*. This is *evaluation problem* or *explanation problem*. Note that it is possible to denote *O* = {*o*1 → *o*2 →…→ *oT*} and the sequence *O* is aforementioned observable stochastic process.
* Given HMM ∆ and an observation sequence *O* = {*o*1, *o*2,…, *oT*} where *ot* Φ, how to find the sequence of states *X* = {*x*1, *x*2,…, *xT*} where *xt* *S* so that *X* is most likely to have produced the observation sequence *O*. This is *uncovering problem*. Note that the sequence *X* is aforementioned state stochastic process.
* Given HMM ∆ and an observation sequence *O* = {*o*1, *o*2,…, *oT*} where *ot* Φ, how to adjust parameters of ∆ such as initial state distribution ∏, transition probability matrix *A*, and observation probability matrix *B* so that the quality of HMM ∆ is enhanced. This is *learning problem*.

This sub-section focuses on the third problem which is learning problem because HMM learning relates to EM algorithm. Before mentioning learning problem, we need to comprehend the important concept of forward-backward procedure related to evaluation problem. Therefore, this sub-section also mentions evaluation problem. Indeed, evaluation problem is solved by forward-backward procedure. According to (Rabiner, 1989, pp. 262-263), there is a so-called *forward-backward procedure* to decrease computational cost for determining the probability *P*(*O|*Δ). Let *αt*(*i*) be the joint probability of partial observation sequence {*o*1, *o*2,…, *ot*} and state *xt*=*si* where , specified by equation 5.3.1.

|  |  |
| --- | --- |
|  | (5.3.1) |

The joint probability *αt*(*i*) is also called *forward variable* at time point *t* and state *si*. The product *αt*(*i*)*aij* where *aij* is the transition probability from state *i* to state *j* counts for probability of join event that partial observation sequence {*o*1, *o*2,…, *ot*} exists and the state *si* at time point *t* is changed to *sj* at time point *t+*1.

(Due to multiplication rule)

(Because the partial observation sequence {*o*1, *o*2,…, *ot*} is independent from next state *xt+*1 given current state *xt*)

(Due to multiplication rule)

Summing product *αt*(*i*)*aij* over all *n* possible states of *xt* produces probability of join event that partial observation sequence {*o*1, *o*2,…, *ot*} exists and the next state is *xt+*1=*sj* regardless of the state *xt*.

The forward variable at time point *t+*1 and state *sj* is calculated on *αt*(*i*) as follows:

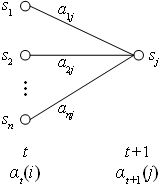
(Due to multiplication rule)

(Due to observations are mutually independent)

Where *bj*(*ot+*1) is the probability of observation *ot+*1 when the state stochastic process is in state *sj*, please see an example of observation probability matrix shown in table 5.3.3. In brief, please pay attention to recurrence property of forward variable specified by equation 5.3.2.

|  |  |
| --- | --- |
|  | (5.3.2) |

The aforementioned construction of forward recurrence equation 5.3.2 is essentially to build up Markov chain, illustrated by figure 5.3.2 (Rabiner, 1989, p. 262).



**Figure 5.3.2.** Construction of recurrence equation for forward variable

According to the forward recurrence equation 5.3.2, given observation sequence *O* = {*o*1, *o*2,…, *oT*}, we have:

The probability *P*(*O|*Δ) is sum of *αT*(*i*) over all *n* possible states of *xT*, specified by equation 5.3.3.

|  |  |
| --- | --- |
|  | (5.3.3) |

The forward-backward procedure to calculate the probability *P*(*O|*Δ), based on forward equation 5.3.2 and 5.3.3, includes three steps as shown in table 5.3.4 (Rabiner, 1989, p. 262).

|  |
| --- |
| 1. Initialization step: Initializing *α*1(*i*) = *bi*(*o*1)*πi* for all 2. Recurrence step: Calculating all *αt+*1(*j*) for all and according to equation 5.3.2. 3. Evaluation step: Calculating the probability according to equation 5.3.3. |

**Table 5.3.4.** Forward-backward procedure based on forward variable to calculate the probability *P*(*O|*Δ)

Thus, evaluation problem is solved by forward-backward procedure shown in table 5.3.4.

There is interesting thing that the forward-backward procedure can be implemented based on so-called *backward variable*. Let *βt*(*i*) be the backward variable which is conditional probability of partial observation sequence {*ot*, *ot+*1,…, *oT*} given state *xt*=*si* where , specified by equation 5.3.4.

|  |  |
| --- | --- |
|  | (5.3.4) |

We have

(Because observations *ot+*1, *ot+*2,…, *oT* are mutually independent)

(Because partial observation sequence *ot+*1, *ot+*2,…, *oT* is independent from state *xt* at time point t)

(Due to multiplication rule)

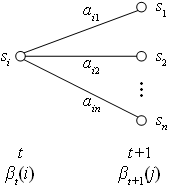
Summing the product *aijbj*(*ot+*1)*βt+*1(*j*) over all *n* possible states of *xt+*1=*sj*, we have:

(Due to the total probability rule)

In brief, the recurrence property of backward variable specified by equation 5.3.5.

|  |  |
| --- | --- |
|  | (5.3.5) |

Where *bj*(*ot+*1) is the probability of observation *ot+*1 when the state stochastic process is in state *sj*, please see an example of observation probability matrix shown in table 5.3.3. The construction of backward recurrence equation 5.3.5 is essentially to build up Markov chain, illustrated by figure 5.3.3 (Rabiner, 1989, p. 263).



**Figure 5.3.3.** Construction of recurrence equation for backward variable

According to the backward recurrence equation 5.3.5, given observation sequence *O* = {*o*1, *o*2,…, *oT*}, we have:

The product *πibi*(*o*1)*β*1(*i*) is:

(Because observations *o*1, *o*2,…, *oT* are mutually independent)

It implies that the probability *P*(*O|*Δ) is:

(Due to the total probability rule)

Shortly, the probability *P*(*O|*Δ) is sum of product *πibi*(*o*1)*β*1(*i*) over all *n* possible states of *x*1=*si*, specified by equation 5.3.6.

|  |  |
| --- | --- |
|  | (5.3.6) |

The forward-backward procedure to calculate the probability *P*(*O|*Δ), based on backward equations 5.3.5 and 5.3.6, includes three steps as shown in table 5.3.5 (Rabiner, 1989, p. 263).

|  |
| --- |
| 1. Initialization step: Initializing *βT*(*i*) = 1for all 2. Recurrence step: Calculating all *βt*(*i*) for all and *t=T*–1, *t=T*–2,…, *t=*1, according to equation 5.3.5. 3. Evaluation step: Calculating the probability *P*(*O|*Δ) according to equation 5.3.6, |

**Table 5.3.5.** Forward-backward procedure based on backward variable to calculate the probability *P*(*O|*Δ)

Thus, evaluation problem is solved by forward-backward procedure shown in table 5.3.5.

Let *γt*(*i*) be joint probability that the stochastic process is in state *si* at time point *t* with observation sequence *O* = {*o*1, *o*2,…, *oT*}, equation 5.3.7 specifies this probability based on forward variable *αt* and backward variable *βt*.

|  |  |
| --- | --- |
|  | (5.3.7) |

The variable *γt*(*i*) is also called *individually optimal criterion* with note that forward variable *αt* and backward variable *βt* are calculated according to recurrence equations 5.3.2 and 5.3.5, respectively. Following is proof of equation 5.3.7.

(Due to Bayes’ rule)

(Due to multiplication rule)

(Because observations *o*1, *o*2,…, *oT* are observed independently)

(According to equations 5.3.1 and 5.3.4 for determining forward variable and backward variable)

Recall that the learning problem is to adjust parameters by maximizing probability of observation sequence *O*, as follows:

Where , , are parameter estimates and so, the purpose of HMM learning problem is to determine them. In the first stage, the well-known Baum-Welch algorithm (Rabiner, 1989). Essentially, EM algorithm is the generalized and multi-usage version of Baum-Welch algorithm. Hence, I only mention how to apply EM into HMM learning. The observation sequence *O* = {*o*1, *o*2,…, *oT*} and state sequence *X* = {*x*1, *x*2,…, *xT*} are observed data and hidden data within context of EM algorithm, respectively. Note *O* and *X* is now represented in sequence. According to EM algorithm, the parameter estimate is determined as follows:

Where Δ*r* = (*Ar*, *Br*, ∏*r*) is the known parameter at the current iteration. Note that we use notation Δ*r* instead of popular notation Δ*t* in order to distinguish iteration indices of EM algorithm from time points inside observation sequence *O* and state sequence *X*. Using equation 2.12 in discrete case, the EM conditional expectation *Q*(Δ|Δ*r*) in accordance with HMM is:

(Because observations *o*1, *o*2,…, *oT* are mutually independent)

(Because each observations *ot* is only dependent on state *xt*)

(Because each state *xt* is only dependent on previous state *xt–*1)

(Due to recurrence onprobability *P*(*x*1, *x*2,…, *xt*))

It is conventional that where *x*0 is pseudo-state, equation 5.3.8 specifies general EM conditional expectation for HMM:

|  |  |
| --- | --- |
|  | (5.3.8) |

Let and are two index functions so that (Ramage, 2007, p. 3)

We have:

Because of the convention , matrix ∏ is degradation case of matrix *A* at time point *t=*1. In other words, the initial probability *πj* is equal to the transition probability *aij* from pseudo-state *x*0 to state *x*1=*sj*.

Note that *n*=|*S*| is the number of possible states and *m*=|Φ| is the number of possible observations.

Shortly, the EM conditional expectation for HMM is specified by equation 5.3.9.

|  |  |
| --- | --- |
|  | (5.3.9) |

Where,

Note that the conditional expectation *Q*(Δ|Δ*r*) is function of Δ. There are two constraints for HMM as follows:

Maximizing *Q*(Δ|Δ*r*) with subject to these constraints is optimization problem that is solved by Lagrange duality theorem (Jia, 2013, p. 8). Original optimization problem mentions minimizing target function but it is easy to infer that maximizing target function shares the same methodology. Let *l*(Δ, *λ*, *μ*) be Lagrange function constructed from *Q*(Δ|Δ*r*) together with these constraints (Ramage, 2007, p. 9), we have equation 5.3.10 for specifying HMM Lagrange function as follows:

|  |  |
| --- | --- |
|  | (5.3.10) |

Where *λ* is *n*-component vector *λ* = (*λ*1, *λ*2,…, *λn*) and *μ* is *m*-component vector *μ* = (*μ*1, *μ*2,…, *μm*). Factors *λi* ≥ 0 and *μj* ≥ 0 are called Lagrange multipliers or Karush-Kuhn-Tucker multipliers (Wikipedia, Karush–Kuhn–Tucker conditions, 2014) or dual variables. The expectation *Q*(Δ|Δ*r*) is specified by equation 5.3.9. The parameter estimate is extreme point of the Lagrange function. According to Lagrange duality theorem (Boyd & Vandenberghe, 2009, p. 216) (Jia, 2013, p. 8), we have:

The parameter estimate is determined by setting partial derivatives of *l*(Δ, *λ*, *μ*) with respect to *aij* and *bj*(*k*) to be zero. The partial derivative of *l*(Δ, *λ*, *μ*) with respect to *aij* is:

Setting the partial derivative to be zero:

The parameter estimate is solution of equation , we have:

It is required to estimate the Lagrange multiplier *λi*. The multiplier estimate is determined by setting the partial derivative of *l*(Δ, *λ*, *μ*) with respect to *λi* to be zero as follows:

Substituting for *aij*, we have:

It implies:

Where, is index function.

Substituting for *λi* inside

We have:

Evaluating the numerator, we have:

(Due to total probability rule)

(Due to multiplication rule)

Evaluating the denominator, we have:

(Due to total probability rule)

(Due to multiplication rule)

It implies

Because of the convention , the estimate is fixed as follows:

The estimate of initial probability is known as specific estimate from pseudo-state *x*0 to state *x*1=*sj*. It means that

Recall that the parameter estimate is determined by setting partial derivatives of *l*(Δ, *λ*, *μ*) with respect to *aij* and *bj*(*k*) to be zero. The parameter estimate was determined. Now it is required to calculate the parameter estimate . The partial derivative of Lagrange function *l*(Δ, *λ*, *μ*) with respect to *bj*(*k*) is:

Setting the partial derivative to be zero:

The parameter estimate is solution of equation , we have:

It is required to estimate the Lagrange multiplier *μj*. The multiplier estimate is determined by setting the partial derivative of *l*(Δ, *λ*, *μ*) with respect to *μj* to be zero as follows:

Substituting for *bj*(*k*) we have:

It implies:

Where, is index function.

Substituting for *μj* inside

We have:

Evaluating the numerator, we have:

(Due to total probability rule)

(Due to multiplication rule)

Note, the expression expresses the sum of probabilities *P*(*O*, *xt*=*sj* | Δ*r*) over *T* time points with condition *ot* = *φk*.

Evaluating the denominator, we have:

(Due to total probability rule)

(Due to multiplication rule)

It implies

In general, the parameter estimate is totally determined as follows:

As a convention, we use notation Δ instead of Δ*r* for denoting known HMM at current iteration of EM algorithm. We have equation 5.3.11 for specifying HMM parameter estimate given current parameter Δ = (*aij*, *bj*(*k*), *πj*) as follows:

|  |  |
| --- | --- |
|  | (5.3.11) |

The parameter estimate is the ultimate solution of the learning problem. As seen in equation 5.3.11, it is necessary to calculate probabilities *P*(*O*, *xt–*1=*si*, *xt*=*sj*) and *P*(*O*, *xt–*1*=si*) when other probabilities *P*(*O*, *xt*=*sj*), *P*(*O*, *x*1=*si*), and *P*(*O*, *x*1=*sj*) are represented by the joint probability *γt* specified by equation 5.3.7.

Let *ξt*(*i*, *j*) is the joint probability that the stochastic process receives state *si* at time point *t*–1 and state *sj* at time point *t* given observation sequence *O* (Rabiner, 1989, p. 264).

Given forward variable *αt* and backward variable *βt*, if , we have:

(Due to multiplication rule)

(Because the partial observation sequence {*o*1, *o*2,…, *ot*} is independent from current state *xt* given previous state *xt–*1)

(Due to multiplication rule)

(Because observations *ot*, *ot+*1, *ot+*2,…, *oT* are mutually independent)

(Due to multiplication rule)

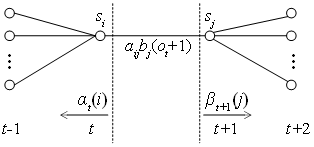
(Due to multiplication rule)

In general, equation 5.3.12 determines the joint probability *ξt*(*i*, *j*) based on forward variable *αt* and backward variable *βt*.

|  |  |
| --- | --- |
|  | (5.3.12) |

Where forward variable *αt* and backward variable *βt* are calculated by previous recurrence formulas 5.3.2 and 5.3.5.

Shortly, the joint probability *ξt*(*i*, *j*) is constructed from forward variable and backward variable, as seen in figure 5.3.4 (Rabiner, 1989, p. 264).



**Figure 5.3.4.** Construction of the joint probability *ξt*(*i*, *j*)

Recall that *γt*(*j*) is the joint probability that the stochastic process is in state *sj* at time point *t* with observation sequence *O* = {*o*1, *o*2,…, *oT*}, specified by previous equation 5.3.7.

According to total probability rule, it is easy to infer that *γt* is sum of *ξt* over all states with , as seen in equation 5.3.13.

|  |  |
| --- | --- |
|  | (5.3.13) |

Deriving from equations 5.3.12 and 5.3.13, we have:

By extending equation 5.3.11, we receive equation 5.3.14 for specifying HMM parameter estimate given current parameter Δ = (*aij*, *bi*(*k*), *πi*) in detailed.

|  |  |
| --- | --- |
|  | (5.3.14) |

The equation 5.3.14 and its proof are found in (Ramage, 2007, pp. 9-12). Followings are interpretations relevant to the joint probabilities *ξt*(*i*, *j*) and *γt*(*j*) with observation sequence *O*.

* The sum expresses expected number of transitions from state *si* to state *sj* (Rabiner, 1989, p. 265).
* The double sum expresses expected number of transitions from state *si* (Rabiner, 1989, p. 265).
* The sum expresses expected number of times in state *sj* and in observation *φk* (Rabiner, 1989, p. 265).
* The sum expresses expected number of times in state *sj* (Rabiner, 1989, p. 265).

Followings are interpretations of the parameter estimate :

* The transition estimate is expected frequency of transitions from state *si* to state *sj*.
* The observation estimate is expected frequency of times in state *sj* and in observation *φk*.
* The initial estimate is (normalized) expected frequency of state *sj* at the first time point (*t=*1).

It is easy to infer that the parameter estimate is based on joint probabilities *ξt*(*i*, *j*) and *γt*(*j*) which, in turn, are based on current parameter Δ = (*aij*, *bj*(*k*), *πj*). The EM conditional expectation is determined by joint probabilities *ξt*(*i*, *j*) and *γt*(*j*); so, the main task of E-step in EM algorithm is essentially to calculate the joint probabilities *ξt*(*i*, *j*) and *γt*(*j*) according to equations 5.3.12 and 5.3.7. The EM conditional expectation gets maximal at estimate and so, the main task of M-step in EM algorithm is essentially to calculate according to equation 5.3.14. The EM algorithm is interpreted in HMM learning problem, as shown in table 5.3.6.

|  |
| --- |
| Starting with initial value for Δ, each iteration in EM algorithm has two steps:   1. *E-step*: Calculating the joint probabilities *ξt*(*i*, *j*) and *γt*(*j*) according to equations 5.3.12 and 5.3.7 given current parameter Δ = (*aij*, *bj*(*k*), *πj*).   Where forward variable *αt* and backward variable *βt* are calculated by previous recurrence equations 5.3.2 and 5.3.5.   1. *M-step*: Calculating the estimate based on the joint probabilities *ξt*(*i*, *j*) and *γt*(*j*) determined at E-step, according to equation 5.3.14.   The estimate becomes the current parameter for next iteration.  EM algorithm stops when it meets the terminating condition, for example, the difference of current parameter Δ and next parameter is insignificant. It is possible to define a custom terminating condition. |

**Table 5.3.6.** EM algorithm for HMM learning problem

The algorithm to solve HMM learning problem shown in table 5.3.6 is known as Baum-Welch algorithm by Leonard E. Baum and Lloyd R. Welch (Rabiner, 1989) which is as special case of EM algorithm for HMM. Please see document “Hidden Markov Models Fundamentals” by Ramage (Ramage, 2007, pp. 8-13) for more details about HMM learning problem.

# **6. Discussions**

The convergence of GEM is based on the assumption that *Q*(Θ’ | Θ) is smooth enough but *Q*(Θ’ | Θ) may not be smooth in practice when *f*(*X* | Θ) may be discrete probability function. For example, when *f*(*X* | Θ) and *k*(*X* | *Y*, Θ) are discrete, equation 2.8 becomes

This discussion section goes beyond traditional variants of GEM algorithm when *Q*(Θ’ | Θ) is not smooth. Therefore, heuristic optimization methods which simulate social behavior, such as particle swarm optimization (PSO) algorithm (Poli, Kennedy, & Blackwell, 2007) and artificial bee colony (ABC) algorithm, are useful in case that there is no requirement of existence of derivative. Moreover, these heuristic methods aim to find global optimizer. I propose an association of GEM and PSO which produces a so-called quasi-PSO-GEM algorithm in which M-step is implemented by one-time PSO (Wikipedia, Particle swarm optimization, 2017). Given current *t*th iteration, Θ(*t*) is modeled as swarm’s best position. Suppose there are *n* particles and each particle *i* has current velocity *Vi*(*t*), current positions Ψ*i*(*t*), and best position Φ*i*(*t*). At each iteration, it is expected that these particles move to swarm’s new best position which is the next parameter Θ(*t*+1). The swarm’s best position at the final iteration is expected as Θ*\**. Table 6.2 is the proposal of quasi-PSO-GEM algorithm.

|  |  |  |  |  |  |  |
| --- | --- | --- | --- | --- | --- | --- |
| *E-step*:  As usual, *Q*(Θ | Θ(*t*)) is determined based on current Θ(*t*) according to equation 2.8. Actually, *Q*(Θ | Θ(*t*)) is formulated as function of Θ.  *M-step* includes four sub-steps:   1. Calculating the next velocity *Vi*(*t*+1) of each particle based on its current velocity *Vi*(*t*), its current positions Ψ*i*(*t*), its best positions Φ*i*(*t*), and the swarm’s best position Θ(*t*):  |  |  | | --- | --- | |  | (6.1) |   Where *ω*, *ϕ*1, and *ϕ*2 are particular parameters of PSO (Poli, Kennedy, & Blackwell, 2007, pp. 3-4) whereas *r* is a random number such that 0 < *r* < 1 (Wikipedia, Particle swarm optimization, 2017).   1. Calculating the next position Ψ*i*(*t*+1) of each particle based on its current position Ψ*i*(*t*) and its current velocity *Vi*(*t*):  |  |  | | --- | --- | |  | (6.2) |  1. If *Q*(Φ*i*(*t*) | Θ(*t*)) < *Q*(Ψ*i*(*t*+1) | Θ(*t*)) then, the next best position of each particle *i* is re-assigned as Φ*i*(*t*+1) = Ψ*i*(*t*+1). Otherwise, such next best position is kept intact as Φ*i*(*t*+1) = Φ*i*(*t*). 2. The next parameter Θ(*t*+1) is the swarm’s new best position over the best positions of all particles:  |  |  | | --- | --- | |  | (6.3) |   If the bias |Θ(*t*+1) – Θ(*t*)| is small enough, the algorithm stops. Otherwise, Θ(*t*+1) and all *Vi*(*t*+1), Ψ*i*(*t*+1), Φ*i*(*t*+1) become current parameters in the next iteration. |

**Table 6.1.** E-step and M-step of the proposed quasi-PSO-GEM

At the first iteration, each particle is initialized with Ψ*i*(1) = Φ*i*(1) = Θ(1) and uniformly distributed velocity *Vi*(1). Note, Θ(1) is initialized arbitrarily. Other termination criteria can be used, for example, *Q*(Θ | Θ(*t*)) is large enough or the number of iterations is large enough.

We cannot prove mathematically convergence of quasi-PSO-GEM but we expect that Θ(*t*+1) resulted from equation 6.3 is an approximation of Θ*\** at the last iteration after a large enough number of iterations. However, quasi-PSO-GEM tendentiously approaches global maximizer of *L*(Θ), regardless of whether *L*(Θ) is concave. Hence, it is necessary to make experiment on quasi-PSO-GEM.

There are many other researches which combine EM and PSO but the proposed quasi-PSO-GEM algorithm has different ideology when it one-time PSO is embed into M-step to maximize *Q*(Θ | Θ(*t*)) and so the ideology of quasi-PSO-GEM is near to the ideology of Newton-Raphson process. With different viewpoint, some other researches combine EM and PSO in order to solving better a particular problem instead of improving EM itself. For example, Ari and Aksoy (Ari & Aksoy, 2010) used PSO to solve optimization problem of the clustering algorithm based on mixture model and EM. Rajeswari and Gunasundari (Rajeswari & Gunasundari, 2016) proposed EM for PSO based weighted clustering. Zhang, Zhuang, Gao, Luo, Ran, and Du (Zhang, et al., 2014) proposed a so-called PSO-EM algorithm to make optimum use of PSO in partial E-step in order solve the difficulty of integrals in normal compositional model. Golubovic, Olcan, and Kolundzija (Golubovic, Olcan, & Kolundzija, 2007) proposed a few modifications of the PSO algorithm which are applied to EM optimization of a broadside antenna array. Tang, Song, and Liu (Tang, Song, & Liu, 2014) proposed a hybrid clustering method based on improved PSO and EM clustering algorithm to overcome drawbacks of EM clustering algorithm. Tran, Vo, and Lee (Tran, Vo, & Lee, 2013) proposed a novel clustering algorithm for image segmentation by employing the arbitrary covariance matrices that uses PSO for the estimation of Gaussian mixture models.

# **References**

Ari, C., & Aksoy, S. (2010). Maximum Likelihood Estimation of Gaussian Mixture Models Using Particle Swarm Optimization. *The 20th International Conference on Pattern Recognition (ICPR 2010)* (pp. 746-749). Istanbul: IEEE. Retrieved February 21, 2018, from www.cs.bilkent.edu.tr/~saksoy/papers/icpr10\_clustering.pdf

Bilmes, J. A. (1998). *A Gentle Tutorial of the EM Algorithm and its Application to Parameter Estimation for Gaussian Mixture and Hidden Markov Models.* International Computer Science Institute, Department of Electrical Engineering and Computer Science. Berkeley: University of Washington. Retrieved from http://melodi.ee.washington.edu/people/bilmes/mypubs/bilmes1997-em.pdf

Borman, S. (2004). *The Expectation Maximization Algorithm - A short tutorial.* University of Notre Dame, Department of Electrical Engineering. South Bend, Indiana: Sean Borman's Home Page.

Boyd, S., & Vandenberghe, L. (2009). *Convex Optimization.* New York, NY, USA: Cambridge University Press.

Burden, R. L., & Faires, D. J. (2011). *Numerical Analysis* (9th Edition ed.). (M. Julet, Ed.) Brooks/Cole Cengage Learning.

Collins, M., & Barzilay, R. (2005). *Advanced Natural Language Processing - The EM Algorithm.* Massachusetts Institute of Technology, Electrical Engineering and Computer Science. MIT OpenCourseWare. Retrieved October 9, 2020, from https://ocw.mit.edu/courses/electrical-engineering-and-computer-science/6-864-advanced-natural-language-processing-fall-2005/lecture-notes/lec5.pdf

Dempster, A. P., Laird, N. M., & Rubin, D. B. (1977). Maximum Likelihood from Incomplete Data via the EM Algorithm. (M. Stone, Ed.) *Journal of the Royal Statistical Society, Series B (Methodological), 39*(1), 1-38.

Dinh, L. T., Pham, D. H., Nguyen, T. X., & Ta, P. D. (2000). *Univariate Analysis - Principles and Practices.* (K. H. Ha, T. V. Ngo, & D. H. Pham, Eds.) Hanoi, Vietnam: Hanoi National University Publisher. Retrieved from http://www.ebook.edu.vn/?page=1.14&view=11156

Fosler-Lussier, E. (1998). *Markov Models and Hidden Markov Models: A Brief Tutorial.* Technical Report TR-98-041, International Computer Science Institute, USA.

Golubovic, R. M., Olcan, D. I., & Kolundzija, B. M. (2007). Particle Swarm Optimization Algorithm and Its Modifications Applied to EM Problems. In B. D. Milovanović (Ed.), *The 8th International Conference on Telecommunications in Modern Satellite, Cable and Broadcasting Services (ELSIKS 2007)* (pp. 427-430). Nis, Serbia: IEEE. doi:10.1109/TELSKS.2007.4376029

Guo, X., Li, Q.-y., & Xu, W.-l. (2017, February). Acceleration of the EM Algorithm Using the Vector Aitken Method and Its Steffensen Form. *Acta Mathematicae Applicatae Sinica, 33*(1), 175-182. doi:10.1007/s10255-017-0648-3

Han, J., & Kamber, M. (2006). *Data Mining: Concepts and Techniques* (2nd Edition ed.). (J. Gray, Ed.) San Francisco, CA, USA: Morgan Kaufmann Publishers, Elsevier.

Hardle, W., & Simar, L. (2013). *Applied Multivariate Statistical Analysis.* Berlin, Germany: Research Data Center, School of Business and Economics, Humboldt University.

Hofmann, T., & Puzicha, J. (1998). *Statistical Models for Co-occurrence Data.* Massachusetts Institute of Technology, Artificial Intelligence Laboratory. MIT Publisher. Retrieved from https://dspace.mit.edu/bitstream/handle/1721.1/7253/AIM-1625.pdf?sequence=2

Hofmann, T., Puzicha, J., & Jordan, M. I. (1998). Learning from Dyadic Data. In M. J. Kearns, S. A. Solla, & D. A. Cohn (Ed.), *Advances in Neural Information Processing Systems 11 (NIPS 1998).* *11*, pp. 466-472. Denver: MIT Press. Retrieved from https://papers.nips.cc/paper/1503-learning-from-dyadic-data

Jebara, T. (2015). *The Exponential Family of Distributions.* Columbia University, Computer Science Department. New York: Columbia Machine Learning Lab. Retrieved April 27, 2016, from http://www.cs.columbia.edu/~jebara/4771/tutorials/lecture12.pdf

Jia, Y.-B. (2013). *Lagrange Multipliers.* Lecture notes on course “Problem Solving Techniques for Applied Computer Science”, Iowa State University of Science and Technology, USA.

Josse, J., Jiang, W., Sportisse, A., & Robin, G. (2018). *Handling missing values.* Inria. Julie Josse. Retrieved October 12, 2020, from http://juliejosse.com/wp-content/uploads/2018/07/LectureNotesMissing.html

Lambers, J. (2009). *Accelerating Convergence.* University of Southern Mississippi, Department of Mathematics. Hattiesburg: University of Southern Mississippi. Retrieved February 15, 2018, from http://www.math.usm.edu/lambers/mat460/fall09/lecture13.pdf

Lansky, D., & Casella, G. (1992). Improving the EM Algorithm. *Computing Science and Statistics*, 420-424. doi:10.1007/978-1-4612-2856-1\_67

McLachlan, G., & Krishnan, T. (1997). *The EM Algorithm and Extensions.* New York, NY, USA: John Wiley & Sons. Retrieved from https://books.google.com.vn/books?id=NBawzaWoWa8C

Meng, X.-L., & Rubin, D. B. (1993, June 1). Maximum likelihood estimation via the ECM algorithm: A general framework. *Biometrika, 80*(2), 267-278. doi:10.2307/2337198

Montgomery, D. C., & Runger, G. C. (2003). *Applied Statistics and Probability for Engineers* (3rd Edition ed.). New York, NY, USA: John Wiley & Sons, Inc.

Nguyen, L. (2015). *Matrix Analysis and Calculus* (1st ed.). (C. Evans, Ed.) Hanoi, Vietnam: Lambert Academic Publishing. Retrieved from https://www.shuyuan.sg/store/gb/book/matrix-analysis-and-calculus/isbn/978-3-659-69400-4

Nguyen, L. (2020, November 2). Learning Dyadic Data and Predicting Unaccomplished Co-occurrent Values by Mixture Model. *Preprints*. doi:10.20944/preprints202011.0038.v1

Poli, R., Kennedy, J., & Blackwell, T. (2007, June). Particle swarm optimization. (M. Dorigo, Ed.) *Swarm Intelligence, 1*(1), 33-57. doi:10.1007/s11721-007-0002-0

Rabiner, L. R. (1989). A tutorial on hidden Markov models and selected applications in speech recognition. (J. H. Trussell, Ed.) *Proceedings of the IEEE, 77*(2), 257-286.

Rai, S. N., & Matthews, D. E. (1993, June). Improving the EM Algorithm. (C. A. McGilchrist, Ed.) *Biometrics, 49*(2), 587-591. doi:10.2307/2532570

Rajeswari, J., & Gunasundari, R. (2016, December). EMPWC: Expectation Maximization with Particle Swarm Optimization based Weighted Clustering for Outlier Detection in Large Scale Data. (C.-H. Lien, & T.-L. Liao, Eds.) *International Journal of Control Theory and Applications (IJCTA), 9*(36), 517-531. Retrieved February 21, 2018, from http://serialsjournals.com/articlesview.php?volumesno\_id=1131&article\_id=14367&volumes\_id=848&journals\_id=268

Ramage, D. (2007). *Hidden Markov Models Fundamentals.* Lecture Notes, Stanford University, USA.

Rao, R. C. (1955, June). Estimation and tests of significance in factor analysis. *Psychometrika, 20*(2), 93-111. doi:10.1007/BF02288983

Rosen, K. H. (2012). *Discrete Mathematics and Its Applications* (7nd Edition ed.). (M. Lange, Ed.) McGraw-Hill Companies.

Schmolze, J. G. (2001). *An Introduction to Hidden Markov Models.* Lecture Notes on course "COMP 232-Knowledge Based Systems", Department of Computer Science, Tufts University.

Sean, B. (2009). *The Expectation Maximization Algorithm - A short tutorial.* University of Notre Dame, Indiana, Department of Electrical Engineering. Sean Borman's Homepage.

StackExchange. (2013, November 19). *Eigenvalues of the product of 2 symmetric matrices*. (Stack Exchange Network) Retrieved February 9, 2018, from Mathematics StackExchange: https://math.stackexchange.com/questions/573583/eigenvalues-of-the-product-of-2-symmetric-matrices

Steorts, R. (2018). *The Multivariate Distributions: Normal and Wishart.* Duke University. Rebecca Steorts Homepage. Retrieved October 21, 2020, from http://www2.stat.duke.edu/~rcs46/lecturesModernBayes/601-module10-multivariate-normal/multivariate-normal.pdf

Ta, P. D. (2014). *Numerical Analysis Lecture Notes.* Vietnam Institute of Mathematics, Numerical Analysis and Scientific Computing. Hanoi: Vietnam Institute of Mathematics. Retrieved 2014

Tang, Z., Song, Y.-Q., & Liu, Z. (2014). Medical Image Clustering Based on Improved Particle Swarm Optimization and Expectation Maximization Algorithm. *The 6th Chinese Conference on Pattern Recognition (CCPR 2014).* *II*, pp. 360-371. Changsha, China: Springer. doi:10.1007/978-3-662-45643-9\_38

Tran, A.-K., Vo, Q.-N., & Lee, G. (2013). Maximum Likelihood Estimation of Gaussian Mixture Models Using PSO for Image Segmentation. In J. Chen, A. Cuzzocrea, & L. T. Yang (Ed.), *The 2013 IEEE 16th International Conference on Computational Science and Engineering (CSE 2013)* (pp. 501-507). Sydney, NSW, Australia: IEEE. doi:10.1109/CSE.2013.81

Wikipedia. (2014, August 4). *Karush–Kuhn–Tucker conditions*. (Wikimedia Foundation) Retrieved November 16, 2014, from Wikipedia website: http://en.wikipedia.org/wiki/Karush–Kuhn–Tucker\_conditions

Wikipedia. (2014, October 10). *Set (mathematics)*. (A. Rubin, Editor, & Wikimedia Foundation) Retrieved October 11, 2014, from Wikipedia website: http://en.wikipedia.org/wiki/Set\_(mathematics)

Wikipedia. (2016, March September). *Exponential family*. (Wikimedia Foundation) Retrieved 2015, from Wikipedia website: https://en.wikipedia.org/wiki/Exponential\_family

Wikipedia. (2017, May 25). *Aitken's delta-squared process*. (Wikimedia Foundation) Retrieved February 15, 2018, from Wikipedia website: https://en.wikipedia.org/wiki/Aitken%27s\_delta-squared\_process

Wikipedia. (2017, February 27). *Commuting matrices*. (Wikimedia Foundation) Retrieved February 9, 2018, from Wikipedia website: https://en.wikipedia.org/wiki/Commuting\_matrices

Wikipedia. (2017, November 27). *Diagonalizable matrix*. (Wikimedia Foundation) Retrieved February 10, 2018, from Wikipedia website: https://en.wikipedia.org/wiki/Diagonalizable\_matrix#Simultaneous\_diagonalization

Wikipedia. (2017, March 2). *Maximum a posteriori estimation*. (Wikimedia Foundation) Retrieved April 15, 2017, from Wikipedia website: https://en.wikipedia.org/wiki/Maximum\_a\_posteriori\_estimation

Wikipedia. (2017, March 7). *Particle swarm optimization*. (Wikimedia Foundation) Retrieved April 8, 2017, from Wikipedia website: https://en.wikipedia.org/wiki/Particle\_swarm\_optimization

Wikipedia. (2017, May 8). *Wolfe conditions*. (Wikimedia Foundation) Retrieved February 20, 2018, from Wikipedia website: https://en.wikipedia.org/wiki/Wolfe\_conditions

Wikipedia. (2018, January 15). *Conjugate prior*. (Wikimedia Foundation) Retrieved February 15, 2018, from Wikipedia website: https://en.wikipedia.org/wiki/Conjugate\_prior

Wikipedia. (2018, January 28). *Gradient descent*. (Wikimedia Foundation) Retrieved February 20, 2018, from Wikipedia website: https://en.wikipedia.org/wiki/Gradient\_descent

Wikipedia. (2018, February 17). *Lipschitz continuity*. (Wikimedia Foundation) Retrieved February 20, 2018, from Wikipedia website: https://en.wikipedia.org/wiki/Lipschitz\_continuity

Wikipedia. (2018, January 7). *Symmetry of second derivatives*. (Wikimedia Foundation) Retrieved February 10, 2018, from Wikipedia website: https://en.wikipedia.org/wiki/Symmetry\_of\_second\_derivatives

Wu, J. C. (1983, March). On the Convergence Properties of the EM Algorithm. *The Annals of Statistics, 11*(1), 95-103. Retrieved from https://projecteuclid.org/euclid.aos/1176346060

Zhang, B., Zhuang, L., Gao, L., Luo, W., Ran, Q., & Du, Q. (2014, May 14). PSO-EM: A Hyperspectral Unmixing Algorithm Based On Normal Compositional Model. (A. Plaza, Ed.) *IEEE Transactions on Geoscience and Remote Sensing, 52*(12), 7782 - 7792. doi:10.1109/TGRS.2014.2319337

Zivot, E. (2009). *Maximum Likelihood Estimation.* Lecture Notes on course "Econometric Theory I: Estimation and Inference (first quarter, second year PhD)", University of Washington, Seattle, Washington, USA.

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