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COINTEGRATION ANALYSIS WITH MIXED-FREQUENCY DATA

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COINTEGRATION ANALYSIS WITH MIXED-FREQUENCY DATA

Abstract

We develop a method for directly modeling cointegrated multivariate time series that are observed in mixed frequencies. We regard lower-frequency data as regularly (or irregularly) missing and treat them with higher-frequency data by adopting a state-space model. This utilizes the structure of multivariate data as well as the available sample information more fully than the methods of transformation to a single frequency, and enables us to estimate parameters including cointegrating vectors and the missing observations of low-frequency data and to construct forecasts for future values. For the maximum likelihood estimation of the parameters in the model, we use an expectation maximization algorithm based on the state-space representation of the error correction model. The statistical efficiency of the developed method is investigated through a Monte Carlo study. We apply the method to a mixed-frequency data set that consists of the quarterly real gross domestic product and the monthly consumer price index.

JEL Code: C13, C22, C32.

Keywords: missing data, Kalman filter, expectation maximization algorithm, forecasting, error correction model, smoothing, maximum likelihood estimation.

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1. Introduction

Multivariate time series that arise in economics and business are often observed in mixed frequencies. For example, data available from the database of the Bureau of Economic Analysis, the Bureau of Labor Statistics, and the Bureau of the Census are often in mixed frequencies, mostly with quarterly, monthly or weekly sampling intervals. Although cointegration, which represents a long-run equilibrium among the components of nonstationary multivariate time series, has been one of the most extensively investigated research topics, especially, in economics and business during the past two decades since Engle and Granger (1987), studies of cointegration have been limited to the case where all the components of a multivariate series are observed at the same frequency.

Data observed in mixed frequencies are usually transformed to a single frequency by temporally aggregating higher-frequency data to lower frequencies, or by interpolating lower-frequency data to higher frequencies. However, temporal aggregation destroys sample information (Zadrozny, 1990). Granger and Siklos (1995) examined the misinterpretation of the long-run component of a time series constructed by temporal aggregation. Marcellino (1999) theoretically derived the effects of temporal aggregation on cointegration such as the asymptotic invariance of cointegrating (CI) rank and vectors. However, he illustrated some possibilities of the loss of power of cointegration tests due to a decline in the number of available observations. Haug (2002) showed in Monte Carlo experiments with various cointegration tests and data generating processes (DGPs) that these power losses indeed occur and assessed their extents in samples of typical size used in empirical work.

For the case of interpolation of mixed-frequency data, Chow and Lin (1971, 1976) used the conventional regression approach. Other researchers, for example, Bernanke *et al.* (1997), Cuche and Hess (2000), and Liu and Hall (2001), used the state-space framework suggested by Harvey and Pierse (1984) where interpolation is based on univariate regression. When the goal is to estimate a multivariate model for forecasting or other purposes, this kind of interpolation is at best an intermediate nuisance and at worst a source of distortion in the data to be used for estimation (Zadrozny, 1990, p. 2).

Recently, Mariano and Murasawa (2003, 2004) considered multivariate models for constructing a new index of economic indicators using mixed-frequency data, which overcome the drawbacks of a univariate approach and exploit the cross-frequency sample information. However, they used differenced data instead of levels data. This causes a loss of information on the long-run dynamics among the variables.

In this paper we develop a method for “directly” modeling the cointegrated multivariate time series with mixed-frequency data, which is based on the state-space representation of the error correction model (ECM). We use the state-space formulation for mixed-frequency data in our development by fully utilizing the

structure of multivariate data as well as the available sample information. For the analysis, we exploit an expectation maximization (EM) algorithm. The method is applied to estimate the parameters and the missing observations of the low-frequency variables and to construct forecasts of their future values.

We introduce notation which is used in this paper: I_m denotes an $m \times m$ identity matrix; O_m , $O_{m \times n}$, and 0_m denote an $m \times m$ zero matrix, an $m \times n$ zero matrix, and an $m \times 1$ zero vector; $\text{vec}(\cdot)$ vectorizes a matrix columnwise from left to right, and $\text{vech}(\cdot)$ vectorizes the lower triangular part of a matrix columnwise; the symbol \otimes denotes the Kronecker product.

The structure of the paper is as follows. In Section 2, we describe preliminary concepts for mixed-frequency data and set up the state-space representation for the ECM of cointegration. In Section 3, we develop the procedures for estimating the parameters of the cointegrated model using the EM algorithm and discuss initial conditions for the procedures. In Section 4, we consider models with different types of deterministic terms. In Sections 5, we comment on smoothing, forecasting, and logarithmic transformation. In Section 6, we conduct Monte Carlo experiments for the investigation of the performance of the developed method. In Section 7, we consider a numerical example to illustrate the method and we conclude the paper in Section 8.

2. Model with mixed-frequency data

In this section, we define mixed-frequency data and the ECM for multivariate cointegrated time series.

2.1 Mixed-frequency data

We define high-frequency variables as those observed at the fundamental shortest sampling interval and define low-frequency variables as those observed at longer sampling intervals, either as temporal aggregates or as skip samples of their high-frequency values. A variable is skip sampled when, for example, it is generated every month but is sampled every third month, say, in the last month of every quarter. See, for example, Zdrozny (1988). We assume that the underlying data generating process of a multivariate time series of mixed-frequency data, composed of both the high and the low-frequency variables, operates at the highest frequency, as in Zdrozny (1990), Mittnik and Zdrozny (2004) and Mariano and Murasawa (2003, 2004). All variables are assumed to be produced at the highest frequency, but some variables are not observed at the highest frequency. For example, we consider a bivariate time series of the consumer price index (CPI) observed monthly and the gross domestic product (GDP) observed quarterly. CPI is the high-frequency variable and GDP is the low-frequency variable. The fundamental interval or highest frequency is monthly and GDP is in

principle ‘produced’ monthly but is observed only quarterly. Variables like GDP observed as temporal aggregates are often called flows, while variables observed at the fundamental frequency are often called stocks.

2.2 Error correction model and state-space representation

Let u_t be an n -dimensional vector autoregressive process of order p , $\text{VAR}(p)$, which operates at the highest frequency with CI rank h , and consider the corresponding error correction form

$$\Delta u_t = ab'u_{t-1} + \sum_{j=1}^{p-1} \Psi_j \Delta u_{t-j} + \varepsilon_t, \quad (1)$$

where a and b are $n \times h$ matrices with $0 < h < n$ and $h = n - d$, Ψ_j , for $j = 1, \dots, p-1$, are $n \times n$ matrices, and ε_t is an independent $N_n(0, \Omega)$ random vector. We assume that the elements of u_t are ordered such that the last $d = n - h$ elements are not cointegrated. This assumption permits the normalization $b = [I_h \ \beta_0]'$ for identification, as in Ahn and Reinsel (1990), where β_0 is a $d \times h$ matrix. The characteristic equation of model (1) has exactly d roots equal to one and all other roots are assumed to be outside the unit circle, so that u_t is cointegrated of order (1,1) (Engle and Granger, 1987).

We reorder the elements of u_t to form $z_t = (z'_{1t}, z'_{2t})'$ such that an $n_1 \times 1$ vector z_{1t} corresponds to high-frequency variables and an $n_2 \times 1$ vector z_{2t} corresponds to low-frequency variables, where $n = n_1 + n_2$. For brevity, we assume that the low-frequency variables are observed as temporal aggregates because the method can be easily modified to accommodate the alternative case of skip-sampled data. The reordering implies $z_t = Vu_t$, where V is an $n \times n$ permutation matrix. Substituting u_t with $V^{-1}z_t$ and pre-multiplying both sides of equation (1) by V , we can rewrite model (1) as

$$\Delta z_t = \alpha \beta' z_{t-1} + \sum_{j=1}^{p-1} \Gamma_j \Delta z_{t-j} + e_t, \quad (2)$$

where $\alpha = Va$, $\beta = Vb$, $\Gamma_j = V\Psi_j V'$, $e_t = V\varepsilon_t$, $e_t \sim N(0, \Sigma)$, and $\Sigma = V\Omega V'$. Also, note that $V^{-1} = V'$. For convenience of estimation, we write $\beta = V_1 I_h + V_2 \beta_0$, where V_1 and V_2 are $n \times h$ and $n \times d$ matrices, such that $V = [V_1 \ V_2]$. The $\text{VAR}(p)$ representation of model (2) is

$$z_t = \sum_{j=1}^p \Phi_j z_{t-j} + e_t, \quad (3)$$

where $\Phi_1 = I_n + \alpha \beta' + \Gamma_1$, $\Phi_j = \Gamma_j - \Gamma_{j-1}$, for $j = 2, \dots, p-1$, and $\Phi_p = -\Gamma_{p-1}$.

In practice, as mentioned above, the low-frequency variables, z_{2t} , are not observed directly at the highest frequency, but either as temporally-aggregated flows or as skip-sampled stocks. Flows can be expressed as $y_{2t} = C_0 z_{2t} + \xi_t$, where $\xi_t = \sum_{j=1}^v C_j z_{2t-j}$ is the lagged part in a temporal aggregation. Here, v denotes the maximum degree of aggregation and the C_j 's are $n_2 \times n_2$ diagonal indicator matrices with zeros and ones on the principal diagonal. By adjusting the diagonal elements of C_j , Zadrozny (1990) suggested a way to treat a variable which is observed directly as a stock, but with a delay.

We now construct a state-space representation of model (3). Let $\tilde{z}_t = (z'_t, \xi'_t)' = (z'_{1t}, z'_{2t}, \xi'_t)'$ be an $\tilde{n} \times 1$ vector with $\tilde{n} = n_1 + 2n_2$. We define the $s \times 1$ state vector $x_t = (\tilde{z}'_t, \dots, \tilde{z}'_{t-r+1})'$, where $r = \max(p, v)$ and $s = \tilde{n}r$ for $t = 1, \dots, T$. Then, we define the state equation, as in Zadrozny (1990), by

$$x_t = Fx_{t-1} + Ge_t, \quad (4)$$

where the initial state, x_0 , is assumed to be a normal random vector with mean vector λ and $s \times s$ covariance matrix Λ . Here, F and G denote $s \times s$ and $s \times n$ matrices defined by

$$F = \begin{bmatrix} \Phi_1^* & \Phi_2^* & \cdots & \Phi_r^* \\ I_{\tilde{n}} & O_{\tilde{n}} & \cdots & O_{\tilde{n}} \\ O_{\tilde{n}} & I_{\tilde{n}} & \cdots & O_{\tilde{n}} \\ \vdots & \vdots & \cdots & \vdots \\ O_{\tilde{n}} & O_{\tilde{n}} & \cdots & O_{\tilde{n}} \end{bmatrix}, \quad G = \begin{bmatrix} I_n \\ O_{(s-n) \times n} \end{bmatrix}, \quad \Phi_j^* = \begin{bmatrix} \Phi_j & O_{n \times n_2} \\ C_j^* & O_{n_2 \times n_2} \end{bmatrix}, \quad C_j^* = \begin{bmatrix} O_{n_2 \times n_1} & C_j \end{bmatrix},$$

$\Phi_j = O_n$, if $j > p$, and $C_j = O_{n_2}$ if $j > v$, so that Φ_j^* and C_j^* are $\tilde{n} \times \tilde{n}$ and $n_2 \times n$ matrices.

Next, assuming no observation errors, we define the measurement equation

$$y_t \equiv \begin{pmatrix} y_{1t} \\ y_{2t} \end{pmatrix} = \begin{pmatrix} H_1 \\ H_2 \end{pmatrix} x_t, \quad (5)$$

where $y_{1t} = z_{1t}$, $H_1 \equiv [I_{n_1} \ O_{n_1 \times (s-n_1)}]$, and $H_2 \equiv [O_{n_2 \times n_1} \ C_0 \ I_{n_2} \ O_{n_2 \times (s-\tilde{n})}]$. The matrix H_1 picks out the high-frequency variables from state vector x_t and the matrix H_2 picks out the low-frequency variables from the state vector. We introduce a new series, y_{2t}^+ , which is observed only at lower frequencies, in order to deal with the missing observations in y_{2t} . As in Brockwell and Davis (1991), we fill in missing observations of y_{2t} with random vectors which are independent of y_t and are distributed independently of the parameters in model (2). Accordingly, we modify measurement equation (5) as

$$y_t^+ \equiv \begin{pmatrix} y_{1t} \\ y_{2t}^+ \end{pmatrix} = \begin{pmatrix} H_1 \\ H_{2t} \end{pmatrix} x_t + \begin{pmatrix} O_{n_1 \times n_2} \\ Q_{2t} \end{pmatrix} w_t, \quad (6)$$

$$H_{2t} = \begin{cases} H_2 & \text{if } y_{2t} \text{ is observable} \\ O_{n_2 \times s} & \text{otherwise} \end{cases}, \quad Q_{2t} = \begin{cases} O_{n_2} & \text{if } y_{2t} \text{ is observable} \\ I_{n_2} & \text{otherwise} \end{cases},$$

and w_t is an independent random vector distributed $N_{n_2}(0, I_{n_2})$. We also define $H_t = [H_1', H_{2t}']'$ and $Q_t = [O_{n_1 \times n_2}', Q_{2t}']'$, which will appear in equations (16) and (20) and in Appendix B. In the implementation, because the realization of w_t is independent of y_t , setting $w_t = 0$ is the preferred simple choice (Brockwell and Davis, 1991; Mariano and Murasawa, 2003). Instead of using w_t , a selection matrix may be used for constructing a measurement equation in order to control the mixed-frequency data, as in Zdrozny (1990).

We note that when the low-frequency variables are stocks, with missing data attributable to skip-sampling, we do not need to define y_{2t} , ξ_t , C_j , and \tilde{z}_t . Instead, we redefine $\tilde{n} = n = n_1 + n_2$, $r = \max(p, 2)$, $\Phi_j^* = \Phi_j$, and $H_2 \equiv [O_{n_2 \times n_1} \quad I_{n_2} \quad O_{n_2 \times (s-n)}]$. We need $r = 2$ in order to construct the state equation even though the autoregressive order is one. Similar adjustments can be applied to more complicated cases, where the low-frequency variables are observed as both stocks and flows.

3. Maximum likelihood estimation of parameters

In this section, we consider maximum likelihood estimation (MLE) of parameters in error correction model (2) in the state-space form (4) and (6).

3.1 EM algorithm

Dempster *et al.* (1977), Shumway and Stoffer (1982), and Watson and Engle (1983) developed and illustrated the EM algorithm for estimating a model in state-space form, when some variables are partly or completely unobserved (latent).

Let $X_s = (x_t; 0 \leq t \leq s)$ and $Y_s^+ = (y_t^+; 1 \leq t \leq s)$ be information sets. In order to develop an EM algorithm for estimating the parameters of the state-space model in (4) and (6), we consider several transformations for forming the likelihood function with respect to the complete data X_T and Y_T^+ . Let

$$\Gamma = [\alpha \quad \Gamma_1 \quad \cdots \quad \Gamma_{p-1}], \quad \Gamma^* = [\alpha V_1' \quad \Gamma_1 \quad \cdots \quad \Gamma_{p-1}],$$

$$A = [I_n \quad O_{n \times n_2} \quad -I_n \quad O_{n \times (s-2n-n_2)}], \quad \text{and} \quad D = [V_2' \quad O_{d \times (s-n)}],$$

where $Dx_{t-1} = V_2' z_{t-1}$. Define

$$B = \begin{bmatrix} \beta' & O_{h \times n_2} & O_{h \times n} & O_{h \times n_2} & O_{h \times n} & \cdots \\ I_n & O_{n \times n_2} & -I_n & O_{n \times n_2} & O_n & \cdots \\ O_n & O_{n \times n_2} & I_n & O_{n \times n_2} & -I_n & \cdots \\ O_n & O_{n \times n_2} & O_n & O_{n \times n_2} & I_n & \cdots \\ \vdots & \vdots & \vdots & \vdots & \vdots & \cdots \end{bmatrix},$$

$$B^* = \begin{bmatrix} I_n & O_{n \times n_2} & O_n & O_{n \times n_2} & O_n & \cdots \\ I_n & O_{n \times n_2} & -I_n & O_{n \times n_2} & O_n & \cdots \\ O_n & O_{n \times n_2} & I_n & O_{n \times n_2} & -I_n & \cdots \\ O_n & O_{n \times n_2} & O_n & O_{n \times n_2} & I_n & \cdots \\ \vdots & \vdots & \vdots & \vdots & \vdots & \cdots \end{bmatrix},$$

where $Bx_{t-1} = (z_{t-1}'\beta, \Delta z_{t-1}', \dots, \Delta z_{t-p+1}')'$, $B^* x_{t-1} = (z_{t-1}', \Delta z_{t-1}', \dots, \Delta z_{t-p+1}')'$, and B and B^* are $\{h + (p-1)n\} \times s$ and $np \times s$ matrices. We consider two transformations for forming the likelihood function and estimating Γ , Σ , and β_0 .

First,

$$Ax_t = \Delta z_t = \alpha \beta' z_{t-1} + \Gamma_1 \Delta z_{t-1} + \cdots + \Gamma_{p-1} \Delta z_{t-p+1} + e_t = \Gamma Bx_{t-1} + e_t, \quad (7)$$

$$\begin{aligned} Ax_t &= \alpha \beta_0' V_2' z_{t-1} + \alpha V_1' z_{t-1} + \Gamma_1 \Delta z_{t-1} + \cdots + \Gamma_{p-1} \Delta z_{t-p+1} + e_t \\ &= \alpha \beta_0' Dx_{t-1} + \Gamma^* B^* x_{t-1} + e_t \\ &= \{\alpha \otimes (Dx_{t-1})'\} \text{vec}(\beta_0) + \Gamma^* B^* x_{t-1} + e_t, \end{aligned} \quad (8)$$

because $\alpha \beta_0' Dx_{t-1} = \text{vec}\{\alpha \beta_0' Dx_{t-1}\}$ and $\text{vec}\{(Dx_{t-1})' \beta_0 \alpha'\} = \{\alpha \otimes (Dx_{t-1})'\} \text{vec}(\beta_0)$, which are obtained using the vectorization rule $\text{vec}(ABC) = (C' \otimes A) \text{vec}(B)$ (Magnus and Neudecker, 1988).

Then, we express the log-likelihood function as

$$\begin{aligned} \log L(\theta; X_T, Y_T^+) &= -\frac{1}{2} \log |\Lambda| - \frac{1}{2} (x_0 - \lambda)' \Lambda^{-1} (x_0 - \lambda) \\ &\quad - \frac{T}{2} \log |\Sigma| - \frac{1}{2} \sum_{t=1}^T (Ax_t - \Gamma Bx_{t-1})' \Sigma^{-1} (Ax_t - \Gamma Bx_{t-1}), \end{aligned} \quad (9)$$

or as

$$\begin{aligned} \log L(\theta; X_T, Y_T^+) &= -\frac{1}{2} \log |\Lambda| - \frac{1}{2} (x_0 - \lambda)' \Lambda^{-1} (x_0 - \lambda) \\ &\quad - \frac{T}{2} \log |\Sigma| - \frac{1}{2} \sum_{t=1}^T [Ax_t - \Gamma^* B^* x_{t-1} - \{\alpha \otimes (Dx_{t-1})'\} \text{vec}(\beta_0)]' \times \end{aligned}$$

$$\Sigma^{-1}[Ax_t - \Gamma^* B^* x_{t-1} - \{\alpha \otimes (Dx_{t-1})'\} \text{vec}(\beta_0)], \quad (10)$$

where terms that do not contain the parameters, $\theta \equiv (\text{vec}(\beta_0)', \text{vec}(\Gamma)', \text{vech}(\Sigma)')$, are omitted. Version (9) of the log-likelihood function is used to estimate the ‘‘stationary’’ parameters in Γ and Σ . Version (10) of the log-likelihood function is used to estimate the remaining ‘‘nonstationary’’ parameters in β_0 . The distribution of w_t in the measurement equation does not have any effect on equations (9) and (10). Because the log likelihood function depends on the unobserved information, X_T , the EM algorithm is applicable, conditional on the observed information, Y_T^+ . Specifically, we define the estimated parameters at iteration $l+1$ as the value of θ which maximizes

$$Q(\theta | \theta^{(l)}) = E_l \left\{ \log L(\theta; X_T, Y_T^+) | Y_T^+ \right\}, \quad (11)$$

where $\theta^{(l)}$ denotes estimated θ after l iterations and $E_l\{\cdot | Y_T^+\}$ denotes the conditional expectation with respect to a density containing $\theta^{(l)}$, given Y_T^+ . Then, using the derivatives of (11) with respect to θ stated in Appendix A, we obtain the following equations for updating $\theta^{(l+1)}$ at the end of l iterations,

$$\begin{aligned} \text{vec}(\beta_0^{(l+1)}) &= \left\{ \alpha^{(l)'} \Sigma^{(l)-1} \alpha^{(l)} \otimes (DM_{11}D') \right\}^{-1} \times \\ &\quad \text{vec} \left[D \left\{ M_{10}A' - M_{11}(\Gamma^{*(l)}B^*)' \right\} \Sigma^{(l)-1} \alpha^{(l)} \right], \end{aligned} \quad (12)$$

$$\Gamma^{(l+1)} = (AM_{01}B'^{(l+1)}) \left(B^{(l+1)}M_{11}B'^{(l+1)} \right)^{-1}, \quad (13)$$

$$\Sigma^{(l+1)} = T^{-1} \left(AM_{00}A' - \Gamma^{(l+1)}B^{(l+1)}M_{10}A' \right), \quad (14)$$

where, in equation (12), $\alpha^{(l)}$, $\Sigma^{(l)}$, and $\Gamma^{*(l)}$ are given by the previous iteration; in equation (13), $B^{(l+1)}$ is given by $\beta_0^{(l+1)}$, according to equation (12); and, in equation (14), $\Gamma^{(l+1)}$ is given by equation (13). In equations (12) to (14), M_{00} , M_{01} , M_{10} , and M_{11} are given by

$$M_{jk} = \sum_{t=1}^T E_l(x_{t-j}x'_{t-k} | Y_T^+) = \sum_{t=1}^T (P_{t-j,t-k}^T + x_{t-j}^T x_{t-k}^T), \quad (15)$$

for $j, k = 0, 1$, where $P_{t-j,t-k}^T$ and x_{t-j}^T are produced by the prediction and updating recursions of the Kalman filter. In Appendix B, we adopt the fixed-interval smoothing algorithm of De Jong (1989), which avoids inversion of large matrices, and, hence, is computationally more efficient than the classical smoothing equations (Durbin and Koopman, 2001). We note that the mean vector, λ , and the covariance matrix, Λ , of the initial state vector, x_0 , cannot be estimated simultaneously. Following Shumway and Stoffer (1982, 2000), we

preset the covariance matrix and estimate the mean vector as $\lambda^{(l+1)} = x_0^T$ by maximizing (11).

We summarize the iterative EM procedure as follows:

- (1) Calculate M_{jk} for $j, k = 0, 1$ using equations (B1) to (B6) in Appendix B, with the initial values $\lambda^{(0)}$, Λ , and $\theta^{(0)}$.
- (2) Estimate $\lambda^{(1)} = x_0^T$ and calculate $\theta^{(1)}$ using equations (12) to (14).
- (3) Iterate on steps (1) and (2) above until the parameter estimates or the likelihood values converge. At each iteration, we calculate the innovations form of the log-likelihood function (Schweppe, 1965),

$$\begin{aligned} \log L(\theta; Y_T^+) &= -\frac{1}{2} \sum_{t=1}^T \log |H_t P_t^{t-1} H_t' + Q_t Q_t'| \\ &\quad - \frac{1}{2} \sum_{t=1}^T (y_t - H_t x_t^{t-1})' (H_t P_t^{t-1} H_t' + Q_t Q_t')^{-1} (y_t - H_t x_t^{t-1}), \end{aligned} \quad (16)$$

and stop when the difference between $\log L(\theta^{(l+1)}; Y_T^+)$ and $\log L(\theta^{(l)}; Y_T^+)$ is less than a predetermined small value.

As an alternative to the EM algorithm, the Newton-Raphson (NR) method can be applied to maximize the likelihood. However, the NR method is more likely to fail because it is very sensitive to initial values; see Shumway and Stoffer (2000) for more comparisons between the NR and EM algorithms. Thus, we recommend using the EM algorithm especially because in cointegration with mixed-frequency data initial values of parameters are difficult to obtain.

3.2 Initialization

To start the EM algorithm with the Kalman filter, we need to specify the initial values, $\lambda^{(0)}$, Λ , and $\theta^{(0)}$. When the state equation is nonstationary, the unconditional distribution of the state vector is not defined. Usually, the initial distribution of x_0 must be specified by a diffuse or noninformative prior because genuine prior information is generally not available (Harvey, 1991). Therefore, we set $\lambda^{(0)} = 0$ and $\Lambda = \delta I$, where δ is a large value, for example, $\delta = 10^8$.

Regarding initial values of parameters, we first obtain an estimate of the nonstationary parameters by using single-frequency data that are usually obtained by transforming the high-frequency variables, y_{1t} , to match the low frequency of y_{2t} . Then, we use the relationship between the nonstationary parameters in mixed-frequency non-temporally-aggregated and single-frequency temporally-aggregated models, estimated using mixed- and single-frequency data, in order to

obtain an initial estimate of the nonstationary parameters in $\beta_0^{(0)}$. For example, see Marcellino (1999) and Pons and Sansó (2005) for the explicit formulas for the relationship in several cases. For the remaining stationary parameters, however, it is not easy to obtain an explicit formula to describe the relationship between the parameters of models of mixed- and single-frequency data. Therefore, treating the initial estimate $\beta_0^{(0)}$ as known, and, thus, fixed in equation (2), we obtain initial estimates of the stationary parameters, $\Gamma^{(0)}$ and $\Sigma^{(0)}$, using the iterative EM algorithm. Henceforth, we call this the “initial EM” algorithm. We use the output from the initial EM algorithm to produce initial values for the “main EM” algorithm which we are advocating. Accordingly, we obtain a better starting value $\lambda^{(0)}$ for the main EM algorithm than the starting value $\lambda^{(0)} = 0$ used for the initial EM algorithm.

In order to prespecify the CI rank in the analysis of cointegration with mixed-frequency data, we use the fact stated by Marcellino (1999) that the CI rank is invariant to temporal aggregation. Then, we can use the CI rank, obtained by applying the CI rank test to the temporally-aggregated single-frequency data.

We select as “best” the VAR(p) model whose MLE yields the lowest values of Akaike’s information criterion (AIC) and Schwartz’s Bayesian information criterion (SBC),

$$\text{AIC} = T^{-1} \{-2 \log L(\hat{\theta}; Y_T^+) + 2 \dim(\theta)\},$$

$$\text{SBC} = T^{-1} \{-2 \log L(\hat{\theta}; Y_T^+) + \dim(\theta) \log T\},$$

where $\log L(\hat{\theta}; Y_T^+)$ is given by equation (16), $\hat{\theta}$ is the MLE of θ , obtained using the main EM algorithm, and $\dim(\theta)$ is the dimension of θ .

4. Model with deterministic terms

In this section, we describe a state-space representation that can accommodate constant and linear trend terms in model (1). To this end, we consider the model

$$\Delta u_t = \gamma_1 + \gamma_2(t-1) + ab'u_{t-1} + \sum_{j=1}^{p-1} \Psi_j \Delta u_{t-j} + \varepsilon_t, \quad (17)$$

where γ_1 and γ_2 are $n \times 1$ vectors. If γ_1 and γ_2 are restricted, they can be reparameterized as $\gamma_j = a\rho_j$ and incorporated into the cointegration relationship as

$$a\rho_1' + a\rho_2'(t-1) + ab'u_{t-1} = a \begin{bmatrix} \rho_1' & \rho_2' & b' \end{bmatrix} \begin{pmatrix} 1 \\ t-1 \\ u_{t-1} \end{pmatrix}.$$

For brevity of exposition, we assume γ_1 and γ_2 are all unrestricted in this section. For cases with restrictions, see Appendix C, where we consider state-space representations for various combinations of deterministic terms, as in Johansen (1996). Then, model (2) is rewritten as

$$\Delta z_t = \mu_1 + \mu_2(t-1) + \alpha\beta'z_{t-1} + \sum_{j=1}^{p-1} \Gamma_j \Delta z_{t-j} + e_t, \quad (18)$$

where $\mu_1 = V\gamma_1$ and $\mu_2 = V\gamma_2$. For model (18), we obtain a state-space representation by replacing state vector x_t with $\tilde{x}_t = (1, t, x_t)'$, so that state equation (4) becomes

$$\tilde{x}_t = \tilde{F}\tilde{x}_{t-1} + \tilde{G}e_t, \quad (19)$$

where

$$\tilde{F} = \begin{bmatrix} 1 & 0 & & \\ 1 & 1 & & \\ \mu_1 & \mu_2 & & \\ 0_{s-n} & 0_{s-n} & & \end{bmatrix} \begin{array}{c} \vdots \\ \vdots \\ \vdots \\ F \end{array}, \quad \tilde{G} = \begin{bmatrix} O_{2 \times n} \\ G \end{bmatrix},$$

and, correspondingly, measurement equation (6) becomes

$$y_t^+ = \tilde{H}_t \tilde{x}_t + Q_t w_t, \quad (20)$$

where $\tilde{H}_t = [O_{n \times 2} \quad H_t]$.

Using state-space representation (19), we can similarly get transformation (7) or (8). First, we define matrices

$$\tilde{A} = [O_{n \times 2} \quad A], \quad \tilde{\Gamma} = [\mu_1 \quad \mu_2 \quad \Gamma], \quad \tilde{\Gamma}^* = [\mu_1 \quad \mu_2 \quad \Gamma^*],$$

$$B_\rho = \begin{bmatrix} 0_h & 0_h & \\ \vdots & \vdots & \\ O_{(p-1)n \times 2} & & B \end{bmatrix}, \quad \tilde{B} = \begin{bmatrix} I_2 & O_{2 \times s} \\ \vdots & \vdots \\ B_\rho \end{bmatrix}.$$

Then, we obtain

$$\tilde{A}\tilde{x}_t = \mu_1 + \mu_2(t-1) + \Gamma B_\rho \tilde{x}_{t-1} + e_t = \tilde{\Gamma} \tilde{B} \tilde{x}_{t-1} + e_t, \quad (21)$$

$$\tilde{A}\tilde{x}_t = \mu_1 + \mu_2(t-1) + \alpha\beta'_0 D x_{t-1} + \Gamma^* B^* x_{t-1} + e_t = \alpha\beta'_0 \tilde{D} \tilde{x}_{t-1} + \tilde{\Gamma}^* \tilde{B}^* \tilde{x}_{t-1} + e_t, \quad (22)$$

where

$$\tilde{B}^* = \begin{bmatrix} I_2 & O_{2 \times s} \\ \vdots & \vdots \\ O_{np \times 2} & B^* \end{bmatrix}, \quad \tilde{D} = [O_{d \times 2} \quad D].$$

Similarly to the case with no deterministic terms, by using equations (21) or (22), we can obtain estimated parameters at iteration $l+1$ of the EM algorithm by using derivatives of conditional expectation (11), namely

$$\text{vec}(\beta_0^{(l+1)}) = \left\{ \alpha'^{(l)} \Sigma^{(l)-1} \alpha^{(l)} \right\} \otimes \left(\tilde{D} \tilde{M}_{11} \tilde{D}' \right)^{-1} \times \text{vec} \left[\tilde{D} \left\{ \tilde{M}_{10} \tilde{A}' - \tilde{M}_{11} \left(\tilde{\Gamma}^{*(l)} \tilde{B}^* \right)' \right\} \Sigma^{(l)-1} \alpha^{(l)} \right], \quad (23)$$

$$\tilde{\Gamma}^{(l+1)} = \left(\tilde{A} \tilde{M}_{01} \tilde{B}'^{(l+1)} \right) \left(\tilde{B}^{(l+1)} \tilde{M}_{11} \tilde{B}'^{(l+1)} \right)^{-1}, \quad (24)$$

$$\Sigma^{(l+1)} = T^{-1} \left(\tilde{A} \tilde{M}_{00} \tilde{A}' - \tilde{\Gamma}^{(l+1)} \tilde{B}^{(l+1)} \tilde{M}_{10} \tilde{A}' \right), \quad (25)$$

where for $j, k = 0, 1$,

$$\tilde{M}_{jk} = \sum_{t=1}^T \text{E}_t \left(\tilde{x}_{t-j} \tilde{x}_{t-k}' \mid Y_T^+ \right) = \sum_{t=1}^T \left(\tilde{P}_{t-j, t-k}^T + \tilde{x}_{t-j}^T \tilde{x}_{t-k}^T \right), \quad (26)$$

and $\tilde{P}_{t-j, t-k}^T$ and \tilde{x}_{t-j}^T are obtained using the Kalman filter, as in the case with no deterministic terms.

5. Comments

One of the important purposes of the paper is to estimate missing or unobserved low-frequency variables, y_{2t} , which satisfy the structure of the cointegrated multivariate time-series model. This can be done by estimating x_t^T , together with its covariance matrix P_t^T , using smoothing equations (B5) and (B6) in Appendix B. One of the advantages of the proposed method is that we can use it to forecast low-frequency variables which are generated jointly with variables observed at higher frequencies. For example, with quarterly GDP and other monthly variables, we can forecast monthly GDP even if GDP is observed only quarterly.

Usually, logarithms of variables are taken before fitting a multivariate time-series model, especially to stabilize the variances of series. This creates no difficulties for high-frequency series (Harvey and Pierce, 1984). However, for a temporally-aggregated low-frequency series, although the sum of the original variables is observed, the logarithm of a sum is not equal to the sum of the logarithms. In such a case, we assume that the logarithms of the data are integrated of order 1. Furthermore, we assume that temporally-aggregated variables are geometric means of unobserved high-frequency variables. This is not a common accounting identity that links high- and low-frequency variables. Usually, low-frequency values are arithmetic means of high-frequency values. In this case, the low-frequency values are geometric means of the high-frequency values. Otherwise, we would have to work with a nonlinear state-space model,

using the extended Kalman filter, as in Anderson and Moore (1979, pp. 193-195). See also Harvey and Pierce (1984) and Mariano and Murasawa (2003).

6. Monte Carlo experiments

Monte Carlo experiments are conducted to investigate the performance of the proposed method. The data generating process we consider is similar to the one in Ahn and Reinsel (1990), except that we consider a 3-dimensional process, specifically,

$$\Delta u_t = (\Delta u_{1t}, \Delta u_{2t}, \Delta u_{3t})' = \gamma_1 + ab'u_{t-1} + \varepsilon_t,$$

where $\varepsilon_t \sim \text{i.i.d. } N_3(0, \Omega)$, for $t = 1, \dots, T$, and γ_1 denotes an unrestricted constant term. The parameters are set at the following values:

$$\begin{aligned} \gamma_1 = (\gamma_{11}, \gamma_{21}, \gamma_{31})' &= (-0.2, 0.1, 0.3)', \quad a = (a_1, a_2, a_3)' = (0.6, 1, 0.4)', \\ b = (b_1, b_2, b_3)' &= (1, -2, 3)' \quad \text{and} \quad \Omega = \begin{pmatrix} 25 & 7.5 & 2.5 \\ 7.5 & 9 & 1.5 \\ 2.5 & 1.5 & 1 \end{pmatrix}, \end{aligned}$$

so that u_t is cointegrated with rank one.

For convenience, we assume that permutation matrix V is an identity matrix, that is, u_{2t} and u_{3t} are not cointegrated. Then, $u_t = z_t$ and the parameters in models (17) and (18) are identical. After generating u_t , we set $y_{1t} = u_{1t}$ and $y_{2t} = u_{2t}$, making them high-frequency variables, and, then, generate the low-frequency variable, y_{3t} , in the following two ways. First (CASE I), we set $y_{3t} = u_{3t} + u_{3t-1} + u_{3t-2}$, for $t = 3, 6, 9, \dots$, which makes y_{3t} a temporally-aggregated flow variable. Second (CASE II), we set $y_{3t} = u_{3t}$, for $t = 3, 6, 9, \dots$, which makes y_{3t} a skip-sampled stock variable. The remaining values of y_{3t} , for $t \neq 3, 6, 9, \dots$, are assumed to be unobserved.

We generate 1,000 replications of the series for sample sizes $T = 120$ and $T = 240$, which represent 10 and 20 years of monthly data, such that the first 50 values are discarded in order to reduce dependence on starting values. We estimate a VAR(1) with an unrestricted constant.

Because, to our best knowledge, there is no other method available to analyze mixed-frequency data with cointegration, we evaluate the performance of the proposed method against the case in which all variables are observed at the highest frequency, that is, there are no missing data. We expect our method to perform nearly as well with intermittently-missing mixed-frequency data as with complete high-frequency data. The same issue arose in Chen and Zadrozny (1998), where a similar Monte Carlo experiment was conducted to evaluate the

performance of their method for estimating a stationary VAR model using mixed-frequency data, relative to using complete high-frequency data.

The method of Ahn and Reinsel (1990) is used for estimating with the complete high-frequency data. As mentioned before, the initial estimate of the cointegrating vector for the proposed method is obtained using low-frequency data obtained by temporally-aggregated high-frequency data. The remaining initial estimates of the stationary parameters are obtained by treating the initial estimate of the cointegrating vector as if it were known.

Tables 1 and 2 contain the results of the simulations for CASEs I and II. We observe that the differences in the tables are generally fairly small and that the main EM method performs well with the mixed-frequency data, compared with the complete data, in terms of biases and root mean-squared errors (RMSEs) of estimated parameters. As expected, neither the initial nor the main EM algorithms perform better using mixed-frequency data than using complete high-frequency data.

It is interesting to compare tables 1 and 2 with tables 2 to 7 in Chen and Zdrozny (1998), which show how MLE and extended Yule-Walker (XYW) parameter estimates deteriorate, in terms of RMSE, when going from complete to mixed-frequency data. In their table 5, the RMSEs of MLE increase about 44%, whereas the best RMSEs of XYW increase about 77% or more. Consider how the RMSEs change in table 1, for $T = 240$, as we move from using complete data to using mixed-frequency data and the main EM algorithm. The RMSEs increase from 5.7% to 65.1%, on average by 40.3%. We conclude that MLEs of parameters in both stationary and nonstationary (cointegrated) VAR processes lose a similar amount of RMSE accuracy when going from complete to mixed-frequency data. However, we should be cautious in drawing this as a general conclusion because it surely also depends on whether the VAR process is stationary or not, on the dimension of the process, and on the sample size.

In all cases, regardless of the value of T , the main EM algorithm performs better than the initial EM, in terms of bias and RMSE. For longer series, biases and RMSEs of both the initial and main algorithm are smaller. In table 3, we report decline rates of RMSEs for estimated parameters in CASEs I and II, when T doubles from 120 to 240. For stationary parameters, a and Ω_{ij} , and nonstationary parameters, b , RMSEs in tables 1 and 2 are consistent with the respective convergence rates of $O_p(T^{-1/2})$ and $O_p(T^{-1})$. RMSEs of stationary parameters generally decline by 29% ($=1 - \sqrt{120/240}$) or more when T doubles from 120 to 240 and RMSEs of nonstationary parameters generally decline by 50% ($=1 - 120/240$) or more when T doubles. However, changes in the RMSEs of stationary parameters, γ , are ambiguous, because their RMSEs decline faster than is predicted for stationary parameters but slower than is predicted for nonstationary parameters. Most importantly, this also occurs when we apply the method of Johansen (1988) or Ahn and Reinsel (1990) to the analysis of complete data and the RMSEs of the stationary parameters, γ , do not decline more than

those of the nonstationary parameters.

Figure 1 shows RMSEs of estimates of missing values of y_{3t} , for $t \neq 3, 6, 9, \dots$, for CASEs I and II. The figure shows all except two starting values for CASE I and one starting value for CASE II. Two final outlying values (0.588 and 0.693, for $t=118$ and 119 , when $T=120$, and 0.556 and 0.626, for $t=238$ and 239 , when $T=240$) are omitted from Figure 1, because they are spuriously large due to endpoint effects, which arise as a result of the backward smoothing recursions (B3) to (B6), with $r_{T+1}=0$ and $R_{T+1}=0$ being a sort of initialization. In Figure 2, this phenomenon is weaker because the aggregation order of the variable is regarded as one. From Figures 1 and 2, we see that as T doubles from 120 to 240, the RMSEs decline by 8.87% and 9.05% in CASEs I and II. Therefore, for stationary parameters, RMSEs of estimates of missing values decline more slowly than the expected convergence rate $O_p(T^{-1/2})$.

Table 1. Performance of EM algorithm in complete versus mixed-frequency data, for CASE I of temporally-aggregated mixed-frequency data

	$T = 120$			$T = 240$		
	Complete data	Mixed-freq. data		Complete data	Mixed-freq. data	
		Init. EM	Main EM		Init. EM	Main EM
$a_1 = 0.6$						
Mean	0.585	0.566	0.578	0.593	0.583	0.588
RMSE	0.049	0.069	0.056	0.035	0.042	0.037
$a_2 = 1$						
Mean	1.007	0.987	0.995	1.002	0.987	0.993
RMSE	0.032	0.064	0.040	0.020	0.036	0.025
$a_3 = 0.4$						
Mean	0.399	0.391	0.396	0.399	0.394	0.397
RMSE	0.011	0.028	0.016	0.007	0.015	0.010
$b_1 = -2$						
Mean	-1.999	-1.949	-1.990	-2.001	-2.003	-2.001
RMSE	0.030	0.171	0.050	0.011	0.058	0.015
$b_2 = 3$						
Mean	2.997	2.896	2.980	3.002	3.009	3.003
RMSE	0.066	0.380	0.112	0.023	0.126	0.033
$\gamma_{11} = -0.2$						
Mean	-0.148	0.013	-0.199	-0.203	-0.473	-0.276
RMSE	0.860	3.568	1.650	0.466	1.669	0.665
$\gamma_{21} = 0.1$						
Mean	0.114	0.362	0.019	0.071	-0.386	-0.052
RMSE	1.229	6.186	2.652	0.590	2.743	0.953
$\gamma_{31} = 0.3$						
Mean	0.314	0.418	0.271	0.294	0.107	0.240
RMSE	0.480	2.457	1.065	0.232	1.103	0.383

Table 1. (Continued)

	$T = 120$			$T = 240$		
	Complete data	Mixed-freq. data		Complete data	Mixed-freq. data	
		Init. EM	Main EM		Init. EM	Main EM
$\Omega_{11} = 25$						
Mean	24.470	27.767	25.891	24.843	26.186	25.483
RMSE	3.114	5.163	3.874	2.240	2.894	2.537
$\Omega_{12} = 7.5$						
Mean	7.374	12.365	9.816	7.469	9.644	8.642
RMSE	1.508	6.390	3.999	1.074	3.066	2.226
$\Omega_{13} = 2.5$						
Mean	2.461	5.209	3.773	2.486	3.690	3.090
RMSE	0.511	3.504	2.175	0.352	1.676	1.173
$\Omega_{22} = 9$						
Mean	8.806	16.121	12.950	8.931	12.434	11.063
RMSE	1.177	8.899	5.957	0.863	4.500	3.352
$\Omega_{23} = 1.5$						
Mean	1.461	5.266	3.590	1.474	3.345	2.572
RMSE	0.307	4.575	3.034	0.216	2.343	1.692
$\Omega_{33} = 1$						
Mean	0.976	2.782	1.931	0.984	1.853	1.461
RMSE	0.132	2.188	1.400	0.092	1.095	0.758

Table 2. Performance of EM algorithm in complete versus mixed-frequency data, for CASE II of skip-sampled mixed-frequency data

	$T = 120$			$T = 240$		
	Complete data	Mixed-freq. data		Complete data	Mixed-freq. data	
		Init. EM	Main EM		Init. EM	Main EM
$a_1 = 0.6$						
Mean	0.585	0.576	0.576	0.593	0.587	0.587
RMSE	0.049	0.060	0.058	0.035	0.039	0.038
$a_2 = 1$						
Mean	1.007	0.989	0.989	1.002	0.990	0.991
RMSE	0.032	0.054	0.047	0.020	0.032	0.029
$a_3 = 0.4$						
Mean	0.399	0.394	0.394	0.399	0.396	0.396
RMSE	0.011	0.022	0.020	0.007	0.013	0.011
$b_1 = -2$						
Mean	-1.999	-1.988	-1.993	-2.001	-2.000	-2.001
RMSE	0.030	0.087	0.055	0.011	0.034	0.019
$b_2 = 3$						
Mean	2.997	2.972	2.984	3.002	3.001	3.002
RMSE	0.066	0.194	0.121	0.023	0.074	0.041
$\gamma_{11} = -0.2$						
Mean	-0.148	0.026	-0.104	-0.203	-0.206	-0.238
RMSE	0.860	2.140	1.631	0.466	1.014	0.704
$\gamma_{21} = 0.1$						
Mean	0.114	0.420	0.196	0.071	0.070	0.014
RMSE	1.229	3.566	2.672	0.590	1.626	1.069
$\gamma_{31} = 0.3$						
Mean	0.314	0.431	0.341	0.294	0.289	0.267
RMSE	0.480	1.417	1.066	0.232	0.654	0.429

Table 2. (Continued)

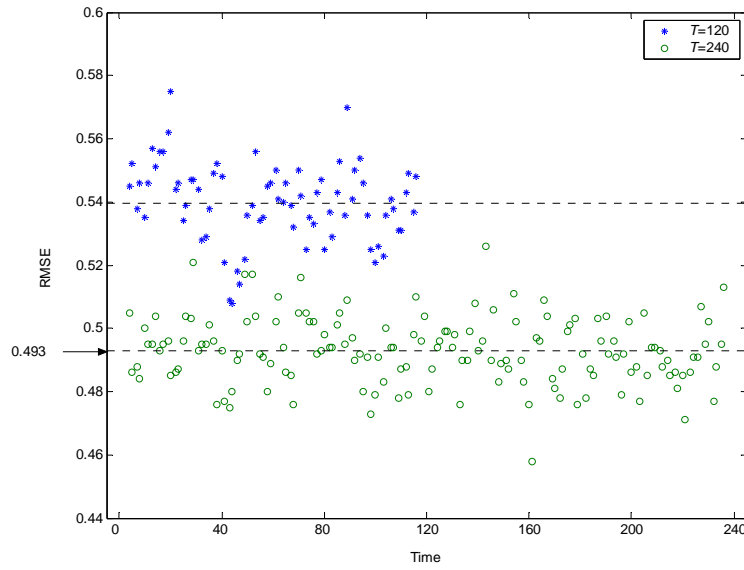
	$T = 120$			$T = 240$		
	Complete data	Mixed-freq. data		Complete data	Mixed-freq. data	
		Init. EM	Main EM		Init. EM	Main EM
$\Omega_{11} = 25$						
Mean	24.470	26.146	25.893	24.843	25.591	25.523
RMSE	3.114	3.932	3.855	2.240	2.585	2.570
$\Omega_{12} = 7.5$						
Mean	7.374	10.341	9.928	7.469	8.893	8.743
RMSE	1.508	4.402	4.134	1.074	2.441	2.350
$\Omega_{13} = 2.5$						
Mean	2.461	3.952	3.769	2.486	3.187	3.124
RMSE	0.511	2.272	2.159	0.352	1.229	1.189
$\Omega_{22} = 9$						
Mean	8.806	14.003	13.338	8.931	11.599	11.291
RMSE	1.177	6.894	6.422	0.863	3.832	3.625
$\Omega_{23} = 1.5$						
Mean	1.461	4.039	3.689	1.474	2.799	2.633
RMSE	0.307	3.388	3.148	0.216	1.862	1.756
$\Omega_{33} = 1$						
Mean	0.976	2.104	1.938	0.984	1.543	1.466
RMSE	0.132	1.521	1.411	0.092	0.811	0.764

Table 3. Decline rates of RMSE in CASEs I and II as T goes from 120 to 240

Parameter	Complete data	CASE I: temp. agg.		Complete data	CASE II: skip sampled	
		Init. EM	Main EM		Init. EM	Main EM
a_1	29%	39%	34%	29%	35%	34%
a_2	38%	44%	38%	38%	41%	38%
a_3	36%	46%	38%	36%	41%	45%
b_1	63%	66%	70%	63%	61%	65%
b_2	65%	67%	71%	65%	62%	66%
γ_{11}	46%	53%	60%	46%	53%	57%
γ_{21}	52%	56%	64%	52%	54%	60%
γ_{31}	52%	55%	64%	52%	54%	60%
Ω_{11}	28%	44%	35%	28%	34%	33%
Ω_{12}	29%	52%	44%	29%	45%	43%
Ω_{13}	31%	52%	46%	31%	46%	45%
Ω_{22}	27%	49%	44%	27%	44%	44%
Ω_{23}	30%	49%	44%	30%	45%	44%
Ω_{33}	30%	50%	46%	30%	47%	46%

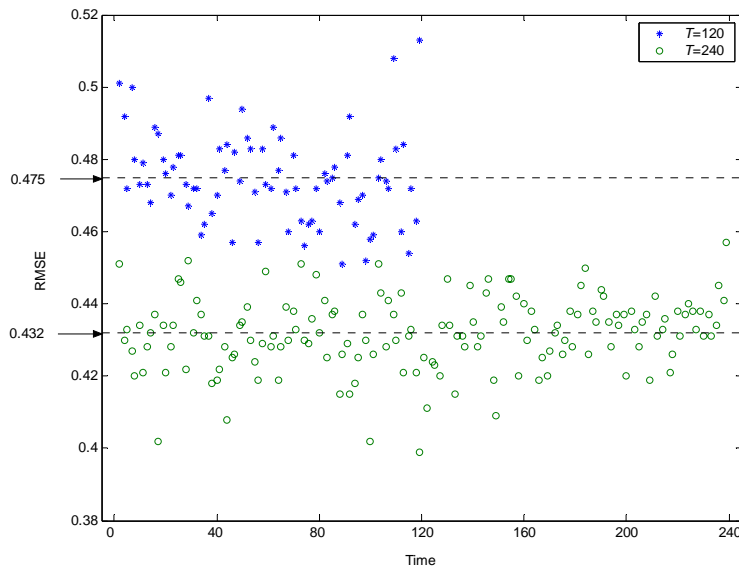
Note: Decline rate = $1 - (\text{RMSE for } T=240) / (\text{RMSE for } T=120)$

Figure 1. RMSE of monthly estimated y_{3t} for CASE I



Note: The dotted lines denote average values of the corresponding RMSE.

Figure 2. RMSE of monthly estimated y_{3t} for CASE II



Note: Same as in Figure.

7. Examples

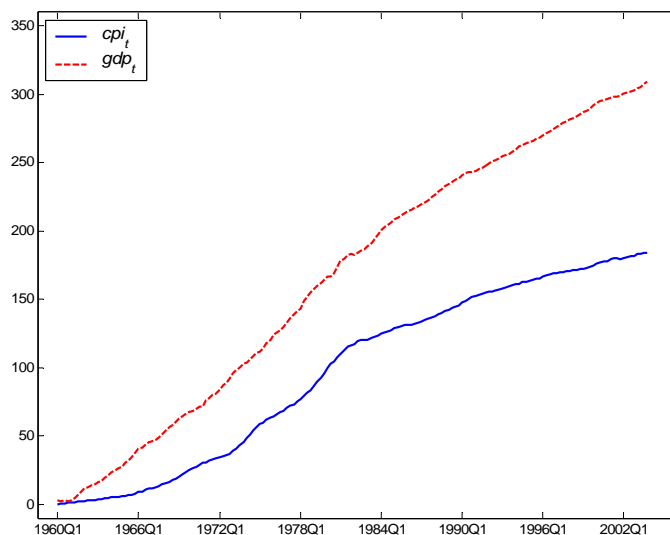
We illustrate the proposed main EM algorithm using mixed-frequency U.S. data of monthly CPI and quarterly GDP, from December 1959 to December 2003, which comprises 176 quarters or 529 months of observations. The CPI and GDP data are from the Bureau of Labor Statistics (www.bls.gov) and from the Bureau of Economic Analysis (www.bea.gov), are seasonally adjusted, have a base CPI value of 100 in 1982-1984, and have a unit of GDP in billions of current dollars. The original data, denoted CPI_t and GDP_t in month t , were transformed to cpi_t and gdp_t by taking natural logarithms and subtracting from them the natural logs at the starting dates, as

$$y_t = \begin{pmatrix} cpi_t \\ gdp_t \end{pmatrix} = \begin{pmatrix} \log(CPI_t) - \log(CPI_{1959:12}) \\ \log(GDP_t) - \log(GDP_{1959:12}) \end{pmatrix}.$$

Henceforth, the lower-case variable names, cpi and gdp , will refer to these transformations of CPI and GDP .

As mentioned in section 3.2, we first estimate the CI rank and the CI vector by using a quarterly single-frequency sample, obtained by picking one monthly value of cpi per quarter (we call this “skip-sampling”) and keeping quarterly gdp as is. In Figure 3, the quarterly single-frequency data indicate that the estimated model may have unrestricted, constant and linear, deterministic terms. The model selection criteria, minimum AIC and SBC, both indicated choosing a VAR(4) model. Then, applying Johansen’s trace test to the VAR(4) model, estimated using the single-frequency data, we obtained a p-value of .0437, which indicated a CI rank of one. The corresponding estimate of the CI vector was $(1, -1.041)'$. For the mixed-frequency observations on $y_t = (cpi_t, gdp_t)'$, we obtained $(1, -3.123)'$ as the estimate of the cointegrating vector by using the relationship in Pons and Sansó (2005).

Figure 3. Quarterly cpi_t and gdp_t from 1960:1 to 2003:12



Note: Quarterly cpi_t are skip samplings of observed monthly cpi_t in a quarter.

Table 4 reports minimum AIC and SBC, used to select the best monthly (highest frequency) model and associated log-likelihood values. Both AIC and SBC select the VAR(2) model,

$$\Delta y_t = \begin{pmatrix} -0.0149 \\ 0.1599 \end{pmatrix} + \begin{pmatrix} -0.0040 \\ -0.0007 \end{pmatrix} t + \begin{pmatrix} -0.0161 \\ -0.0022 \end{pmatrix} (1, -3.1289) y_{t-1} \\ + \begin{pmatrix} 0.5156 & 0.0046 \\ 0.1250 & -0.0042 \end{pmatrix} \Delta y_{t-1} + \varepsilon_t,$$

$$\hat{\text{var}}(\varepsilon_t) = \begin{pmatrix} 0.0463 & -0.0036 \\ -0.0036 & 0.0388 \end{pmatrix},$$

as the best VAR(p) model. For this model, we computed monthly smoothed estimates of gdp_t and monthly forecasts of cpi_t and gdp_t .

Table 5 shows monthly smoothed estimates of gdp_t , from 2003:1 to 2003:12, as examples of monthly smoothed estimates of gdp_t in the sample period. Two types of estimates are computed, temporally aggregated and not temporally aggregated. The temporally-aggregated estimates reflect low-frequency quarterly aggregated gdp_t , i.e., sums of three consecutive monthly gdp_t . The not-temporally-aggregated estimates reflect high-frequency monthly disaggregated gdp_t , i.e., monthly gdp_t by itself. We note that in the table the not-temporally aggregated estimates are multiplied by 3 in order to be in quarterly

form comparable to observed quarterly gdp_t .

Tables 6 and 7 state out-of-sample (that is, out of model estimation sample) forecasts of cpi_t and gdp_t for 2004 and compare these with true values. The tables show smaller forecast errors when using mixed-frequency data than when using single-frequency data. Specifically, the forecast errors for mixed-frequency data are smaller on average by about 52% in Table 6 and about 32% in Table 7.

Table 4. Model selection criteria of optimal VAR(p) for mixed-frequency data

P	Log likelihood function ($\times 10^6$)	AIC ($\times 10^3$)	SBC ($\times 10^3$)
1	-1.7246	6.5326	6.5327
2	-1.7245	6.5323	6.5324
3	-1.7245	6.5323	6.5324
4	-1.7245	6.5323	6.5325
5	-1.7245	6.5324	6.5326
6	-1.7249	6.5337	6.5340
7	-1.7251	6.5346	6.5349
8	-1.7254	6.5356	6.5359
9	-1.7250	6.5341	6.5344
10	-1.7245	6.5325	6.5329

Table 5. Monthly smoothed estimates of in-sample gdp_t , 2003:1 to 2003:12

Year: month	Observed	Temp. agg.	Not temp. agg.
2003:1		303.33	303.75
2003:2		303.74	304.14
2003:3	304.15	304.15	304.53
2003:4		304.56	305.01
2003:5		304.96	305.34
2003:6	305.44	305.44	305.97
2003:7		306.06	306.87
2003:8		306.82	307.62
2003:9	307.55	307.55	308.16
2003:10		308.14	308.64
2003:11		308.59	308.97
2003:12	308.93	308.93	309.18

Note: “Temporally aggregated” and “not temporally aggregated” estimates in columns 3 and 4, respectively, reflect quarterly sums of monthly values ending in the indicated month and monthly values for that month multiplied by 3, in order to be in quarterly form comparable to observed quarterly gdp in column 2.

Table 6. Monthly out-of-sample forecasts of cpi_t

Year: month	Observed	Single-frequency		Mixed-frequency	
		Forecast	Error	Forecast	Error
2004:1	184.39			184.11	0.28
2004:2	184.71			184.32	0.39
2004:3	185.14	183.99	1.14	184.53	0.61
2004:4	185.35			184.73	0.62
2004:5	185.94			184.93	1.01
2004:6	186.20	184.39	1.81	185.13	1.07
2004:7	186.15			185.32	0.83
2004:8	186.20			185.50	0.70
2004:9	186.36	184.61	1.75	185.69	0.67
2004:10	186.94			185.87	1.07
2004:11	187.20			186.04	1.16
2004:12	187.20	184.75	2.45	186.22	0.98

Note: “Out-of-sample” means for months beyond earlier months, from 1959:12 to 2003:12, used to estimate the model which was used to produce the forecasts.

Table 7. Monthly out-of-sample forecasts of gdp_t

Year: month	Observed	Single-frequency		Mixed-frequency		
		Forecast	Error	Forecast Low Freq.	Error	Forecast High Freq.
2004:1				309.22		309.51
2004:2				309.51		309.84
2004:3	310.71	309.47	1.24	309.83	0.88	310.14
2004:4				310.15		310.47
2004:5				310.46		310.77
2004:6	312.30	310.23	2.07	310.78	1.52	311.10
2004:7				311.09		311.40
2004:8				311.41		311.73
2004:9	313.65	310.74	2.91	311.72	1.93	312.03
2004:10				312.04		312.36
2004:11				312.35		312.66
2004:12	315.11	311.22	3.89	312.66	2.45	312.96

Note: Same as in Table 6.

8. Conclusion

We have developed and illustrated a method, for estimating a multivariate cointegrated VAR model with mixed-frequency time-series data, by using a state-space representation of an error correction model. The method allows us not only to estimate such a model using mixed-frequency data, but also to estimate missing or unobserved high-frequency values of the low-frequency variables. Monte Carlo experiments, applied to mixed-frequency data, with missing observations, and to single-frequency data, with complete observations, indicate that the proposed method performs well with missing data due to mixed frequencies.

Appendix A

To compute estimated parameters $\Gamma^{(l+1)}$ and $\Sigma^{(l+1)}$ at iteration $l+1$ of the EM algorithm, we differentiate equation (11) with respect to Γ and Σ^{-1} , using equation (9), and obtain

$$\frac{\partial Q(\theta | \theta^{(l)})}{\partial \Gamma} = -2\Sigma^{-1} \left\{ A \sum_{t=1}^T E_t(x_t x_{t-1}' | Y_T^+) B' - \Gamma B \sum_{t=1}^T E_t(x_{t-1} x_{t-1}' | Y_T^+) B' \right\}, \quad (\text{A1})$$

$$\begin{aligned} \frac{\partial Q(\theta | \theta^{(l)})}{\partial \Sigma^{-1}} = & \frac{T}{2} \{2\Sigma - \text{diag}(\Sigma)\} - \frac{1}{2} \left[2 \sum_{t=1}^T E_t \left\{ (Ax_t - \Gamma Bx_{t-1})(Ax_t - \Gamma Bx_{t-1})' | Y_T^+ \right\} \right. \\ & \left. - \text{diag} \left(\sum_{t=1}^T E_t \left\{ (Ax_t - \Gamma Bx_{t-1})(Ax_t - \Gamma Bx_{t-1})' | Y_T^+ \right\} \right) \right], \quad (\text{A2}) \end{aligned}$$

where $\text{diag}(\cdot)$ denotes the diagonal matrix formed from the diagonal elements of the argument matrix. To find the estimated parameters $\tilde{\beta}^{(l+1)}$, we differentiate equation (11) with respect to β_0 , using equation (10), and obtain

$$\begin{aligned} \frac{\partial Q(\theta | \theta^{(l)})}{\partial \text{vec}(\beta_0)} = & -2 \sum_{t=1}^T E_t \left[\left\{ \alpha' \otimes (Dx_{t-1}) \right\} \Sigma^{-1} (Ax_t - \Gamma^* B^* x_{t-1}) | Y_T^+ \right] \\ & + 2 \sum_{t=1}^T E_t \left[\left\{ \alpha' \otimes (Dx_{t-1}) \right\} \Sigma^{-1} \left\{ \alpha \otimes (Dx_{t-1})' \right\} | Y_T^+ \right] \text{vec}(\beta_0). \quad (\text{A3}) \end{aligned}$$

By noting that $\left\{ \alpha' \otimes (Dx_{t-1}) \right\} \Sigma^{-1} Ax_t = \text{vec} \left\{ Dx_{t-1} (x_t' A' \Sigma^{-1}) \alpha \right\}$, we can rewrite equation (A3) as

$$\frac{\partial Q(\theta | \theta^{(l)})}{\partial \text{vec}(\beta_0)} = -2 \text{vec} \left\{ D \sum_{t=1}^T E_t(x_{t-1} x_t' | Y_T^+) A' \Sigma^{-1} \alpha \right\}$$

$$\begin{aligned}
& + 2\text{vec}\left\{D\sum_{t=1}^T E_l(x_{t-1}x'_{t-1} | Y_T^+)(\Gamma^* B^*)'\Sigma^{-1}\alpha\right\} \\
& + 2\left[(\alpha'\Sigma^{-1}\alpha)\otimes\left\{D\sum_{t=1}^T E_l(x_{t-1}x'_{t-1} | Y_T^+)D'\right\}\right]\text{vec}(\beta_0).
\end{aligned}$$

Appendix B

Let $x_t^s = E_l(x_t | Y_s^+)$, $P_t^s = \text{cov}_l(x_t | Y_s^+)$ and $P_{t,t-1}^s = \text{cov}_l(x_t, x_{t-1} | Y_s^+)$, where $E_l(\cdot | Y_s^+)$ and $\text{cov}_l(\cdot | Y_s^+)$ denote the conditional expectation and conditional covariance with respect to the density based on $\theta^{(l)}$.

We calculate the prediction and updating recursions using the following equations (for example, Shumway and Stoffer, 2000). For $t = 1, \dots, T$,

$$x_t^{t-1} = Fx_{t-1}^{t-1}, \quad P_t^{t-1} = FP_{t-1}^{t-1}F' + GG', \quad (\text{B1})$$

$$x_t^t = x_t^{t-1} + K_t(y_t^+ - H_t x_t^{t-1}), \quad P_t^t = P_t^{t-1} - K_t H_t P_t^{t-1}, \quad (\text{B2})$$

where $K_t = P_t^{t-1}H_t'(H_t P_t^{t-1}H_t' + Q_t Q_t')^{-1}$. We start iterations (B1) and (B2) by setting $x_0^0 = \lambda$ and $P_0^0 = \Lambda$. In order to calculate x_t^T , P_t^T and $P_{t,t-1}^T$ using equations (B1) and (B2), for $t = T, \dots, 1$, we iterate over the backwards recursions

$$r_t = H_t'(H_t P_t^{t-1}H_t' + Q_t Q_t')^{-1}(y_t^+ - H_t x_t^{t-1}) + L_t' r_{t+1} \quad (\text{B3})$$

$$R_t = H_t'(H_t P_t^{t-1}H_t' + Q_t Q_t')^{-1}H_t + L_t' R_{t+1} L_t \quad (\text{B4})$$

where $r_{T+1} = 0$, $R_{T+1} = 0$, and $L_t = F(I_s - K_t H_t)$. Following Durbin and Koopman (2001), we obtain the smoothing equations

$$x_t^T = x_t^{t-1} + P_t^{t-1} r_t, \quad P_t^T = P_t^{t-1} - P_t^{t-1} R_t P_t^{t-1}, \quad (\text{B5})$$

for $t = 1, \dots, T$, and

$$P_{t+1,t}^T = (I_s - P_{t+1}^t R_{t+1}) L_t P_t^{t-1}, \quad (\text{B6})$$

for $t = 1, \dots, T-1$.

Appendix C

Here, we discuss the modifications of state-space equations (19) to (22) which are necessary in order for the state-space representation to incorporate the constant and linear deterministic terms of equation (17). We follow the classifications of the restrictions given by Johansen (1996).

TYPE 1 restrictions are $\gamma_1 = \alpha\rho_1$ or $\mu_1 = \alpha\rho_1$ and $\gamma_2 = 0$ or $\mu_2 = 0$. In this case, ρ_1 can be put into the cointegration relationship by replacing β_0 with $\tilde{\beta}_0 = (\rho_1', \beta_0')$, so that ρ_1 is estimated instead of μ_1 . TYPE 2 restrictions are an unrestricted γ_1 and $\gamma_2 = 0$. The model is easily correspondingly changed by deleting the terms involving γ_2 and μ_2 , which are now zero. TYPE 3 restrictions are an unrestricted γ_1 and $\gamma_2 = \alpha\rho_2$ or $\mu_2 = \alpha\rho_2$. In this case, ρ_2 is put into the cointegration relationship and β_0 is replaced by $\tilde{\beta}_0 = (\rho_2', \beta_0')$. Table 8 states these modifications explicitly in algebraic notation.

Table 8. Definitions of state-space components for Type I to Type III restrictions on deterministic terms

	Type 1	Type 2	Type 3
\tilde{x}_t	$(1, x_t)'$	$(1, x_t)'$	$(1, t, x_t)'$
\tilde{F}	$\begin{bmatrix} 1 & O_{1 \times s} \\ \alpha \rho_1 & F \\ 0_{s-n} & \end{bmatrix}$	$\begin{bmatrix} 1 & O_{1 \times s} \\ \mu_1 & F \\ 0_{s-n} & \end{bmatrix}$	$\begin{bmatrix} 1 & 0 & O_{2 \times s} \\ 1 & 1 & \\ \mu_1 & \alpha \rho_2 & F \\ 0_{s-n} & 0_{s-n} & \end{bmatrix}$
\tilde{G}	$\begin{bmatrix} O_{1 \times n} \\ G \end{bmatrix}$	$\begin{bmatrix} O_{1 \times n} \\ G \end{bmatrix}$	$\begin{bmatrix} O_{2 \times n} \\ G \end{bmatrix}$
\tilde{H}_t	$[O_{n \times 1} \ H_t]$	$[O_{n \times 1} \ H_t]$	$[O_{n \times 2} \ H_t]$
\tilde{A}	$[O_{n \times 1} \ A]$	$[O_{n \times 1} \ A]$	$[O_{n \times 2} \ A]$
$\tilde{\Gamma}$	Γ	$[\mu_1 \ \Gamma]$	$[\mu_1 \ \Gamma]$
$\tilde{\Gamma}^*$	$[0_n \ \Gamma^*]$	$[\mu_1 \ \Gamma^*]$	$[\mu_1 \ 0_n \ \Gamma^*]$
β_0	$\tilde{\beta}_0 = (\rho_1', \beta_0)'$	β_0	$\tilde{\beta}_0 = (\rho_2', \beta_0)'$
B_ρ	$\begin{bmatrix} \rho_1' & B \\ O_{(p-1)n \times 1} & \end{bmatrix}$	$[O_{\{h+(p-1)n\} \times 1} \ B]$	$\begin{bmatrix} 0_h \ \rho_2' & B \\ O_{(p-1)n \times 2} & \end{bmatrix}$
\tilde{B}	B_ρ	$\begin{bmatrix} 1 & O_{1 \times s} \\ B_\rho \end{bmatrix}$	$\begin{bmatrix} 1 & O_{1 \times (s+1)} \\ B_\rho \end{bmatrix}$
\tilde{B}^*	$\begin{bmatrix} 1 & O_{1 \times s} \\ O_{np \times 1} & B^* \end{bmatrix}$	$\begin{bmatrix} 1 & O_{1 \times s} \\ O_{np \times 1} & B^* \end{bmatrix}$	$\begin{bmatrix} I_2 & O_{2 \times s} \\ O_{np \times 2} & B^* \end{bmatrix}$
\tilde{D}	$\begin{bmatrix} 1 & O_{1 \times s} \\ O_{d \times 1} & D \end{bmatrix}$	$[O_{d \times 1} \ D]$	$\begin{bmatrix} 0 & 1 & O_{1 \times s} \\ O_{d \times 2} & D \end{bmatrix}$

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