A two-step estimator for large approximate dynamic factor models based on Kalman filtering*

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Abstract

This paper shows consistency of a two step estimator of the parameters of a dynamic approximate factor model when the panel of time series is large (n large). In the first step, the parameters are first estimated from an OLS on principal components. In the second step, the factors are estimated via the kalman smoother. This projection allows to consider dynamics in the factors and heteroskedasticity in the idiosyncratic variance. The analysis provides theoretical backing for the estimator considered in Giannone, Reichlin, and Sala (2004) and Giannone, Reichlin, and Small (2005).

JEL Classification: C51, C32, C33.

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1 Introduction

We consider a "large" panel of time series and assume that it can be represented by an approximate factor structure whereby the dynamics of each series is split in two orthogonal components – one capturing the bulk of cross-sectional comovements and driven by few common factors and the other being composed of poorly cross-correlated elements. This model has been introduced by Chamberlain and Rothschild (1983) and generalized to a dynamic framework by Forni, Hallin, Lippi, and Reichlin (2000); Forni and Lippi (2001) and Stock and Watson (2002a,b).

As in many other papers in the literature, this paper studies the estimation of the common factors and consistency and rates for the size of the cross-section n and the sample size T going to infinity.

The literature has extensively studied the particular case in which the factors are estimated by principal components (Bai, 2003; Bai and Ng, 2002; Forni, Hallin, Lippi, and Reichlin, 2005b; Forni, Giannone, Lippi, and Reichlin, 2005a; Stock and Watson, 2002a,b). It has been shown that the latter are (n,T) consistent estimates of a rotation of the factors. Consistency is achieved even if principal components do not exploit likely features of the data generating process, such as heterogeneous signal to noise ratio (cross-sectional heteroscedasticity of the idiosyncratic component), dynamic of the factors and dynamic in the idiosyncratic component.

The literature has also studied a number of methods to exploit those features. Forni, Hallin, Lippi, and Reichlin (2005b) has proposed a two-step approach based on principal components in the frequency domain to exploit, when extracting the common factors, the cross-sectional heteroscedasticity of the idiosyncratic component and the dynamic properties of the data; Boivin and Ng (2003) and Forni and Reichlin (2001) have used iteratively re-weighted principal components and Boivin and Ng (2005), D'Agostino and Giannone (2005), Stock and Watson (2005) have studied the empirical relevance of such efficiency improvements. Finally, Giannone, Reichlin, and Sala (2004) and Giannone, Reichlin, and Small (2005) have introduced a parametric time domain two-step estimator involving principal components and Kalman filter to exploit both factor dynamics and idiosyncratic heteroscedacticity.

This paper develops the parametric approach to study these potential efficiency improvements in a unified framework.

We parameterize the dynamics of the factors as in Forni, Giannone, Lippi, and Reichlin (2005a). The parameters of the model can then be estimated by simple least squares by treating the principal components as if they were the true common factors. These estimated parameters can be used to project onto the observations. We consider three cases, each corresponding to an estimator under different forms of misspecification: factor dynamics, idiosyncratic heteroscedacticity and idiosyncratic dynamics (principal components); factor and idiosyncratic dynamics (reweighted principal components); idiosyncratic dynamics only (Kalman filter).

Each projection corresponds to a different two-step estimator whereby the first step involves the estimation of the parameters and the second step the application of the Kalman smoother.

We prove consistency for such estimators and design an empirical exercise that

allows to evaluate the efficiency improvement in small sample for the dynamic and the heteroscedasticity case.

We should stress that the use of the Kalman filter, beside achieving possible efficiency improvements, allows useful empirical applications. First, the treatment of unbalanced panels, particularly interesting for forecasting current quarter GDP at dates in which not all data included in the panel are released (see Giannone, Reichlin, and Sala, 2004; Giannone, Reichlin, and Small, 2005). Second, "cleaning", through the second step, the estimate of the factors, allows a better reconstruction of the common shocks considered in the structural factor model Giannone, Reichlin, and Sala (2004). Finally, such parametric approach allows to easily evaluate uncertainty in the estimates of the factors as shown in both the papers just cited.

Let us finally note that similar reasoning to that applied to this paper can be applied to use principal components to initialize the algorithm for maximum likelihood estimation. We analyze such approach in the empirical section while we study consistency of maximum likelihood estimator in a separate paper Doz, Giannone, and Reichlin (2005).

The paper is organized as follows. Section two introduces models and assumptions. Section three analyzes the projections under the different cases and shows, for known parameters, how to extract the n consistent factors under different hypothesis on specification error. Section four contains the main propositions which show consistency and (n, T) rates for the two step estimators. Section five presents the empirical application on both real and artificial data. Section six concludes.

2   The Models

We consider the following model:

$$X_t = \Lambda_0^* F_t + \xi_t$$

where

$X_t = (x_{1t}, ..., x_{nt})'$ is a $(n \times 1)$ stationary process

$\Lambda_0^* = (\lambda_{0,ij}^*)$ is the $n \times r$ matrix of factor loadings

$F_t = (f_{1t}, ..., f_{rt})'$ is a $(r \times 1)$ stationary process (common factors)

$\xi_t = (\xi_{1t}, ..., \xi_{nt})'$ is a $(n \times 1)$ stationary process (idiosyncratic component)

(F_t) and (ξ_t) are two orthogonal processes

Note that X_t, Λ_0^*, ξ_t depend on n but, in this paper, we drop the subscript for sake of simplicity.

The general idea of the model is that the observable variables can be decomposed in two orthogonal unobserved processes: the common component driven by few common shocks which captures the bulk of the covariation between the time series, and the

idiosyncratic component which is driven by n shocks generating dynamics which is series specific or local.

We have the following decomposition of the covariance matrix of the observables:

$$\Sigma_0 = \Lambda_0^* \Phi_0^* \Lambda_0^{*\prime} + \Psi_0$$

where $\Psi_0 = \mathrm{E}[\xi_t \xi_t']$ and $\Phi_0^* = \mathrm{E}[F_t F_t']$. It is well-known that the factors are defined up to a pre-multiplication by an invertible matrix, so that it is possible to choose $\Phi_0^* = I_r$. Even in this case, the factors are defined up to a pre-multiplication by an orthogonal matrix, a point that we make more precise below.

We also have the following decomposition of the auto-covariance matrix of order h of the observables:

$$\Sigma_0(h) = \Lambda_0^* \Phi_0^*(h) \Lambda_0^{*\prime} + \Psi_0(h)$$

where $\Sigma_0(h) = \mathrm{E}[x_t x_{t-h}']$, $\Phi_0^*(h) = \mathrm{E}[F_t F_{t-h}']$, and $\Psi_0(h) = \mathrm{E}[\xi_t \xi_{t-h}']$

Remark 1: Bai (2003); Bai and Ng (2002) and Stock and Watson (2002a) consider also some form of non-stationarity. Here we do not do it for simplicity. The main arguments used in what follows still hold under the assumption of weak time dependence of the common and the idiosyncratic component.

More precisely, we make the following set of assumptions:

(A1) *For any n, (X_t) is a stationary process with zero mean and finite second order moments.*

(A2) *The x_{it}'s have uniformly bounded variance : $\exists M / \forall (i,t) V x_{it} = \sigma_{0,ii} \leq M$*

(A3) - *(F_t) and (ξ_t) are independent processes.*
   - *(F_t) admits a Wold representation: $F_t = C_0(L)\varepsilon_t = \sum_{k=0}^{+\infty} C_k \varepsilon_{t-k}$ such that: $\sum_{k=0}^{+\infty} \|C_k\| < +\infty$, and ε_t admits finite moments of order four.*
   - *For any n, (ξ_t) admits a Wold representation: $\xi_t = D_0(L)v_t = \sum_{k=0}^{+\infty} D_k v_{t-k}$ where $\sum_{k=0}^{+\infty} \|D_k\| < +\infty$ and v_t is a strong white noise such that: $\exists M / \forall (i,t) E v_{it}^4 \leq M$*

Note first that (v_t) and $D_0(L)$ are not nested matrices: when n increases because a new observation is added to x_t, a new observation is also added to ξ_t but the innovation process and the filter $D_0(L)$ entirely change.

A convenient way to parameterize the dynamics is to further assume that the common factors following a VAR process so that (A3) can be replaced by the following assumption (see Forni et al., 2005a, for a discussion):

(A3') - A VAR approximation for the factors: $A_0^*(L)F_t = u_t$ where $A_0^*(z) \neq 0$ for $|z| \leq 1$ and $A_0^*(0) = I_r$.

- Independence between the shocks driving the factor and the idiosyncratic processes: $(u_t', v_t')' \sim \text{WN}(0, \Delta)$, with Δ a diagonal matrix.

For any n, we denote by $\bar{\psi}_0 = \frac{1}{n} \sum_{j=1}^n \text{E}\xi_{it}^2$, and in the whole paper, $A_0^*(L)$, Ψ_0, $D_0(L)$, $\bar{\psi}_0$ denote the true values of the parameters.

Given the size of the cross-section n, the model is identified provided that the number of common factors (r) is small with respect to the size of the cross-section (n), and the idiosyncratic component is orthogonal at all leads and lags, i.e. $D_0(L)$ is a diagonal matrix (exact factor model). This version of the model is what proposed by Engle and Watson, 1981 and estimated by them by Maximum Likelihood [1]. In what follows, we will not impose such restriction and work under the assumption of some form of weak correlation among idiosyncratic components (approximate factor model) as in the n large, new generation factor literature. There are different ways to impose identifying assumptions that restrict the cross-correlation of the idiosyncratic elements and preserve the commonality of the common component as n increases. We will assume that the Chamberlain and Rothschild (1983)'s conditions are satisfied. More precisely, denoting by $\lambda_{min}(A)$ and $\lambda_{max}(A)$ the smallest and the greatest eigenvalues of a matrix A, and by $\|A\| = (\lambda_{max}(A'A))^{1/2}$, we make the two following assumptions:

(CR1) $\liminf_{n \to \infty} \frac{1}{n} \lambda_{min}(\Lambda_0^{*\prime} \Lambda_0^*) > 0$
(pervasiveness of the common component)

(CR2) $\exists \bar{\lambda} / \sum_{h \in \mathbb{Z}} \|\Psi_0(h)\| < \bar{\lambda}$ (limitation of cross-sectional time autocorrelation of the idiosyncratic component)

Note that assumption (CR2) is achieved as soon as the two following assumptions are made:

- $\exists M / \forall n \ \|\text{E}[v_t v_t']\| \leq M$

- $\exists M / \forall n \ \sum_{k=0}^{+\infty} \|D_k\| \leq M$

We also suppose, as in Forni et al. (2004), that all the eigenvalues of $\Lambda_0^{*\prime} \Lambda_0^*$ diverge at the same rate, which is equivalent to the following further assumption:

(CR3) $\limsup_{n \to \infty} \frac{1}{n} \lambda_{max}(\Lambda_0^{*\prime} \Lambda_0^*)$ *is finite*

Finally, we make the two next assumptions:

(A4) $\inf_n \lambda_{min}(\Psi_0) = \underline{\lambda} > 0$

---

[1] Identification conditions for the model for a fixed cross-sectional dimensions (n) are studied in Geweke and Singleton (1980).

(A5) $\Lambda_0^{*\prime}\Lambda_0^*$ has distinct eigenvalues [2].

Remark 2: These assumptions are slightly different than those introduced by Stock and Watson (2002a) and Bai and Ng (2002) but have a similar role. They have been generalized for the dynamic case by Forni et al. (2000) and Forni and Lippi (2001)

As we said before, the common factors, and the factor loadings, are identified up to a normalization. In order to give a precise statement of the consistency results in our framework, we will use here a particular normalization. Let us define:

- D_0 as the diagonal matrix whose diagonal entries are the eigenvalues of $\Lambda_0^{*\prime}\Lambda_0^*$ in decreasing order,

- Q_0 as the matrix of a set of unitary eigenvectors associated with D_0,

- $\Lambda_0 = \Lambda_0^* Q_0$, so that $\Lambda_0'\Lambda_0 = D_0$ and $\Lambda_0\Lambda_0' = \Lambda_0^*\Lambda_0^{*\prime}$,

- $P_0 = \Lambda_0 D_0^{-1/2}$ so that $P_0'P_0 = I_r$,

- $G_t = Q_0'F_t$.

With these new notations, the model can also be written as:

$$X_t = \Lambda_0 G_t + \xi_t \tag{2.1}$$

We then have : $\mathrm{E}[G_t G_t'] = I_r$, and $\mathrm{E}[G_t G_{t-h}'] = \Phi_0(h) = Q_0'\Phi_0^*(h)Q_0$ for any h . It then follows that:

$$\Sigma_0 = \Lambda_0^*\Lambda_0^{*\prime} + \Psi_0 = \Lambda_0\Lambda_0' + \Psi_0$$

and that, for any h: $\Sigma_0(h) = \Lambda_0^*\Phi_0^*(h)\Lambda_0^{*\prime} + \Psi_0(h) = \Lambda_0\Phi_0(h)\Lambda_0' + \Psi_0(h)$.

Note that, in the initial representation of the model, the matrices Λ_0^* are supposed to be nested (when an observation is added to x_t, a line is added to the matrix Λ_0^*), whereas the Λ_0 matrix is entirely modified. However, as Q_0 is invertible, G_t and F_t have the same range, likewise Λ_0 and Λ_0^* have the same range[3]. In addition, assumptions (A1) to (A5) and (CR1) to (CR2) are satisfied if we replace Λ_0^* with Λ_0, and F_t with G_t. If also assumption (A3') holds then G_t also has a VAR representation. Indeed, as $Q_0 G_t = F_t$, we have: $A_0^*(L)G_t = u_t$, and $Q_0'A_0^*(L)G_t = Q_0'u_t$. We then can write:

$$A_0(L)G_t = w_t,$$

with $A_0(L) = Q_0'A_0^*(L)Q_0, w_t = Q_0'u_t^*, A_0(z) \neq 0$ for $|z| \leq 1$, and $A_0(0) = I_r$.

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[2] This assumption is usual in this framework, and is made to avoid useless mathematical complications. However, in case of multiple eigenvalues, the results would remained unchanged.

[3] It is worth noticing that Q_0 is uniquely defined up to a sign change of its columns and that G_t is uniquely defined up to a sign change of its components (this will be used below). Indeed, as $\Lambda_0^{*\prime}\Lambda_0^*$ is supposed to have distinct eigenvalues, Q_0 is uniquely defined up to a sign change of its columns. Then, if Δ is a diagonal matrix whose diagonal terms are ± 1, and if Q_0 is replaced by $Q_0\Delta$, Λ_0 is replaced by $\Lambda_0\Delta$ and G_t is replaced by ΔG_t.

3 Approximating projections

We are interested in extracting the factors G_t from the observables $X_1, ..., X_T$ where T is the sample size. In particular we are interested in the following linear projection:

$$G_{t|T} = \text{Proj}_\Omega[G_t|X_s, s \leq T]$$

where $\Omega = \{\Lambda^*, A^*(L), D(L)\}$ denotes the triple defining the model, which can be equivalently written as $\Omega = \{\Lambda, A(L), D(L)\}$, if we use the new parameterization of the model.

If the model is Gaussian, i.e. if u_t and v_t are normally distributed, then

$$\text{Proj}_\Omega[G_t|X_s, s \leq T] = \text{E}_\Omega[G_t|X_s, s \leq T]$$

Moreover, if the projection is taken under the true parameter values, $\Omega_0 = \{\Lambda_0, A_0(L), D_0(L)\}$, then we have optimality in mean square sense.

In what follows, we consider other projections of G_t, associated to misspecified models. Although not optimal, these projections also give consistent approximations of G_t. The simplest projection is obtained under the triple $\Omega_0^{R1} = \left\{\Lambda_0, I_r, \sqrt{\bar{\psi}_0}I_n\right\}$, that is under an approximating model according to which the common factors are white noise with covariance I_r and the idiosyncratic components are cross-sectionally homoscedastic white noises with variance $\bar{\psi}_0$. We have:

$$\text{Proj}_{\Omega_0^{R1}}[G_t|X_s, s \leq T] = \text{E}_{\Omega_0^{R1}}[G_t X_t'] \left[\text{E}_{\Omega_0^{R1}}[X_t X_t']\right]^{-1} X_t = \Lambda_0' \left(\Lambda_0 \Lambda_0' + \bar{\psi}_0 I_n\right)^{-1} X_t.$$

Simple calculations show that, when Ψ_{0R} is an invertible matrix of order n:

$$\left(\Lambda_0 \Lambda_0' + \Psi_{0R}\right)^{-1} = \Psi_{0R}^{-1} - \Psi_{0R}^{-1}\Lambda_0 \left(\Lambda_0' \Psi_{0R}^{-1}\Lambda_0 + I_r\right)^{-1} \Lambda_0' \Psi_{0R}^{-1}$$

and that the previous expression can also be written as:

$$\text{Proj}_{\Omega_0^{R1}}[G_t|X_s, s \leq T] = \left(\Lambda_0' \bar{\psi}_0^{-1}\Lambda_0 + I_r\right)^{-1} \Lambda_0' \bar{\psi}_0^{-1} X_t = \left(\Lambda_0' \Lambda_0 + \bar{\psi}_0 I_r\right)^{-1} \Lambda_0' X_t$$

which is, by assumption CR1, asymptotically equivalent to the OLS regression of X_t on the factor loadings Λ_0.

It is clear that, under conditions CR1 and CR2, such simple OLS regression provides a consistent estimate of the unobserved common factors as the cross-section becomes large[4]. In particular,

$$\text{Proj}_{\Omega_0^{R1}}[G_t|X_s, s \leq T] \xrightarrow{m.s.} G_t \text{ as } n \to \infty$$

---

[4]Notice that here the term consistency could be misleading since we are supposing that the parameters of the model are known. We will consider the case of joint estimation of parameters and factors in the next section.

Indeed, given the factor model representation, and the definition of Λ_0, we have:

$$
\begin{aligned}
(\Lambda_0'\Lambda_0 + \bar{\psi}_0 I_r)^{-1}\Lambda_0'X_t &= (D_0 + \bar{\psi}_0 I_r)^{-1}\Lambda_0'X_t \\
&= (D_0 + \bar{\psi}_0 I_r)^{-1}\Lambda_0'(\Lambda_0 G_t + \xi_t) \\
&= (D_0 + \bar{\psi}_0 I_r)^{-1}D_0 G_t + (D_0 + \bar{\psi}_0 I_r)^{-1}\Lambda_0'\xi_t
\end{aligned}
$$

Under CR1, the first term converges to the unobserved common factors G_t, since $(D_0 + \bar{\psi}_0 I_r)^{-1}D_0 \to I_r$, as $n \to \infty$.

Moreover, the last term converges to zero in mean square since

$$
\begin{aligned}
&\mathrm{E}_{\Omega_0}\left[(D_0 + \bar{\psi}_0 I_r)^{-1}\Lambda_0'\xi_t\xi_t'\Lambda_0(D_0 + \bar{\psi}_0 I_r)^{-1}\right] \\
&= (D_0 + \bar{\psi}_0 I_r)^{-1}\Lambda_0'\Psi_0\Lambda_0(D_0 + \bar{\psi}_0 I_r)^{-1} \\
&\leq \lambda_{max}(\Psi_0)(D_0 + \bar{\psi}_0 I_r)^{-1}D_0(D_0 + \bar{\psi}_0 I_r)^{-1} \to 0 \text{ as } n \to \infty
\end{aligned}
$$

by assumptions $(CR2)$ and $(A4)$.

If we denote $G_{t/T,R1} = \mathrm{Proj}_{\Omega_0^{R1}}[G_t|X_s, s \leq T]$, we then have:

$$
G_{t/T,R1} - G_t = O_P\left(\frac{1}{\sqrt{n}}\right) \quad \text{as } n \to \infty
$$

Remark 3 Notice that in traditional factor models, where n is considered fixed, the factors are indeterminate and can only be approximated with an approximation error that depends inversely on the signal to noise variance ratio. The n large analysis shows that under suitable conditions, the approximation error goes to zero for n large.

This simple estimator is the most efficient if $\Omega_0 = \Omega_0^{R1}$. This is the model assumed in the Probabilistic Principal Components framework for i.i.d. data (static). However, if there are dynamics in the common factors $(A_0(L) \neq I_r)$ and if the idiosyncratic components have dynamics and are not spherical $(D_0(L) \neq \sqrt{\bar{\psi}_0}I_n)$ we still estimate consistently the unobserved common factors, as $n \to \infty$.

If the size of the idiosyncratic component is not the same across series, a more efficient estimator can be obtained by exploiting such heterogeneity by giving less weight to series with larger idiosyncratic component. Denoting $\Psi_{0d} = \mathrm{diag}(\psi_{0,11}, ..., \psi_{0,nn})$, this can be done by running the projection under the triple

$$
\Omega_0^{R2} = \left\{\Lambda_0, I_r, \Psi_{0d}^{1/2}\right\}
$$

Using the same kind of calculations as those we used in the previous case, with $\Psi_{0R} = \Psi_{0d}$ instead of $\Psi_{0R} = \bar{\psi}_0 I_n$, the following estimated factors are:

$$
\mathrm{Proj}_{\Omega_0^{R2}}[G_t|X_s, s \leq T] = \Lambda_0'(\Lambda_0\Lambda_0' + \Psi_{0d})^{-1}X_t = \left(\Lambda_0'\Psi_{0d}^{-1}\Lambda_0 + I_r\right)^{-1}\Lambda_0'\Psi_{0d}^{-1}X_t
$$

This estimator is thus obtained as the previous one, up to the fact that X_t and, of course, Λ_0 have been weighted, with weight given by $\sqrt{\psi_{0,11}}, ..., \sqrt{\psi_{0,nn}}$. This is the model assumed in the traditional (exact) Factor Analysis framework for i.i.d. data (static). It is then straightforward to obtain the same consistency result as in the previous case. If $G_{t/t,R2} := \mathrm{Proj}_{\Omega_0^{R2}}[G_t | X_s, s \leq T]$, then:

$$G_{t/T,R2} \overset{m.s.}{\longrightarrow} G_t \text{ and } G_{t/T,R2} - G_t = O_P\left(\frac{1}{\sqrt{n}}\right) \text{ as } n \to \infty$$

Further efficiency improvements could be obtained by non diagonal weighting scheme, i.e. by running the projection under the triple $\left\{\Lambda_0, I_r, \Psi_0^{1/2}\right\}$. This might be empirically relevant since, although limited asymptotically by assumption CR2, the idiosyncratic cross-sectional correlation may affect results in finite sample. We will not consider such projections since non diagonal weighting schemes raise identifiability problems in finite samples. Practically, they require the estimation of too many parameters and result in running out of degree of freedom in estimation (see next Section).

On the other hand, the estimators considered above do not take into consideration the dynamics of the factors and the idiosyncratic component. For this reason the factors are extracted by projecting only on contemporaneous observations. Since the model can be written in a state space form, the projection under more general dynamic structure can be computed using Kalman smoothing techniques.

Two particular cases in which the Kalman smoother can be used to exploit the dynamics of the common factors are:

$$\Omega_0^{R3} = \left\{\Lambda_0, A_0(L), \sqrt{\bar{\psi}_0} I_n\right\}$$

$$\Omega_0^{R4} = \left\{\Lambda_0, A_0(L), \Psi_{0d}^{1/2}\right\}$$

It is then possible, in this more general framework, to show the following result:

Proposition 1 Under assumptions A1, A2, A3', A4, A5, CR1, CR2 and CR3, if $G_{t/T,R} := \mathrm{Proj}_{\Omega_0^R}[G_t | X_s, s \leq T]$ with $R = R3$ or $R = R4$, then:

$$G_{t/T,R} - G_t = O_P\left(\frac{1}{\sqrt{n}}\right) \quad \text{as } n \to \infty$$

Under such parametrization, the computational complexity of the Kalman smoothing techniques depends mainly on the dimension of the transition equation which, under the parameterizations above, is independent of n and depends only on the number of the common factors.

In summary, the factors can be consistently estimated, as n become larger, by simple static projection of the observable on the factor loadings. However, efficiency improvements can be obtained by exploiting the cross-sectional heteroscedasticity of the

idiosyncratic components through weighted regressions (parametrization Ω_0^{R2}) and by exploiting the dynamics of the factors through the Kalman smoother (parametrizations $\Omega_0^{R3}, \Omega_0^{R4}$).

Individual idiosyncratic dynamics could also be taken into account when performing the projections. This would require to specify an autoregressive model for the idiosyncratic components or a reparameterization of the model as in Quah and Sargent (1992), to capture idiosyncratic dynamics by including lagged observable variable.

4 Estimation of the Parameters

The discussion in the previous section assumed that the parameters were known and focused on the extraction of the factors. In this section we will consider the problem of the estimation of the parameters as well as the resulting estimation of the factors.

The estimation of the full model is not feasible since it is not possible to fully parameterize parsimoniously the DGP of the idiosyncratic component since in most applications the cross-sectional items have no natural order. Moreover, models that explicitly take into account cross-correlation are not identified in general. In addition, the treatment of the idiosyncratic dynamics, even at the univariate level, is problematic since it can create computational problems.

However, as we have seen above, if the factor loadings were known, the factors could be consistently estimated, even if the projections were not computed under the correct specification. Does robustness with respect to misspecification still hold if the parameters are estimated?

Let us consider first the ML estimation under the approximating model $\Omega^{R1} = \left\{ \Lambda, I_r, \sqrt{\bar{\psi}} I_n \right\}$. The log-likelihood of the model is given by:

$$
\begin{aligned}
\mathcal{L}^T(\Lambda, I_r, \bar{\psi} I_n) \ &= -\tfrac{nT}{2} \log(2\pi) - \tfrac{T}{2} \log|\Sigma| - \tfrac{1}{2} \sum_{t=1}^{T} X_t' \Sigma^{-1} X_t \\
&= -\tfrac{nT}{2} \log(2\pi) - \tfrac{T}{2} (\log|\Sigma| + \operatorname{tr}\Sigma^{-1} S)
\end{aligned}
$$

where $S = \frac{1}{T} \sum_{t=1}^{T} X_t X_t'$ and $\Sigma = \Lambda\Lambda' + \bar{\psi} I_n$.

The model is identified under the normalization condition that $\Lambda'\Lambda$ is a diagonal matrix, with diagonal elements in decreasing order of magnitude. If we denote by $\hat{d}_j$ the j-th eigenvalue of S, in decreasing order of magnitude[5], by $\hat{p}_j$ the relative eigenvector and write $\hat{D}$ for the $(r \times r)$ diagonal matrix with diagonal elements $\hat{d}_j$, $j = 1 \ldots r$, and $\hat{P} := (\hat{p}_1, \ldots, \hat{p}_r)$, the associated maximum-likelihood estimates are given by[6]:

$$
\hat{\bar{\psi}}_{R1} = \frac{1}{n-r} \operatorname{trace}(S - \hat{D}); \qquad \hat{\Lambda}_{R1} = \hat{P} \left(\hat{D} - \hat{\bar{\psi}}_{R1} I_r \right)^{1/2}
$$

---

[5] It is always assumed that those eigenvalues are all distinct, in order to avoid useless mathematical complications. Under assumption A7, this will be asymptotically true, due to the fact that S converges to Σ_0

[6] See e.g. Lawley and Maxwell (1963) for a derivation of the first order conditions.

Denoting by $\hat{\Omega}^{R1} = \left\{ \hat{\Lambda}_{R1}, I_r, \sqrt{\hat{\bar{\psi}}_{R1}} I_n \right\}$ the model associated with the estimated parameters, we get:

$$
\begin{aligned}
\hat{G}_{t/T,R1} &= \text{Proj}_{\hat{\Omega}^{R1}}[G_t | X_s, s \leq T] = \left(\hat{\Lambda}'_{R1} \hat{\Lambda}_{R1} + \hat{\bar{\psi}}_{R1} I_r \right)^{-1} \hat{\Lambda}'_{R1} X_t \\
&= \hat{D}^{-1} \left(\hat{D} - \hat{\bar{\psi}}_{R1} I_r \right)^{1/2} \hat{P}' X_t
\end{aligned}
$$

It can then be shown (see corollary below) that $\hat{G}_{t/T,R1}$ is asymptotically equivalent to the normalized sample principal components $\hat{G}_t = \hat{D}^{-1/2} \hat{P}' X_t$.

Hence, principal components can be seen as an asymptotic equivalent of the Maximum Likelihood estimator for the factor loadings of the approximate factor model, in a situation in which the probability model is not correctly specified: the true model satisfies conditions CR1 to CR3, is dynamic and approximate, while we restrict the approximating model to be static and the idiosyncratic component to be spherical. This is what White (1982) named as Quasi Maximum Likelihood estimator. Properties of this estimator are studied in (Doz et al., 2005)

Under our set of assumptions, it can be shown that principal components give consistent estimators of the span of the common factors, and of associated factors loadings, when both the cross-section and the sample size go to infinity. This result has been shown by Forni et al. (2005a). Similar results, under alternative assumptions have been derived Bai (2003), Bai and Ng (2002) and Stock and Watson (2002a). However, we give our own proof of this result in appendix A.2, in order to make the paper self-contained and to prove the following propositions of this section.

Proposition 2 If assumptions (CR1) to (CR3), (A1) to (A5) hold, then Λ_0 can be defined[7] so as the following property holds:

$$
\hat{G}_t - G_t = O_P \left(\frac{1}{\sqrt{n}} \right) + O_P \left(\frac{1}{\sqrt{T}} \right), \quad \text{as } n, T \to \infty
$$

Consistency results can be reinterpreted as follows: "the bias arising from this misspecification of the data generating process of the idiosyncratic component and the dynamic properties of the factors is negligible if the cross-sectional dimension is large enough".

As suggested by Forni et al. 2005, the VAR coefficients $A_0(L)$ can be estimated by OLS regression of $\hat{G}_{t/T,R1}$, or equivalently of $\hat{G}_t$, on their own past. More precisely, the following OLS regression:

$$
\hat{G}_t = \hat{A}_1 \hat{G}_{t-1} + \dots + \hat{A}_p \hat{G}_{t-p} + \hat{w}_t
$$

gives consistent estimates of the $A_{0,k}$ matrices. The following proposition states the consistency results for the estimators of the loading and idiosyncratic matrices, as well as for the estimated VAR coefficient matrices.

---

[7]As Λ_0 is defined up to a sign change of its columns, and G_t is defined up to the sign of its components, the consistency result holds up to a given value of these signs.

Proposition 3 Under the same assumptions as in proposition 2, if $\hat{\Lambda} = \hat{P}\hat{D}^{1/2}$ is the estimator of Λ_0 associated to PCA and if $\hat{\Psi} = S - \hat{\Lambda}\hat{\Lambda}'$ the following properties hold:

i) For any i, j: $\hat{\lambda}_{ij} - \lambda_{0,ij} = O_P\left(\frac{1}{n}\right) + O_P\left(\frac{1}{\sqrt{T}}\right)$

ii) For any (i, j): $\hat{\psi}_{ij} - \psi_{0,ij} = O_P\left(\frac{1}{n}\right) + O_P\left(\frac{1}{\sqrt{T}}\right)$

iii) If $\hat{\Gamma}_{\hat{G}}(h)$ denotes the sample autocovariance of order h of the estimated principal components: $\hat{\Gamma}_{\hat{G}}(h) = \frac{1}{T-h}\sum_{t=h+1}^{T}\hat{G}_t\hat{G}'_{t-h}$, then for any h:

$$\hat{\Gamma}_{\hat{G}}(h) - \Phi_0(h) = O_P\left(\frac{1}{n}\right) + O_P\left(\frac{1}{\sqrt{T}}\right)$$

and the result is uniform in $h \in \mathbb{Z}$

(iv) For any $s = 0, ..., p$: $\hat{A}_s - A_{0,s} = O_P\left(\frac{1}{n}\right) + O_P\left(\frac{1}{\sqrt{T}}\right)$

Corollary 1 Under the same assumptions as in proposition 2:

i) $\hat{G}_{t/T,R1} - G_t = O_P\left(\frac{1}{\sqrt{n}}\right) + O_P\left(\frac{1}{\sqrt{T}}\right)$, as $n, T \to \infty$

ii) the properties which are stated in proposition 3 still hold if $\hat{\Lambda}$ is replaced by $\hat{\Lambda}_{R1}$ and if $\hat{\Psi}$ is replaced by $\hat{\Psi}_{R1} = S - \hat{\Lambda}_{R1}\hat{\Lambda}'_{R1}$.

The propositions and corollary above show that principal components are asymptotically equivalent to maximum likelihood estimators under a spherical assumption for the idiosyncratic component. Moreover, they show that, with principal components, we can estimate consistently not only the the common factor, but also the factor loadings, the variance of the idiosyncratic component and the VAR filter of the common factors. The latter estimates can then be used to get a more efficient estimates of the common factors. Let us denote by $\hat{\psi}_{ii}$ and $\hat{A}(L)$ such estimates, then we can obtain a new estimate of the factors by computing new projections. We consider three cases:

1) Weighted Principal components

$$\hat{G}_{t/T,R2} = \text{Proj}_{\hat{\Omega}^{R2}}[G_t | X_s, s \leq T]$$

were $\hat{\Omega}^{R2} = \left\{\hat{\Lambda}, I_r, \text{diag}(\hat{\psi}_{11}, ..., \hat{\psi}_{nn})^{1/2}\right\}$. This is asymptotically equivalent to principal components on weighted observations where the weights are the inverse of the standard deviation of the estimated idiosyncratic components. This estimator has been considered in Forni and Reichlin (2000), Boivin and Ng (2004), Forni, Hallin, Lippi and Reichlin (2005).

2)Principal components and Kalman filtering with no reweighting

$$\hat{G}_{t/T,R3} = \text{Proj}_{\hat{\Omega}^{R3}}[G_t|X_s, s \leq T]$$

were $\hat{\Omega}^{R3} = \left\{\hat{\Lambda}, \hat{A}(L), \sqrt{\hat{\bar{\psi}}}I_n\right\}$. This estimator does not take into account the non-sphericity of the idiosyncratic components, but only exploits the common factor dynamics.

3) Principal components and Kalman filtering with reweighting

$$\hat{G}_{t/T,R4} = \text{Proj}_{\hat{\Omega}^{R4}}[G_t|X_s, s \leq T]$$

were $\hat{\Omega}^{R4} = \left\{\hat{\Lambda}, \hat{A}(L), \text{diag}(\hat{\psi}_{11}, ..., \hat{\psi}_{nn})^{1/2}\right\}$. This projection is estimated using the Kalman filter proposed by Giannone, Reichlin and Small (2005) and applied by Giannone, Reichlin and Sala, 2005. Such estimator exploits both the non-sphericity of the idiosyncratic component and the dynamics of the common factors.

Consistency of this three new estimates of the common factors, follows from the consistency of the principal components. First, it is straightforward to extend the proof of Propositions 2 and 3 in order to obtain the consistency of the weighted PCA estimates (see appendix A.3). Second the consistency of the two Kalman filter estimates stems from the consistency of the associated unfiltered estimates, and the proofs are identical in the $\hat{\Omega}^{R3}$ and $\hat{\Omega}^{R4}$ frameworks. If we denote by Ω_0^R the model under consideration, and by $\hat{\Omega}^R$ the associated set of parameters, obtained at the first step of the estimation procedure, so that:

- if $R = R3$: $\hat{\Omega}^R = \left\{\hat{\Lambda}, \hat{A}(L), \sqrt{\hat{\bar{\psi}}}I_n\right\}$

- if $R = R4$: $\hat{\Omega}^R = \left\{\hat{\Lambda}, \hat{A}(L), \text{diag}(\hat{\psi}_{0,11}, ..., \hat{\psi}_{0,nn})^{1/2}\right\}$

then, the consistency of the associated estimates can be stated in the following proposition:

Proposition 4 Denote $\hat{G}_{t/T,R} = \widehat{\text{Proj}}_{\hat{\Omega}^R}[G_t|X_s, s \leq T]$ with $R = R3$ or $R4$.
If $\limsup \frac{T}{n^3} = O(1)$, the following result holds under assumptions (CR1) to (CR3), (A1), (A2), (A3'), (A4) and (A5):

$$\hat{G}_{t/T,R} - G_t = O_P\left(\frac{1}{\sqrt{n}}\right) + O_P\left(\frac{1}{\sqrt{T}}\right) \quad \text{as } n, T \rightarrow \infty$$

The procedure outlined above consists in computing the common factors through principal components. We then use the common factors to estimate the parameters. With this set of parameters we then reestimate the common factors according to the selected approximating model, in order to improve the efficiency of the estimates. What if we iterate such procedure? From the new estimated factors, we can estimate a new set of parameters which in turn can then be used to reestimate the common factors and so on. If, at each iteration the least squares estimates of the parameters are computed using expected sufficient statistics, then such iterative procedure is nothing that the EM algorithm by Dempster and Rubin (1977) and introduced in small scale dynamic factor models by Engle and Watson (1981). Quah and Sargent (1992) used such algorithm for large cross-sections, but their approach was disregarded in subsequent literature. The algorithm is very powerful since at each step the likelihood increases, and hence, under regularity conditions, it converges to the Maximum Likelihood solution. For details about the estimation with state space models see Engle and Watson (1981) and Quah and Sargent (1992). The algorithm is feasible for large cross-sections for two reasons. First, as stressed above, its complexity is mainly due to the number of factors, which in our framework is independent of the size of the cross-section and typically very small. Second, since the algorithm is initialized with consistent estimates (Principal Component), the number of iterations required for convergence is expected to be limited, in particular when the cross-section is large. The asymptotic properties of quasi maximum likelihood estimates for large cross-section and under an approximate factor structure is developed in Doz et al. (2005).

5 Empirics

In this section we run a simulation study to asses the performances of our estimator. The model from which we simulate is standard in the literature. A similar model has been used, for example, in Stock and Watson (2002a).

Let us define it below (in what follows, in order to have simpler notations, we drop the zero subscript for the true value of the parameters which we had previously used to study the consistency of the estimates).

$x_{it} = \sum_{j=1}^{r} \lambda_{ij}^* f_{jt} + \xi_{it}, i = 1, .., n$, in vector notation $X_t = \Lambda^* F_t + \xi_t$

$A(L)F_t = u_t$, with u_t i.i.d. $\mathcal{N}(0, I_r); i, j = 1, ..., r$

$D(L)\xi_t = v_t$ with v_t i.i.d. $\mathcal{N}(0, \mathcal{T})$

$a_{ij}(L) = \begin{cases} 1 - \rho L & \text{if} \quad i = j \\ 0 & \text{if} \quad i \neq j \end{cases}$

$d_{ij}(L) = \begin{cases} \sqrt{\alpha_i}(1 - dL) & \text{if} \quad i = j \\ 0 & \text{if} \quad i \neq j \end{cases} ; i, j = 1, ..., n$

λ_{ij}^* i.i.d. $\mathcal{N}(0, 1), i = 1, ..., n; j = 1, .., r$

$$\alpha_i = \frac{\beta_i}{1-\beta_i} \frac{1}{T} \sum_{t=1}^{T} \left(\sum_{j=1}^{r} \lambda_{ij}^* f_{jt} \right)^2 \text{ with } \beta_i \text{ i.i.d. } \mathcal{U}([u, 1-u])$$

$$\mathcal{T}_{ij} = \tau^{|i-j|} \frac{1}{1-d^2}, i, j = 1, ..., n$$

Notice that we allow for instantaneous cross-correlation between the idiosyncratic elements. Since $\mathcal{T}$ is a Toeplitz matrix, the cross-correlation among idiosyncratic elements is limited and it is easily seen that Assumption A (ii) is satisfied. The coefficient τ controls for the amount of cross-correlation. The exact factor model corresponds to $\tau = 0$.

The coefficient β_i is the ratio between the variance of the idiosyncratic component, ξ_{it}, and the variance of the common component, $\sum_{j=1}^{r} \lambda_{ij}^* f_{jt}$. The is also known as the noise to signal ratio. In our simulation this ratio is uniformly distributed with an average of 50%. If $u = .5$ then the standardized observations have cross-sectionally homoscedastic idiosyncratic components.

Notice that if $\tau = 0, d = 0$, our approximating model is well specified (with the usual notational convention that $0^0 = 1$)and hence the approximating model R4 is well specified. If $\tau = 0$, $d = 0$, $\rho = 0$, we have a static exact factor model with heteroscedastic idiosyncratic component and model R2 is correctly specified while principal components are not the most efficient estimator. Finally, if $\tau = 0$, $d = 0$, $u = 1/2$, we have a spherical, static factor model on standardized variables, situation in which the approximating model R1 is correctly specified and principal components on standardized variables provide the most efficient, maximum likelihood, estimates.

We generate the model for different sizes of the cross-section, $n = 10, 25, 50, 100$, and for sample size $T = 50$, 100. We perform 2500 Monte-Carlo repetitions. We draw 50 times the parameters β_i, $i = 1, ..., n$, and λ_{ij}^*, $i = 1, ..., n$; $j = 1, .., r$. Then, for each draw of the parameters, we generate the 50 times the shocks u_t and ξ_t.

As stressed in the introduction, an advantage of having a parameterized model is that it is possible to extract the common factors from panel at the end of the sample due to the unsynchronous data releases (see Giannone et al., 2004, 2005, for an application to real time nowcasting and forecasting output and inflation). To study the performance of our models, for each sample size T and cross-sectional dimension n, we generate the data under the following pattern of data availability,

$$x_{it} \text{ available for } t = 1, ..., T - j \text{ if } i \leq (j+1)\frac{n}{5}$$

that is all the variables are observed for $t = 1, ..., T - 4$, we name this a balanced panel; 80% of the data are available at time $T - 3$; 60% are available at time $T - 2$; 40% are available at time $T - 1$; 20% are available at time T.

At each repetition, the parameters $\hat{\Lambda}, \hat{A}(L)$ and $\hat{\psi}_{ii}, i = 1, ..., n$ are estimated on the balanced part of the panel, $x_{it}, i = 1, ..., n, t = 1, ..., T - 4$. Data are standardized so as to have mean zero and variance equal to one. Such standardization is typically applied in empirical analysis since principal components are not scale invariant.

We consider the factor extraction under the approximating models studied in the previous section and summarized below.

$$\hat{\Omega}^{R1} = \left\{ \hat{\Lambda}, I_r, \sqrt{\hat{\bar{\psi}}} I_n \right\}$$

$$\hat{\Omega}^{R2} = \left\{ \hat{\Lambda}, I_r, \mathrm{diag}(\hat{\psi}_{11}, ..., \hat{\psi}_{nn})^{1/2} \right\}$$

$$\hat{\Omega}^{R3} = \left\{ \hat{\Lambda}, \hat{A}(L), \sqrt{\hat{\bar{\psi}}} I_n \right\}$$

$$\hat{\Omega}^{R4} = \left\{ \hat{\Lambda}, \hat{A}(L), \mathrm{diag}(\hat{\psi}_{11}, ..., \hat{\psi}_{nn})^{1/2} \right\}.$$

We compute the estimates by applying the Kalman smoother using the estimated parameters: $\hat{G}_{t/T,R} = \widehat{\mathrm{Proj}}_{\hat{\Omega}^R}[G_t | X_s, s \leq T]$, for $R = R1, R2, R3, R4$. The pattern of data availability can be taken into account when estimating the common factors, by modifying the idiosyncratic variance when performing the projections:

- if x_{it} is available, then $\mathrm{E}\xi_{it}^2 = \hat{\bar{\psi}}$ for the projections $R1, R3$ and $\mathrm{E}\xi_{it}^2 = \hat{\psi}_{ii}$ is x_{it} for the projections $R2, R4$

- if x_{it} is not available, then $\mathrm{E}\xi_{it}^2 = \infty$ is x_{it}

The estimates of the common factor can hence be computed running the Kalman smoother with time varying parameters (see Giannone et al., 2004, 2005).

We measure the performance of the different estimators as:

$$\Delta_{t,R} = \mathrm{Trace} \left(F_t - \hat{Q}'_R \hat{G}_{t/T,R} \right) \left(F_t - \hat{Q}'_R \hat{G}_{t/T,R} \right)'$$

where $\hat{Q}_R$ is the OLS coefficient from the regression of F_t on $\hat{G}_{t/T,R}$ estimated using observations up to time $T - 4$, that is: $\hat{Q}_R = \sum_{t=1}^{T-4} F_t \hat{G}'_{t/T,R} \left(\sum_{t=1}^{T-4} \hat{G}_{t/T,R} \hat{G}'_{t/T,R} \right)^{-1}$. This OLS regression is performed since the common factors are identified only up to a rotation. Indeed, we know from the previous sections that $\hat{G}_{t/T,R}$ is a consistent estimator of $G_t = Q'F_t$, where Q is a rotation matrix such that $Q'\Lambda^{*\prime}\Lambda Q$ is diagonal, with diagonal terms in decreasing order. Thus, it can be easily checked that, as $E(F_t F_t') = I_r$, $\hat{Q}_R$ is a consistent estimator of:

$$\mathrm{plim} \left(\tfrac{1}{T} \sum_{t=1}^{T-4} F_t \hat{G}'_{t/T,R} \right) \left(\tfrac{1}{T} \sum_{t=1}^{T-4} \hat{G}_{t/T,R} \hat{G}'_{t/T,R} \right)^{-1}$$
$$= \mathrm{plim} \left(\tfrac{1}{T} \sum_{t=1}^{T-4} F_t G_t' \right) \left(\tfrac{1}{T} \sum_{t=1}^{T-4} G_t G_t' \right)^{-1}$$
$$= \mathrm{plim} \left(\tfrac{1}{T} \sum_{t=1}^{T-4} F_t F_t' \right) Q Q' \left(\tfrac{1}{T} \sum_{t=1}^{T-4} G_t G_t' \right)^{-1} Q$$
$$= Q$$

so that $\hat{Q}'_R \hat{G}_{t/T,R}$ is a consistent estimator of F_t.

Table 1:

j	T=50						T=100			
	$n=5$	$n=10$	$n=25$	$n=50$	$n=100$		$n=5$	$n=10$	$n=25$	$n=50$	$n=100$
$\Delta_{j,R4}$: evaluation of the Kalman filter with cross-sectional heteroscedasticity											
-4	0.45	0.35	0.30	0.29	0.28		0.34	0.23	0.19	0.18	0.17
-3	0.45	0.35	0.30	0.28	0.28		0.36	0.24	0.19	0.18	0.17
-2	0.47	0.36	0.30	0.28	0.27		0.37	0.26	0.20	0.18	0.17
-1	0.50	0.39	0.31	0.29	0.27		0.40	0.29	0.21	0.18	0.17
0	0.57	0.44	0.34	0.30	0.28		0.48	0.35	0.25	0.21	0.19
$\Delta_{j,R4}/\Delta_{j,R1}$: relative performances of simple Principal components											
-4	0.97	0.97	0.98	0.99	0.99		0.95	0.94	0.96	0.98	0.99
-3	0.95	0.95	0.97	0.98	0.99		0.93	0.93	0.96	0.98	0.98
-2	0.92	0.93	0.97	0.98	0.99		0.90	0.91	0.95	0.97	0.98
-1	0.88	0.89	0.95	0.97	0.98		0.84	0.85	0.92	0.95	0.97
0	0.80	0.82	0.90	0.95	0.98		0.73	0.75	0.85	0.92	0.96
$\Delta_{j,R4}/\Delta_{j,R2}$: relative performances of Weighted Principal components											
-4	0.98	0.98	0.99	1.00	1.00		0.96	0.98	0.99	1.00	1.00
-3	0.96	0.97	0.99	1.00	1.00		0.95	0.96	0.99	1.00	1.00
-2	0.94	0.96	0.99	1.00	1.00		0.93	0.95	0.99	1.00	1.00
-1	0.90	0.92	0.98	0.99	1.00		0.86	0.89	0.97	0.99	1.00
0	0.81	0.84	0.94	0.98	1.00		0.75	0.78	0.91	0.97	1.00
$\Delta_{j,R4}/\Delta_{j,R3}$: relative performances of the Kalman filter with cross-sectional homoscedasticity											
-4	1.00	0.99	0.99	0.99	1.00		1.00	0.97	0.97	0.98	0.99
-3	0.99	0.99	0.98	0.99	0.99		1.00	0.97	0.96	0.98	0.98
-2	0.99	0.98	0.98	0.98	0.99		0.98	0.96	0.96	0.97	0.98
-1	0.98	0.98	0.98	0.98	0.99		0.97	0.96	0.96	0.96	0.97
0	0.97	0.98	0.99	0.97	0.98		0.96	0.96	0.94	0.95	0.96

We compute the distance for each repetition and then compute the averages $(\bar{\Delta}_{t,R})$.

Table 1 summarizes the results of the Montecarlo experiment for one common factors $r=1$ and the following specification: $\rho=.9$, $d=.5$, $\tau=.5$, $u=.1$.

We report the following measures of performance for the last 5 observations to analyze how data availability affects the estimates. The Kalman filter with cross-sectional heteroscedasticity $R4$ is used as a benchmark and we report $\bar{\Delta}_{T-j,R4}$. The smaller the measure, the more accurate are the estimates of the common factors. In addition, we report $\bar{\Delta}_{T-j,R4}/\bar{\Delta}_{T-j,R1}$, $\bar{\Delta}_{T-j,R4}/\bar{\Delta}_{T-j,R2}$, $\bar{\Delta}_{T-j,R4}/\bar{\Delta}_{T-j,R3}$. A number smaller then 1 indicates that the projection under $R4$ is more accurate.

Results show five main features:

1. For any j fixed, $\bar{\Delta}_{T-j,R4}$ decreases as n and T increase, that is the precision of the estimated common factors increases with the size of the cross-section n and the sample size T.

2. For any combination of n and T, $\bar{\Delta}_{T-j,R4}$ increases as j decreases, reflecting the fact that the more numerous are the available data, the higher the precision of the common factor estimates.

3. $\bar{\Delta}_{T-j,R4} < \bar{\Delta}_{T-j,R3} < \bar{\Delta}_{T-j,R2} < \bar{\Delta}_{T-j,R1}$, for all n,T,j. This result indicates

that the less miss-specified is the model used for the projection, the more accurate are the estimated factors. This suggests that taking into account cross-sectional heteroskedasticity and the dynamic of the common factors helps extracting the common factor.

4. For any combination of n and T, $\bar{\Delta}_{T-j,R4}/\bar{\Delta}_{T-j,R}$ (for $R = R1$ to $R3$) decreases as j decreases. That is, the efficiency improvement is more relevant when it is harder to extract the factors (i.e. the less numerous are the available data).

5. As n, T increase $\bar{\Delta}_{T-j,R4}/\bar{\Delta}_{T-j,R}$ tends to one, for all j and for $R = R1$ to $R3$; that is the performance of the different estimators tends to become very similar. This reflects the fact that all the estimates are consistent for large cross-sections.

Summarizing, the two steps estimator of approximate factor models works well in finite sample. Because it models explicitly dynamics and cross-sectional heteroscedasticity, it dominates principal components. Efficiency improvements are relevant when the factor extraction is difficult, that is, when the available data are less numerous.

6   Conclusions

We have shown (n, T) consistency and rates of common factors estimated via a two step procedure whereby, in the first step, the parameters of a dynamic approximate factor model are first estimated by a OLS regression of the variables on principal components and, in the second step, given the estimated parameters, the factors are estimated by the Kalman smoother.

This procedure allows to take into account, in the estimation of the factors, both factor dynamics and idiosyncratic heteroskedasticity, features that are likely to be relevant in the panels of data typically used in empirical applications in macroeconomics.

Our empirical analysis shows a slight improvement for n small which however disappears in a panel of medium size ($n = 70$).

The parametric approach studied in this paper provides the theoretical justification for two applications of factor models in large cross-sections: treatment of unbalanced panels (Giannone, Reichlin, and Sala, 2004; Giannone, Reichlin, and Small, 2005) and estimation of shocks in structural factor models (Giannone, Reichlin, and Sala, 2004). The approach can also be used to evaluate estimation uncertainty around the common factors as in the papers just cited.

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A Appendix

A.1 Consistency of Kalman Smoothing: population results

Notice first that, as stressed in Section 2, G_t has the same stochastic properties as F_t. In particular: $\sum_{h=-\infty}^{+\infty} \|\Phi_0(h)\| < +\infty$ and $\sum_{h=-\infty}^{+\infty} \Phi_0(h)$ is an invertible matrix.

Let us denote :

$$G_{t/T,R} = \text{Proj}_{\Omega_0^R}[G_t|X_s, s \leq T] = \sum_{s=1}^{T} M_s\left(G_t, \Omega_0^R\right) X_s$$

We want to show that $G_{t/T,R}$ is a consistent estimate of G_t and that this property is true even if Ω_0^R is misspecified due to the fact that the true matrix Ψ_0 is a non-diagonal matrix and the idiosyncratic components are autocorrelated.

We use the following notations:

- $P_0 = D_0^{-1/2}\Lambda_0$, so that P_0 and Λ_0 span the same subspaces and $P_0'P_0 = I_r$,

- $\mathbf{X}_T = (X_1', ... X_T')'$, $\mathbf{G}_T = (G_1', ... G_T')'$, $\mathbf{Z}_T = (\xi_1', ... \xi_T')'$,

- $\mathbf{M}\left(G_t, \Omega_0^R\right) = \left(M_1\left(G_t, \Omega_0^R\right), ..., M_T\left(G_t, \Omega_0^R\right)\right)$,

- E denotes the expectation of a random variable, under the true model Ω_0,

- $E_{\Omega_0^R}$ denotes the expectation of a random variable, when Ω_0^R is the model which is considered,

- when (Y_t) is a stationary process: $\Gamma_Y(h) = E(Y_tY_{t-h}')$ and $\Gamma_{Y,R}(h) = E_{\Omega_0^R}(Y_tY_{t-h}')$.

With these notations:
$$\mathbf{X}_T = (I_T \otimes \Lambda_0)\mathbf{G}_T + \mathbf{Z}_T,$$

$$\mathbf{M}\left(G_t, \Omega_0^R\right) = E_{\Omega_0^R}(G_t\mathbf{X}_T')(E_{\Omega_0^R}(\mathbf{X}_T\mathbf{X}_T'))^{-1},$$

$$\text{and}: G_{t/T,R} = \mathbf{M}\left(G_t, \Omega_0^R\right)\mathbf{X}_T = E_{\Omega_0^R}(G_t\mathbf{X}_T')(E_{\Omega_0^R}(\mathbf{X}_T\mathbf{X}_T'))^{-1}\mathbf{X}_T.$$

Before proving the proposition, we finally introduce a last notation, in order to simplify the calculations. When (Y_t) is a stationary process and $\mathbf{Y}_T = (Y_1', ... Y_T')'$, we denote:

$$\Sigma_Y = E(\mathbf{Y}_T\mathbf{Y}_T') \text{ and } \Sigma_{Y,R} = E_{\Omega_0^R}(\mathbf{Y}_T\mathbf{Y}_T')$$

so that we can finally write:

$$G_{t/T,R} = E_{\Omega_0^R}(G_t\mathbf{X}_T')\Sigma_{X,R}^{-1}\mathbf{X}_T$$

Notice that, when $R = R3$ or $R = R4$, the DGP of (G_t) is correctly specified, so that $\Sigma_G = \Sigma_{G,R}$. On the contrary, $\Sigma_{Z,R}$ is not equal to Σ_Z.

Before proving the proposition, we need the following results:

Lemma 1 *Under assumptions A1, A2, A3', A4, A5, CR1, CR2 and CR3, the following properties hold for $R = R3$ and $R = R4$:*

i) $E_{\Omega_0^R}(G_t \mathbf{X}'_T) = (\Gamma_{G,R}(t-1)\Lambda'_0, ..., \Gamma_{G,R}(t-T)\Lambda'_0) = \Sigma_{G,R}(I_T \otimes \Lambda'_0)$

ii) $\Sigma_{X,R} = (I_T \otimes \Lambda_0)\Sigma_{G,R}(I_T \otimes \Lambda'_0) + I_T \otimes \Psi_{0R}$
    where $\Psi_{0,R3} = \Gamma_{\xi,R3}(0) = \bar{\psi}_0 I_n$ *and* $\Psi_{0,R4} = \Gamma_{\xi,R4}(0) = \mathrm{diag}(\psi_{0,11}, ...\psi_{0,nn})$

iii) $\Sigma_{G,R} = \Sigma_G$, $\|\Sigma_G\| = O(1)$ *and* $\|\Sigma_G^{-1}\| = O(1)$

iv) $\|\Sigma_Z\| = O(1)$

Proof

i) As $X_t = \Lambda_0 G_t + \xi_t$, we get: $\mathbf{X}_T = (I_T \otimes \Lambda_0)\mathbf{G}_T + \mathbf{Z}_T$.

It then immediately follows from assumptions (A3) and (A4) that:

$$E_{\Omega_0^R}(G_t \mathbf{X}'_T) = E_{\Omega_0^R}(G_t \mathbf{G}'_T)(I_T \otimes \Lambda'_0)$$

where: $E_{\Omega_0^R}(G_t \mathbf{G}'_T)(I_T \otimes \Lambda'_0) = (\Gamma_{G,R}(t-1)\Lambda'_0, ..., \Gamma_{G,R}(t-T)\Lambda'_0)$.

ii) It also follows from assumptions (A3) and (A4) that:

$$\Sigma_{X,R} = E_{\Omega_0^R}(\mathbf{X}_T \mathbf{X}'_T) = (I_T \otimes \Lambda_0)E_{\Omega_0^R}(\mathbf{G}_T \mathbf{G}'_T)(I_T \otimes \Lambda'_0) + E_{\Omega_0^R}(\mathbf{Z}_T \mathbf{Z}'_T)$$
$$= (I_T \otimes \Lambda_0)\Sigma_{G,R}(I_T \otimes \Lambda'_0) + \Sigma_{Z,R}$$

Further, as (ξ_t) is supposed to be a white noise in both Ω^{R3} and Ω^{R4} specifications, we also have

$$\Sigma_{Z,R} = I_T \otimes \Gamma_{\xi,R}(0) = I_T \otimes \Psi_{0R}$$

iii) We have already noticed that, when $R = R3$ or $R = R4$, the model is correctly specified for (G_t), so that $\Sigma_{GR} = \Sigma_G$.

For any $\omega \in [-\pi, +\pi]$, let us now denote by $S_G(\omega)$ the spectral density matrix of (G_t) calculated in ω. In order to show the two announced properties, it is sufficient to show that if:

$$m = \mathrm{Min}_{\omega \in [-\pi,+\pi]}\lambda_{min}(S_G(\omega)) \quad \text{and} \quad M = \mathrm{Max}_{\omega \in [-\pi,+\pi]}\lambda_{max}(S_G(\omega))$$

then: $2\pi m \leq \lambda_{min}(\Sigma_G)$ and $2\pi M \geq \lambda_{max}(\Sigma_G)$.

Indeed, as we know, from assumption (A3), that $m > 0$ and $M < \infty$, the result will then follow from the fact that:

$$\|\Sigma_G\| = \lambda_{max}(\Sigma_G) \text{ and } \|(\Sigma_G)^{-1}\| = \frac{1}{\lambda_{min}(\Sigma_G)}.$$

In order to show this property, we generalize to the r-dimensionnal process (G_t) the proof which is given by Brockwell and Davis, 1987 (proposition 4.5.3) in the univariate

case.

If $\mathbf{x} = (x_1', ...x_T')'$ is a non-random vector of $\mathbb{R}^{rT}$ such that: $\|x\|^2 = \sum_{t=1}^{T} \|x_t\|^2 = 1$, we can write:

$$\mathbf{x}'\Sigma_G\mathbf{x} = \sum_{t=1}^{T}\sum_{\tau=1}^{T} x_t'\Gamma_G(t-\tau)x_\tau = \sum_{t=1}^{T}\sum_{\tau=1}^{T} x_t'\Phi_0(t-\tau)x_\tau$$

We thus get:

$$
\begin{aligned}
\mathbf{x}'\Sigma_G\mathbf{x} &= \sum_{1\leq t,\tau\leq T} x_t' \left(\int_{-\pi}^{+\pi} S_G(\omega)e^{-i\omega(t-\tau)}d\omega \right) x_\tau \\
&= \int_{-\pi}^{+\pi} \left(\sum_{1\leq t,\tau\leq T} x_t'S_G(\omega)x_\tau e^{-i\omega(t-\tau)} \right) d\omega \\
&= \int_{-\pi}^{+\pi} \left(\sum_{1\leq t\leq T} x_t'e^{-i\omega t} \right) S_G(\omega) \left(\sum_{1\leq \tau\leq T} x_\tau e^{i\omega\tau} \right) d\omega \\
&\in \left[m\int_{-\pi}^{+\pi} \| \sum_{1\leq t\leq T} x_t'e^{-i\omega t}\|^2 d\omega, M\int_{-\pi}^{+\pi} \| \sum_{1\leq t\leq T} x_t'e^{-i\omega t}\|^2 d\omega \right]
\end{aligned}
$$

Now:

$$
\begin{aligned}
\int_{-\pi}^{+\pi} \| \sum_{1\leq t\leq T} x_t'e^{-i\omega t}\|^2 d\omega &= \int_{-\pi}^{+\pi} \left(\sum_{1\leq t,\tau\leq T} x_t'e^{-i\omega t}x_\tau e^{-i\omega\tau} \right) d\omega \\
&= \sum_{1\leq t,\tau\leq T}\int_{-\pi}^{+\pi} x_t'x_\tau e^{-i\omega(t-\tau)}d\omega = 2\pi\sum_{1\leq t\leq T} x_t'x_t = 2\pi\sum_{1\leq t\leq T} \|x_t\|^2 = 2\pi
\end{aligned}
$$

We thus obtain that any eigenvalue of Σ_G belongs to $[2\pi m, 2\pi M]$, which gives the announced result.

iv) For any $\omega \in [-\pi, +\pi]$, let us now denote by $S_\xi(\omega)$ the spectral density matrix of (ξ_t) calculated in ω. If $x = (x_1, ...x_n)'$ is a non-random vector of $\mathbb{C}^n$ such that: $\|x\|^2 = x'\bar{x} = 1$, we have:

$$x'S_\xi(\omega)\bar{x} = \frac{1}{2\pi}\sum_{h\in\mathbb{Z}} x'\Gamma_\xi(h)e^{i\omega h}\bar{x} = \frac{1}{2\pi}\sum_{h\in\mathbb{Z}} x'\Psi_0(h)e^{i\omega h}\bar{x}$$

so that:

$$|x'S_\xi(\omega)\bar{x}| \leq \frac{1}{2\pi}\sum_{h\in\mathbb{Z}} |x'\Psi_0(h)\bar{x}| \leq \frac{1}{2\pi}\sum_{h\in\mathbb{Z}} \|\Psi_0(h)\|$$

It then results from assumption (CR2) that for any n, and for any $\omega \in [-\pi, +\pi]$: $\lambda_{max}S_\xi(\omega) \leq \frac{1}{2\pi}\bar{\lambda}$ so that

$$\text{Max}_{\omega\in[-\pi,+\pi]}\lambda_{max}(S_\xi(\omega)) \leq \frac{1}{2\pi}\bar{\lambda}$$

Applying the same result as in (iii), we then get:

$$\|\Sigma_{\xi,R}\| \leq \bar{\lambda}$$

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Proof of Proposition 1

From lemma 1 (ii), we know that: $\Sigma_{X,R} = (I_T \otimes \Lambda_0)\Sigma_{G,R}(I_T \otimes \Lambda_0') + \Sigma_{Z,R}$.

Using the same kind of formula as the formula we have used to calculate Σ_0^{-1}, it can be easily checked that:

$$\Sigma_{X,R}^{-1} = \Sigma_{Z,R}^{-1} - \Sigma_{Z,R}^{-1}(I_T \otimes \Lambda_0)\left(\Sigma_{G,R}^{-1} + (I_T \otimes \Lambda_0')\Sigma_{Z,R}^{-1}(I_T \otimes \Lambda_0)\right)^{-1}(I_T \otimes \Lambda_0')\Sigma_{Z,R}^{-1}$$

Using the fact that $\Sigma_{Z,R}^{-1} = I_T \otimes \Psi_{0,R}^{-1}$, we then get:

$$\begin{aligned}
(I_T \otimes \Lambda_0')\Sigma_{X,R}^{-1} &= I_T \otimes \Lambda_0'\Psi_{0,R}^{-1} - I_T \otimes \Lambda_0'\Psi_{0,R}^{-1}\Lambda_0(\Sigma_{G,R}^{-1} + I_T \otimes \Lambda_0'\Psi_{0,R}^{-1}\Lambda_0)^{-1}I_T \otimes \Lambda_0'\Psi_{0,R}^{-1} \\
&= \left(\Sigma_{G,R}^{-1} + I_T \otimes \Lambda_0'\Psi_{0,R}^{-1}\Lambda_0 - I_T \otimes \Lambda_0'\Psi_{0,R}^{-1}\Lambda_0\right)(\Sigma_{G,R}^{-1} + I_T \otimes \Lambda_0'\Psi_{0,R}^{-1}\Lambda_0)^{-1}I_T \otimes \Lambda_0'\Psi_{0,R}^{-1} \\
&= \Sigma_{G,R}^{-1}(\Sigma_{G,R}^{-1} + I_T \otimes \Lambda_0'\Psi_{0,R}^{-1}\Lambda_0)^{-1}I_T \otimes \Lambda_0'\Psi_{0R}^{-1}
\end{aligned}$$

Then, as:

$$G_{t/T,R} = E_{\Omega_0^R}(G_t\mathbf{G}_T')(I_T \otimes \Lambda_0')\Sigma_{X,R}^{-1}\mathbf{X}_T$$

we get:

$$G_{t/T,R} = E_{\Omega_0^R}(G_t\mathbf{G}_T')\Sigma_{G,R}^{-1}(\Sigma_{G,R}^{-1} + I_T \otimes \Lambda_0'\Psi_{0R}^{-1}\Lambda_0)^{-1}(I_T \otimes \Lambda_0'\Psi_{0R}^{-1})\mathbf{X}_T$$

Finally, if we denote by $\mathbf{U}_t'$ the $(r \times rT)$ matrix defined by: $\mathbf{U}_t' = (0, ...I_r, 0...0)$, we have: $E_{\Omega_0^R}(G_t\mathbf{G}_T') = E_{\Omega_0^R}(\mathbf{U}_t'\mathbf{G}_T\mathbf{G}_T') = \mathbf{U}_t'\Sigma_{G,R}$ so that:

$$\begin{aligned}
G_{t/T,R} &= \mathbf{U}_t'\Sigma_{G,R}\Sigma_{G,R}^{-1}(\Sigma_{G,R}^{-1} + I_T \otimes \Lambda_0'\Psi_{0R}^{-1}\Lambda_0)^{-1}(I_T \otimes \Lambda_0'\Psi_{0R}^{-1})\mathbf{X}_T \\
&= \mathbf{U}_t'(\Sigma_{G,R}^{-1} + I_T \otimes \Lambda_0'\Psi_{0R}^{-1}\Lambda_0)^{-1}(I_T \otimes \Lambda_0'\Psi_{0R}^{-1})\mathbf{X}_T
\end{aligned}$$

Before proving the proposition, let us first recall a relation, which we use in that proof as well as in others. If A and B are two square invertible matrices, it is possible to write write: $B^{-1} - A^{-1} = B^{-1}(A - B)A^{-1}$, so that the relation:

$$(A + H)^{-1} = A^{-1} - (A + H)^{-1}HA^{-1} \qquad (R)$$

also gives a Taylor expansion of the inversion operator at order zero when H is small with respect to A.

Using relation (R), and denoting $M_0 = \Lambda_0'\Psi_{0R}^{-1}\Lambda_0$, we then get:

$$\begin{aligned}
G_{t/T,R} &= \mathbf{U}_t'\left(I_T \otimes M_0^{-1} - (\Sigma_{G,R}^{-1} + I_T \otimes M_0)^{-1}\Sigma_{G,R}^{-1}(I_T \otimes M_0^{-1})\right)(I_T \otimes \Lambda_0'\Psi_{0R}^{-1})\mathbf{X}_T \\
&= \mathbf{U}_t'(I_T \otimes M_0^{-1}\Lambda_0'\Psi_{0R}^{-1})\mathbf{X}_T - \mathbf{U}_t'(\Sigma_{G,R}^{-1} + I_T \otimes M_0)^{-1}\Sigma_{G,R}^{-1}(I_T \otimes M_0^{-1}\Lambda_0'\Psi_{0R}^{-1})\mathbf{X}_T \\
&= M_0^{-1}\Lambda_0'\Psi_{0R}^{-1}X_t - \mathbf{U}_t'(\Sigma_{G,R}^{-1} + I_T \otimes M_0)^{-1}\Sigma_{G,R}^{-1}(I_T \otimes M_0^{-1}\Lambda_0'\Psi_{0R}^{-1})\mathbf{X}_T
\end{aligned}$$

Let us denote $G^1_{t/T,R}$ the first term of the previous summation. We can write:

$$\begin{aligned}
G^1_{t/T,R} &= (\Lambda'_0 \Psi^{-1}_{0R} \Lambda_0)^{-1} \Lambda'_0 \Psi^{-1}_{0R} X_t \\
&= (\Lambda'_0 \Psi^{-1}_{0R} \Lambda_0)^{-1} \Lambda'_0 \Psi^{-1}_{0R} (\Lambda_0 G_t + \xi_t) \\
&= G_t + (\Lambda'_0 \Psi^{-1}_{0R} \Lambda_0)^{-1} \Lambda'_0 \Psi^{-1}_{0R} \xi_t
\end{aligned}$$

with:

$$\begin{aligned}
E[\|(\Lambda'_0 \Psi^{-1}_{0R} \Lambda_0)^{-1} \Lambda'_0 \Psi^{-1}_{0R} \xi_t\|^2] &= E\left[tr(\Lambda'_0 \Psi^{-1}_{0R} \Lambda_0)^{-1} \Lambda'_0 \Psi^{-1}_{0R} \xi_t \xi'_t \Psi^{-1}_{0R} \Lambda_0 (\Lambda'_0 \Psi^{-1}_{0R} \Lambda_0)^{-1}\right] \\
&= tr\left((\Lambda'_0 \Psi^{-1}_{0R} \Lambda_0)^{-1} \Lambda'_0 \Psi^{-1}_{0R} \Psi_0 \Psi^{-1}_{0R} \Lambda_0 (\Lambda'_0 \Psi^{-1}_{0R} \Lambda_0)^{-1}\right)
\end{aligned}$$

As $\Psi^{-1/2}_{0R} \Psi_0 \Psi^{-1/2}_{0R} \leq \frac{\lambda_{max}(\Psi_0)}{\lambda_{min}(\Psi_{0R})} I_n$, we get:

$$(\Lambda'_0 \Psi^{-1}_{0R} \Lambda_0)^{-1} \Lambda'_0 \Psi^{-1}_{0R} \Psi_0 \Psi^{-1}_{0R} \Lambda_0 (\Lambda'_0 \Psi^{-1}_{0R} \Lambda_0)^{-1} \leq \frac{\lambda_{max}(\Psi_0)}{\lambda_{min}(\Psi_{0R})} (\Lambda'_0 \Psi^{-1}_{0R} \Lambda_0)^{-1}$$

so that: $E\left[\|(\Lambda'_0 \Psi^{-1}_{0R} \Lambda_0)^{-1} \Lambda'_0 \Psi^{-1}_{0R} \xi_t\|^2\right] = O_P\left(\frac{1}{n}\right)$ by assumptions (CR1) and (CR2).

We have thus obtained:

$$G^1_{t/T,R} = G_t + O_P\left(\frac{1}{\sqrt{n}}\right)$$

Turning to the second term of the summation, it can in turn be decomposed in two parts. Indeed, as $\mathbf{X}_T = (I_T \otimes \Lambda_0)\mathbf{G}_T + \mathbf{Z}_T$, we can write:

$$\mathbf{U}'_t (\Sigma^{-1}_{G,R} + I_T \otimes M_0)^{-1} \Sigma^{-1}_{G,R} (I_T \otimes M^{-1}_0 \Lambda'_0 \Psi^{-1}_{0R}) \mathbf{X}_T = G^2_{t/T,R} + G^3_{t/T,R}$$

with:

$$\begin{aligned}
G^2_{t/T,R} &= \mathbf{U}'_t (\Sigma^{-1}_{G,R} + I_T \otimes M_0)^{-1} \Sigma^{-1}_{G,R} (I_T \otimes M^{-1}_0 \Lambda'_0 \Psi^{-1}_{0R})(I_T \otimes \Lambda_0)\mathbf{G}_T \\
&= \mathbf{U}'_t \left(\Sigma^{-1}_{G,R} + I_T \otimes M_0\right)^{-1} \Sigma^{-1}_{G,R} (I_T \otimes (\Lambda'_0 \Psi^{-1}_{0R} \Lambda_0)^{-1} \Lambda'_0 \Psi^{-1}_{0R} \Lambda_0)\mathbf{G}_T \\
&= \mathbf{U}'_t \left(\Sigma^{-1}_{G,R} + I_T \otimes M_0\right)^{-1} \Sigma^{-1}_{G,R} \mathbf{G}_T
\end{aligned}$$

and:

$$G^3_{t/T,R} = \mathbf{U}'_t (\Sigma^{-1}_{G,R} + I_T \otimes M_0)^{-1} \Sigma^{-1}_{G,R} (I_T \otimes M^{-1}_0 \Lambda'_0 \Psi^{-1}_{0R}) \mathbf{Z}_T$$

We can write:

$$\begin{aligned}
&E\left[\|G^2_{t/T,R}\|^2\right] \\
&= tr\left[\mathbf{U}'_t \left(\Sigma^{-1}_{G,R} + I_T \otimes M_0\right)^{-1} \Sigma^{-1}_{G,R} E\left[\mathbf{G}_T \mathbf{G}'_T\right] \Sigma^{-1}_{G,R} \left(\Sigma^{-1}_{G,R} + I_T \otimes M_0\right)^{-1} \mathbf{U}_t\right]
\end{aligned}$$

As $E\left[\mathbf{G}_T \mathbf{G}'_T\right] = \Sigma_G = \Sigma_{G,R}$, we then get:

$$E\left[\|G^2_{t/T,R}\|^2\right] = tr\left[\mathbf{U}'_t \left(\Sigma^{-1}_{G,R} + I_T \otimes M_0\right)^{-1} \Sigma^{-1}_{G,R} \left(\Sigma^{-1}_{G,R} + I_T \otimes M_0\right)^{-1} \mathbf{U}_t\right]$$

As $\Sigma_{G,R}^{-1} \leq \lambda_{max}(\Sigma_{G,R}^{-1})I_{rT}$, with $\lambda_{max}(\Sigma_{G,R}^{-1}) = \|\Sigma_{G,R}^{-1}\|$, we have:

$$\left(\Sigma_{G,R}^{-1} + I_T \otimes M_0\right)^{-1} \Sigma_{G,R}^{-1} \left(\Sigma_{G,R}^{-1} + I_T \otimes M_0\right)^{-1} \leq \|\Sigma_{G,R}^{-1}\| \left(\Sigma_{G,R}^{-1} + I_T \otimes M_0\right)^{-2}$$

Now: $\Sigma_{G,R}^{-1} + I_T \otimes M_0 \geq I_T \otimes M_0$ so that: $\left(\Sigma_{G,R}^{-1} + I_T \otimes M_0\right)^{-1} \leq I_T \otimes M_0^{-1}$.

We then get:

$$E\left[\|G_{t/T,R}^2\|^2\right] \leq \|\Sigma_{G,R}^{-1}\| tr\left[\mathbf{U}_t'(I_T \otimes M_0^{-2})\mathbf{U}_t\right] = \|\Sigma_{G,R}^{-1}\| tr\left[M_0^{-2}\right] = O\left(\frac{1}{n^2}\right)$$

It then follows from assumptions (CR1) and (CR3) and from lemma 1 (iii) that:

$$G_{t/T,R}^2 = O_P\left(\frac{1}{n}\right)$$

If we use the same type of properties that we have used for the study of $G_{t/T,R}^2$, we can

write:

$$E\left[\|G_{t/T,R}^3\|^2\right] = tr\left[\begin{array}{c} \mathbf{U}_t'(\Sigma_{G,R}^{-1} + I_T \otimes M_0)^{-1}\Sigma_{G,R}^{-1}(I_T \otimes M_0^{-1}\Lambda_0'\Psi_{0R}^{-1})\Sigma_Z \\ \times (I_T \otimes \Psi_{0R}^{-1}\Lambda_0 M_0^{-1})\Sigma_{G,R}^{-1}(\Sigma_{G,R}^{-1} + I_T \otimes M_0)^{-1}\mathbf{U}_t \end{array}\right]$$

We thus get:

$$E\left[\|G_{t/T,R}^3\|^2\right]$$
$$\leq \|\Sigma_{G,R}^{-1}(I_T \otimes M_0^{-1}\Lambda_0'\Psi_{0R}^{-1})\Sigma_Z(I_T \otimes M_0^{-1}\Lambda_0'\Psi_{0R}^{-1})\Sigma_{G,R}^{-1}\| tr\left[\mathbf{U}_t'(\Sigma_{G,R}^{-1} + I_T \otimes M_0)^{-2}\mathbf{U}_t\right]$$
$$\leq \|\Sigma_{G,R}^{-1}\|^2 \|(I_T \otimes M_0^{-1}\Lambda_0'\Psi_{0R}^{-1})\|^2 \|\Sigma_Z\| tr\left[\mathbf{U}_t'(\Sigma_{G,R}^{-1} + I_T \otimes M_0)^{-2}\mathbf{U}_t\right]$$

From lemma 1 (iii) and (iv) we know that $\|\Sigma_{G,R}^{-1}\| = O(1)$ and $\|\Sigma_Z\| = O(1)$.

Further, using assumptions (CR1) and (CR2), we can write, as before:

$$tr\left[\mathbf{U}_t'(\Sigma_{G,R}^{-1} + I_T \otimes M_0)^{-2}\mathbf{U}_t\right] \leq tr\left[\mathbf{U}_t'(I_T \otimes M_0)^{-2}\mathbf{U}_t\right] = tr(M_0^{-2}) = O\left(\frac{1}{n^2}\right)$$

and:
$$\|(I_T \otimes M_0^{-1}\Lambda_0'\Psi_{0R}^{-1})\| = \|M_0^{-1}\Lambda_0'\Psi_{0R}^{-1}\| = O\left(\frac{1}{\sqrt{n}}\right)$$

It then follows that: $E\left[\|G_{t/T,R}^3\|^2\right] = O\left(\frac{1}{n^3}\right)$, so that:

$$G_{t/T,R}^3 = O_P\left(\frac{1}{n\sqrt{n}}\right)$$

which completes the proof of the proposition.

A.2 Consistency of PCA

Lemma 2 Under assumptions (CR1) to (CR3), (A1) to (A5), the following properties hold, as $n, T \to \infty$:

i) $\frac{1}{n}\|S - \Lambda_0\Lambda_0'\| = O\left(\frac{1}{n}\right) + O_P\left(\frac{1}{\sqrt{T}}\right)$

ii) $\frac{1}{n}\|\hat{D} - D_0\| = O\left(\frac{1}{n}\right) + O_P\left(\frac{1}{\sqrt{T}}\right)$

iii) $n\|\hat{D}^{-1} - D_0^{-1}\| = O_P\left(\frac{1}{n}\right) + O_P\left(\frac{1}{\sqrt{T}}\right)$

iv) $D_0\hat{D}^{-1} = I_r + O_p\left(\frac{1}{n}\right) + O_p\left(\frac{1}{\sqrt{T}}\right)$

Proof

i) $\frac{1}{n}\|S - \Lambda_0\Lambda_0'\| \leq \frac{1}{n}\|S - \Sigma_0\| + \frac{1}{n}\|\Sigma_0 - \Lambda_0\Lambda_0'\|.$

As $\Sigma_0 = \Lambda_0^*\Lambda_0^{*\prime} + \Psi_0 = \Lambda_0\Lambda_0' + \Psi_0$, we have by assumption (CR2) :

$$\frac{1}{n}\|\Sigma_0 - \Lambda_0\Lambda_0'\| = \frac{1}{n}\|\Psi_0\| = O\left(\frac{1}{n}\right)$$

We also have:

$$S = \frac{1}{T}\sum_{t=1}^{T} X_t X_t' = \Lambda_0\frac{1}{T}\sum_{t=1}^{T} G_t G_t' \Lambda_0' + \Lambda_0\frac{1}{T}\sum_{t=1}^{T} G_t \xi_t' + \frac{1}{T}\sum_{t=1}^{T} \xi_t G_t' \Lambda_0' + \frac{1}{T}\sum_{t=1}^{T} \xi_t \xi_t'$$

so that:

$$\frac{1}{n}(S - \Sigma_0) = \frac{1}{n}\Lambda_0\left(\frac{1}{T}\sum_{t=1}^{T} G_t G_t' - I_r\right)\Lambda_0' + \frac{1}{n}\left(\Lambda_0\frac{1}{T}\sum_{t=1}^{T} G_t \xi_t' + \frac{1}{T}\sum_{t=1}^{T} \xi_t G_t' \Lambda_0'\right)$$
$$+ \frac{1}{n}\left(\frac{1}{T}\sum_{t=1}^{T} \xi_t \xi_t' - \Psi_0\right)$$

Then, using assumptions (A3) and (CR2) and a multivariate extension of the proof given in the univariate case by Brockwell and Davies (1991, pp226-227), it is possible to show that:

$$\mathrm{E}\left(\|\frac{1}{T}\sum_{t=1}^{T} G_t G_t' - I_r\|^2\right) = O\left(\frac{1}{T}\right) \text{ and } \mathrm{E}\left(\|\frac{1}{T}\sum_{t=1}^{T} \xi_t \xi_t' - \Psi_0\|^2\right) = O\left(\frac{n^2}{T}\right)$$

so that:

$$\|\frac{1}{T}\sum_{t=1}^{T} G_t G_t' - I_r\| = O_P\left(\frac{1}{\sqrt{T}}\right) \text{ and } \|\frac{1}{T}\sum_{t=1}^{T} \xi_t \xi_t' - \Psi_0\| = O_P\left(\frac{n}{\sqrt{T}}\right)$$

It also follows from these assumptions that: $\|\frac{1}{T}\sum_{t=1}^{T} G_t \xi_t'\| = O_P\left(\frac{\sqrt{n}}{\sqrt{T}}\right)$. Indeed, we can write:

$$\|\frac{1}{T}\sum_{t=1}^{T} G_t \xi_t'\|^2 = \|\frac{1}{T}\sum_{t,s} G_t \xi_t' \xi_s G_s'\| \leq tr\left(\frac{1}{T}\sum_{t,s} G_t \xi_t' \xi_s G_s'\right)$$

As (G_t) and (ξ_t) are two independent processes, we have:

$$
\begin{aligned}
\mathrm{E}\left[tr\left(\tfrac{1}{T}\sum_{t,s}G_t\xi_t'\xi_s G_s'\right)\right] &= tr\left(\tfrac{1}{T}\sum_{t,s}\mathrm{E}(\xi_t'\xi_s)\mathrm{E}(G_t G_s')\right) \\
&= \tfrac{1}{T^2}\sum_{t,s}tr\left(\Psi_0(t-s)\right)tr\left(\Phi_0(s-t)\right) \\
&\leq \tfrac{1}{T^2}\sum_{t,s}|tr\left(\Psi_0(t-s)\right)||tr\left(\Phi_0(s-t)\right)| \\
&\leq \tfrac{nr}{T^2}\sum_{t,s}\|\Psi_0(t-s)\|\|\Phi_0(s-t)\| \\
&= \tfrac{nr}{T}\sum_{h=-T+1}^{T-1}(1-\tfrac{|h|}{T})\|\Psi_0(h)\|\|\Phi_0(-h)\| \\
&\leq \tfrac{nr}{T}\sum_{h\in\mathbb{Z}}\|\Psi_0(h)\|\|\Phi_0(-h)\| \\
&\leq \tfrac{nr}{T}Max_{h\in\mathbb{Z}}\|\Psi_0(h)\|\sum_{h\in\mathbb{Z}}\|\Phi_0(h)\|
\end{aligned}
$$

We thus obtain: $\mathrm{E}\left[\|\tfrac{1}{T}\sum_{t=1}^{T}G_t\xi_t'\|^2\right] = O_P\left(\tfrac{n}{T}\right)$ and the result follows.

ii) $\hat{D}$ is the diagonal matrix of the r first eigenvalues of S, in decreasing order. D_0 is a diagonal matrix which is equal to $\Lambda_0'\Lambda_0$. It is then also equal to the diagonal matrix of the r first eigenvalues of $\Lambda_0\Lambda_0'$ in decreasing order.

Further, if we denote by $\lambda_1(A) \geq \lambda_2(A) \geq \cdots \geq \lambda_n(A)$ the ordered eigenvalues of a symmetric matrix A, we can write, from Weyl theorem, that for any $j = 1,...r$:

$$
|\lambda_j(S) - \lambda_j(\Lambda_0\Lambda_0')| \leq \|S - \Lambda_0\Lambda_0'\|
$$

(see for instance, Horn and Johnson (1990) p.181). The result then immediately follows from (i).

iii) By assumptions (CR1) and (CR3), we know that $\tfrac{1}{n}D_0 = O(1)$ and that $\left(\tfrac{1}{n}D_0\right)^{-1} = O(1)$. It then results from (ii) that the eigenvalues of $\tfrac{1}{n}\hat{D}$ and of $\left(\tfrac{1}{n}\hat{D}\right)^{-1}$ are $O_P(1)$, so that $\tfrac{1}{n}\hat{D} = O_P(1)$ and $\left(\tfrac{1}{n}\hat{D}\right)^{-1} = O_P(1)$. The result the follows from (ii) and from the decomposition:

$$
n\left(\hat{D}^{-1} - D_0^{-1}\right) = \left(\tfrac{1}{n}\hat{D}\right)^{-1}\tfrac{1}{n}\left(\hat{D}-D_0\right)\left(\tfrac{1}{n}D_0\right)^{-1} = O_P\left(\tfrac{1}{n}\right) + O_P\left(\tfrac{1}{\sqrt{T}}\right)
$$

iv) $D_0\hat{D}^{-1} = I_r + \tfrac{D_0}{n}\left[\left(\tfrac{\hat{D}}{n}\right)^{-1} - \left(\tfrac{D_0}{n}\right)^{-1}\right]$.

The result then follows from (iii) and assumption CR3.

Lemma 3 Let us denote $\hat{A} = \hat{P}'P_0$, with $\hat{A} = (\hat{a}_{ij})_{1\leq i,j\leq r}$.

The following properties hold:

i) $\hat{a}_{ij} = O_P\left(\tfrac{1}{n}\right) + O_P\left(\tfrac{1}{\sqrt{T}}\right)$ for $i \neq j$

ii) $\hat{a}_{ii}^2 = 1 + O_P\left(\tfrac{1}{n}\right) + O_P\left(\tfrac{1}{\sqrt{T}}\right)$ for $i = 1,...r$

Proof

i) As $S\hat{P} = \hat{P}\hat{D}$ we have $\hat{P} = S\hat{P}\hat{D}^{-1}$ and:

$$\hat{P}'P_0 = \hat{D}^{-1}\hat{P}'SP_0 = \hat{D}^{-1}\hat{P}'\left(S - \Lambda_0\Lambda_0'\right)P_0 + \hat{D}^{-1}\hat{P}'\Lambda_0\Lambda_0'P_0$$

As $\Lambda_0 = P_0 D_0^{1/2}$, and $P_0'P_0 = I_r$, we have: $\Lambda_0\Lambda_0'P_0 = P_0 D_0$. We then get:

$$\hat{P}'P_0 = \left(\frac{\hat{D}}{n}\right)^{-1}\hat{P}'\left(\frac{S - \Lambda_0\Lambda_0'}{n}\right)P_0 + \left(\frac{\hat{D}}{n}\right)^{-1}\hat{P}'P_0\left(\frac{D_0}{n}\right)$$

As we saw in lemma 1, assumptions (CR1) and (CR3) imply that $\frac{D_0}{n}$ and $\left(\frac{D_0}{n}\right)^{-1}$ are $O(1)$ and that $\frac{\hat{D}}{n}$ and $\left(\frac{\hat{D}}{n}\right)^{-1}$ are $O_P(1)$. As $\hat{P}'\hat{P} = I_r$ and $P_0'P_0 = I_r$, it follows that $\hat{P}'P_0 = O_P(1)$. Thus, lemma 1 (i) and (iii) imply that:

$$\hat{P}'P_0 = O_P\left(\frac{1}{n}\right) + O_P\left(\frac{1}{\sqrt{T}}\right) + \left(\frac{D_0}{n}\right)^{-1}\hat{P}'P_0\left(\frac{D_0}{n}\right).$$

or equivalently that:

$$\hat{A} = D_0^{-1}\hat{A}D_0 + O_P\left(\frac{1}{n}\right) + O_P\left(\frac{1}{\sqrt{T}}\right).$$

For any i and j the previous relation states that:

$$\hat{a}_{ij} = \frac{d_{0,jj}}{d_{0,ii}}\hat{a}_{ij} + O_P\left(\frac{1}{n}\right) + O_P\left(\frac{1}{\sqrt{T}}\right)$$

For $i \neq j$, we assume, from assumption (A7), that $d_{0,jj} \neq d_{0,ii}$. We then obtain:

$$\hat{a}_{ij} = O_P\left(\frac{1}{n}\right) + O_P\left(\frac{1}{\sqrt{T}}\right) \text{ for } i \neq j.$$

ii) To study the asymptotic behavior of $\hat{a}_{ii}$, let us now use the relation

$$\hat{D} = \hat{P}'S\hat{P}$$

which implies, together with lemma 1 (i), that:

$$\frac{\hat{D}}{n} = \hat{P}'\frac{S}{n}\hat{P} = \hat{P}'\frac{\Lambda_0\Lambda_0'}{n}\hat{P} + O_P\left(\frac{1}{n}\right) + O_P\left(\frac{1}{\sqrt{T}}\right)$$

or, equivalently, that:

$$\frac{\hat{D}}{n} = \hat{P}'P_0\frac{D_0}{n}P_0'\hat{P} + O_P\left(\frac{1}{n}\right) + O_P\left(\frac{1}{\sqrt{T}}\right)$$

It then follows from lemma 1 (ii) that:

$$\frac{D_0}{n} = \hat{P}'P_0\frac{D_0}{n}P_0'\hat{P} + O_P\left(\frac{1}{n}\right) + O_P\left(\frac{1}{\sqrt{T}}\right)$$

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or equivalently that:

$$\frac{D_0}{n} = \hat{A}\frac{D_0}{n}\hat{A}' + O_P\left(\frac{1}{n}\right) + O_P\left(\frac{1}{\sqrt{T}}\right)$$

Thus, for $i = 1, ...r$:

$$\frac{d_{0,ii}}{n} = \sum_{k=1}^{r}\frac{d_{0,kk}}{n}\hat{a}_{ik}^2 + O_P\left(\frac{1}{n}\right) + O_P\left(\frac{1}{\sqrt{T}}\right)$$

and:

$$\frac{d_{0,ii}}{n}\left(1 - \hat{a}_{ii}^2\right) = \sum_{k\neq i}\frac{d_{0,kk}}{n}\hat{a}_{ik}^2 + O_P\left(\frac{1}{n}\right) + O_P\left(\frac{1}{\sqrt{T}}\right)$$

From result (i), we know that $\hat{a}_{ik} = O_P\left(\frac{1}{n}\right) + O_P\left(\frac{1}{\sqrt{T}}\right)$ for $i \neq k$. As $\frac{D_0}{n} = O_P(1)$, it then follows that:

$$\hat{a}_{ii}^2 = 1 + O_P\left(\frac{1}{n}\right) + O_P\left(\frac{1}{\sqrt{T}}\right) \text{ for } i = 1, ...r$$

Lemma 4 Under assumptions (CR1) to (CR3), (A1) to (A5), P_0 and $\hat{P}$ can be defined so as the following properties hold, as $n, T \rightarrow \infty$:

(i) $\hat{P}'P_0 = I_r + O_P\left(\frac{1}{n}\right) + O_P\left(\frac{1}{\sqrt{T}}\right)$

(ii) $\|\hat{P} - P_0\|^2 = O_P\left(\frac{1}{n}\right) + O_P\left(\frac{1}{\sqrt{T}}\right)$

(iii) $\tau_{in}'(\hat{\Lambda} - \Lambda_0) = O_P\left(\frac{1}{n}\right) + O_P\left(\frac{1}{\sqrt{T}}\right), \quad i = 1, ..., n$

where τ_{in} the ith denotes the ith vector of the canonical basis in $\mathbb{R}^n$.

Proof

i) We have seen before that P_0 is uniquely defined up to a sign change of each of its columns, and that this implies that G_t is uniquely defined for any t up to a sign change of each of its components. As $\hat{P}$ is also defined up to a sign change of its columns, it is thus possible to suppose that P_0 and $\hat{P}$ are chosen such that the diagonal terms of $\hat{A} = \hat{P}'P_0$ are positive. In such a case, lemma 2 (ii) implies that:

$$\hat{a}_{ii} = 1 + O_P\left(\frac{1}{n}\right) + O_P\left(\frac{1}{\sqrt{T}}\right) \text{ for } i = 1, ...r$$

We then obtain from lemma 2 (i) that: $\hat{P}'P_0 = I_r + O_P\left(\frac{1}{n}\right) + O_P\left(\frac{1}{\sqrt{T}}\right)$.

ii) Let $x \in \mathbb{R}^n$ a non-random vector such that $\|x\| = 1$. As $\hat{P}'\hat{P} = I_r$ and $P_0'P_0 = I_r$ we have:

$$x'(\hat{P} - P_0)'(\hat{P} - P_0)x = x'(2I_r - \hat{P}'P_0 - P_0'\hat{P})x$$

It then follows from (i) that $x'(\hat{P} - P_0)'(\hat{P} - P_0)x = O_P\left(\frac{1}{n}\right) + O_P\left(\frac{1}{\sqrt{T}}\right)$. As this is true for any $x \in \mathbb{R}^n$, it then follows that

$$\|\hat{P} - P_0\|^2 = O_P\left(\frac{1}{n}\right) + O_P\left(\frac{1}{\sqrt{T}}\right)$$

iii) We have $\hat{P} = S\hat{P}\hat{D}^{-1}$ and $\Sigma_0 = P_0 D_0 P_0' + \Psi_0$, so that

$$
\begin{aligned}
\tau_{in}'(\hat{\Lambda} - \Lambda_0) &= \tau_{in}'(\hat{P}\hat{D}^{1/2} - P_0 D_0^{1/2}) \\
&= \tau_{in}'(S\hat{P}\hat{D}^{-1/2} - P_0 D_0^{1/2}) \\
&= \tau_{in}'\left((S - \Sigma_0)\hat{P}\hat{D}^{-1/2} + (P_0 D_0 P_0' + \Psi_0)\hat{P}\hat{D}^{-1/2} - P_0 D_0^{1/2}\right) \\
&= \tau_{in}'(S - \Sigma_0)\hat{P}\hat{D}^{-1/2} + \tau_{in}'\Psi_0\hat{P}\hat{D}^{-1/2} + \tau_{in}' P_0 D_0 \left(P_0'\hat{P} - D_0^{-1/2}\hat{D}^{1/2}\right)\hat{D}^{-1/2}
\end{aligned}
$$

In order to study the first term, let us first notice that: $\|\tau_{in}'(S - \Sigma_0)\| = \left(\sum_{j=1}^n (s_{ij} - \sigma_{0,ij})^2\right)^{1/2}$.

Using the same arguments as in the proof of Lemma 2 (i), we have

$$\mathrm{E}\|\tau_{in}'(S - \Sigma_0)\|^2 \leq \sum_{j=1}^n \mathrm{E}(s_{ij} - \sigma_{0,ij})^2 = O\left(\frac{n}{T}\right)$$

so that $\tau_{in}'(S - \Sigma_0) = O_P\left(\frac{\sqrt{n}}{\sqrt{T}}\right)$.

As $\hat{P}'\hat{P} = I_r$, we know that $\hat{P} = O_P(1)$. Then, using $\hat{D}^{-1/2} = O_P\left(\frac{1}{\sqrt{n}}\right)$, it follows that:

$$\tau_{in}'(S - \Sigma_0)\hat{P}\hat{D}^{-1/2} = O_P\left(\frac{1}{\sqrt{T}}\right).$$

Turning to the second term, we have: $\|\tau_{in}'\Psi_0\| \leq \|\Psi_0\| = O(1)$, by assumption (CR2). As $\hat{P} = O_P(1)$ and $\hat{D}^{-1/2} = O_P\left(\frac{1}{\sqrt{n}}\right)$, we get:

$$\tau_{in}'\Psi_0\hat{P}\hat{D}^{-1/2} = O_P\left(\frac{1}{\sqrt{n}}\right)$$

Finally, $\tau_{in}' P_0 D_0 \left(P_0'\hat{P} - D_0^{-1/2}\hat{D}^{1/2}\right)\hat{D}^{-1/2} = \tau_{in}'\Lambda_0 D_0^{1/2}\left(P_0'\hat{P} - D_0^{-1/2}\hat{D}^{1/2}\right)\hat{D}^{-1/2}$.

As $Vx_{it} = \|\tau_{in}'\Lambda_0\|^2 + \psi_{0,ii}$, it follows from assumption (A2) that $\tau_{in}'\Lambda_0 = O(1)$.

Further, $\left(P_0'\hat{P} - D_0^{-1/2}\hat{D}^{1/2}\right) = O_P\left(\frac{1}{\sqrt{n}}\right) + O_P\left(\frac{1}{\sqrt{T}}\right)$ by lemma 2(iv) and lemma 4 (i). As $\hat{D}^{-1/2} = O_P\left(\frac{1}{\sqrt{n}}\right)$ and $D_0^{1/2} = O\left(\sqrt{n}\right)$, it then follows that:

$$\tau_{in}' P_0 D_0 \left(P_0'\hat{P} - D_0^{-1/2}\hat{D}^{1/2}\right)\hat{D}^{-1/2} = O_P\left(\frac{1}{\sqrt{n}}\right) + O_P\left(\frac{1}{\sqrt{T}}\right)$$

which completes the proof.

Proof of proposition 2

We can write:

$$\begin{aligned}
\hat{G}_t - G_t &= \hat{D}^{-1/2}\hat{P}'X_t - G_t \\
&= \hat{D}^{-1/2}\hat{P}'\left(\Lambda_0 G_t + \xi_t\right) - G_t \\
&= \left(\hat{D}^{-1/2}\hat{P}'P_0 D_0^{1/2} - I_r\right)G_t + \xi_t \\
&= \hat{D}^{-1/2}\left(\hat{P}'P_0 - \hat{D}^{1/2}D_0^{-1/2}\right)D_0^{1/2}G_t + \hat{D}^{-1/2}\hat{P}'\xi_t
\end{aligned}$$

Lemma 2 (iv) and lemma 4 (i) give: $\hat{P}'P_0 - \hat{D}^{1/2}D_0^{-1/2} = O_P\left(\frac{1}{\sqrt{n}}\right) + O_P\left(\frac{1}{\sqrt{T}}\right)$.

Then, applying lemma 2 (iv) a second time, and using the fact that $G_t = O_P(1)$, we get:

$$\hat{D}^{-1/2}\left(\hat{P}'P_0 - \hat{D}^{1/2}D_0^{-1/2}\right)D_0^{1/2}G_t = O_P\left(\frac{1}{\sqrt{n}}\right) + O_P\left(\frac{1}{\sqrt{T}}\right).$$

In order to study $\hat{D}^{-1/2}\hat{P}'\xi_t$, let us first decompose ξ_t as: $\xi_t = P_0 P_0'\xi_t + P_{0\perp}P_{0\perp}'\xi_t$ where $P_{0\perp}$ is a $(n \times (n-r))$ matrix whose columns form an orthonormal basis of the orthogonal space of P_0. We then obtain:

$$\hat{D}^{-1/2}\hat{P}'\xi_t = \hat{D}^{-1/2}\hat{P}'P_0 P_0'\xi_t + \hat{D}^{-1/2}\hat{P}'P_{0\perp}P_{0\perp}'\xi_t.$$

First, let us notice that $P_0'\xi_t = O_P(1)$ and that $P_{0\perp}'\xi_t = O_P(\sqrt{n})$.
Indeed, we can write:

$$E\left(\|P_0'\xi_t\|\right)^2 = E\left(\xi_t'P_0 P_0'\xi_t\right) = E\left(tr\left(P_0'\xi_t\xi_t'P_0\right)\right) = tr\left(P_0'\Psi_0 P_0\right) \leq r\lambda_1\left(\Psi_0\right) = O(1)$$

and $E\left(\|P_{0\perp}'\xi_t\|\right)^2 = E\left(tr\left(P_{0\perp}'\xi_t\xi_t'P_{0\perp}\right)\right) = tr\left(P_{0\perp}'\Psi_0 P_{0\perp}\right) \leq (n-r)\lambda_1\left(\Psi_0\right) = O(n).$

As lemma 2 (iii) implies that: $\hat{D}^{-1} = O_P(\frac{1}{n})$, we then get from lemma 4 (i) that:

$$\hat{D}^{-1/2}\hat{P}'P_0 P_0'\xi_t = O_P\left(\frac{1}{\sqrt{n}}\right).$$

In order to study the second term, let us first show that:

$$\hat{P}'P_{0\perp} = O_P\left(\frac{1}{n}\right) + O_P\left(\frac{1}{\sqrt{T}}\right)$$

Indeed, if we use: $\hat{P} = S\hat{P}\hat{D}^{-1}$, we can write:

$$\hat{P}'P_{0\perp} = \hat{D}^{-1}\hat{P}'SP_{0\perp}$$

As P_0 and Λ_0 have the same range, $P_{0\perp}'\Lambda_0 = 0$, so that we also have:

$$\hat{P}'P_{0\perp} = \hat{D}^{-1}\hat{P}'(S - \Lambda_0\Lambda_0')P_{0\perp}$$

We then get:

$$\hat{P}' P_{0\perp} = \left(\frac{\hat{D}}{n}\right)^{-1} \hat{P}' \frac{S - \Lambda_0 \Lambda_0'}{n} P_{0\perp}.$$

As $P_{0\perp}' P_{0\perp} = I_{n-r}$, we have: $P_{0\perp} = O(1)$. It then follows from lemma 2 (i) and (ii) that:

$$\hat{P}' P_{0\perp} = O_P\left(\frac{1}{n}\right) + O_P\left(\frac{1}{\sqrt{T}}\right)$$

Then, as $\hat{D}^{-1/2} = O_P\left(\frac{1}{\sqrt{n}}\right)$, and $P_{0\perp}' \xi_t = O_P(\sqrt{n})$, it follows that:

$$\hat{D}^{-1/2} \hat{P}' P_{0\perp} P_{0\perp}' \xi_t = O_P\left(\frac{1}{n}\right) + O_P\left(\frac{1}{\sqrt{T}}\right)$$

which completes the proof of the proposition.

Proof of Proposition 3

i) Since $\hat{\lambda}_{ij} - \lambda_{0,ij} = \tau'_{in}\left(\hat{\Lambda} - \Lambda_0\right)\tau_{jn}$, the result is an immediate consequence of Lemma 4 (iii).

ii) As $\hat{\Psi} = S - \hat{\Lambda}\hat{\Lambda}'$, and as $\Psi_0 = \Sigma_0 - \Lambda_0\Lambda'_0$, the result follows from (i), and from the fact that for any (i,j): $s_{ij} - \sigma_{0,ij} = O_P\left(\frac{1}{n}\right) + O_P\left(\frac{1}{\sqrt{T}}\right)$.

iii) Consider the sample autocovariance of the estimated principal components

$$\hat{\Gamma}_{\hat{G}}(h) = \frac{1}{T-h}\sum_{t=h+1}^{T}\hat{G}_t\hat{G}'_{t-h} = \hat{D}^{-1/2}\hat{P}'S(h)\hat{P}\hat{D}^{-1/2}$$

with $S(h) = \frac{1}{T-h}\sum_{t=h+1}^{T}X_tX'_{t-h}$.

For any h, we can decompose $\hat{\Gamma}_{\hat{G}}(h)$ as:

$$\hat{\Gamma}_{\hat{G}}(h) = \hat{D}^{-1/2}\hat{P}'\Lambda_0\Phi_0(h)\Lambda'_0\hat{P}\hat{D}^{-1/2} + \hat{D}^{-1/2}\hat{P}'\left(S(h) - \Lambda_0\Phi_0(h)\Lambda'_0\right)\hat{P}\hat{D}^{-1/2}$$

First, we can write:

$$\hat{D}^{-1/2}\hat{P}'\Lambda_0\Phi_0(h)\Lambda'_0\hat{P}\hat{D}^{-1/2} = \hat{D}^{-1/2}\hat{P}'P_0D_0^{1/2}\Phi_0(h)D_0^{1/2}P'_0\hat{P}\hat{D}^{-1/2}$$

It then follows from lemma 2 (iv), lemma 4 (i), and the fact that $\Phi_0(h) = O(1)$ that:

$$\hat{D}^{-1/2}\hat{P}'\Lambda_0\Phi_0(h)\Lambda'_0\hat{P}\hat{D}^{-1/2} = \Phi_0(h) + O_P\left(\frac{1}{n}\right) + O_P\left(\frac{1}{\sqrt{T}}\right)$$

Then, under assumption (A3) and (CR2), it is possible to extend what has been done in Lemma 2(i) for $h = 0$, and to show that: $\frac{1}{n}\|S(h) - \Lambda_0\Phi_0(h)\Lambda'_0\| = O_P\left(\frac{1}{n}\right) + O_P\left(\frac{1}{\sqrt{T}}\right)$. Indeed, if we decompose $S(h)$ as:

$$S(h) = \frac{1}{T-h}\left[\Lambda_0\sum_{t=h+1}^{T}G_tG'_{t-h}\Lambda'_0 + \Lambda_0\sum_{t=h+1}^{T}G_t\xi'_{t-h} + \sum_{t=h+1}^{T}\xi_tG'_{t-h}\Lambda'_0 + \sum_{t=h+1}^{T}\xi_t\xi'_{t-h}\right]$$

we get:

$$\frac{1}{n}(S(h) - \Lambda_0\Phi_0(h)\Lambda'_0) = \frac{1}{n}\Lambda_0\left(\frac{1}{T-h}\sum_{t=h+1}^{T}G_tG'_{t-h} - \Phi_0(h)\right)\Lambda'_0$$
$$+ \frac{1}{n}\left(\Lambda_0\frac{1}{T-h}\sum_{t=h+1}^{T}G_t\xi'_{t-h} + \frac{1}{T-h}\sum_{t=h+1}^{T}\xi_tG'_{t-h}\Lambda'_0\right)$$
$$+ \frac{1}{n}\left(\frac{1}{T-h}\sum_{t=h+1}^{T}\xi_t\xi'_{t-h} - \Psi_0(h)\right) + \frac{1}{n}\Psi_0(h)$$

Then, using assumptions (A3) and (CR2) and a multivariate extension of the proof given in the univariate case by Brockwell and Davies (1991, pp226-227), it is possible, as in lemma 2 (i), to show that:

$$\cdot\ \mathrm{E}\left(\|\frac{1}{T-h}\sum_{t=h+1}^{T}G_tG'_{t-h} - \Phi_0(h)\|^2\right) = O\left(\frac{1}{T}\right)$$

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· $\mathrm{E}\left(\|\frac{1}{T-h}\sum_{t=h+1}^{T}\xi_t\xi'_{t-h}-\Psi_0(h)\|^2\right)=O\left(\frac{n^2}{T}\right)$

so that:

· $\|\frac{1}{T-h}\sum_{t=h+1}^{T}G_tG'_{t-h}-\Phi_0(h)\|=O_P\left(\frac{1}{\sqrt{T}}\right)$

· $\|\frac{1}{T-h}\sum_{t=h+1}^{T}\xi_t\xi'_{t-h}-\Psi_0(h)\|=O_P\left(\frac{n}{\sqrt{T}}\right)$

Using the and the same kind of arguments as we have used in lemma 2 (i), it then also follows that:

$$\|\frac{1}{T-h}\sum_{t=h+1}^{T}G_t\xi'_{t-h}\|=O_P\left(\frac{\sqrt{n}}{\sqrt{T}}\right)$$

From assumptions ((CR1), we also have: $\|\Lambda_0\|=O\left(\frac{1}{\sqrt{n}}\right)$ and $\|\Psi_0(h)\|=O(1)$, so that:

$$\frac{1}{n}\|S(h)-\Lambda_0\Phi_0(h)\Lambda'_0\|=O_P\left(\frac{1}{n}\right)+O_P\left(\frac{1}{\sqrt{T}}\right)$$

Finally, as $\hat{D}^{-1/2}\hat{P}'\left(S(h)-\Lambda_0\Phi_0(h)\Lambda'_0\right)\hat{P}\hat{D}^{-1/2}=(\frac{\hat{D}}{n})^{-1/2}\hat{P}'\frac{S(h)-\Lambda_0\Phi_0(h)\Lambda'_0}{n}\hat{P}(\frac{\hat{D}}{n})^{-1/2}$, and $\frac{\hat{D}}{n}=O_P(1)$, it follows that

$$\hat{D}^{-1/2}\hat{P}'\left(S(h)-\Lambda_0\Phi_0(h)\Lambda'_0\right)\hat{P}\hat{D}^{-1/2}=O_P\left(\frac{1}{n}\right)+O_P\left(\frac{1}{\sqrt{T}}\right)$$

iv) Let us first recall that any VAR(p) model can be written in a VAR(1) form. More precisely, if we denote: $G_t^{(p)}=(G'_t,G'_{t-1},\ldots,G'_{t-p+1})'$, we can write:

$$G_t^{(p)}=A_0^{(p)}G_{t-1}^{(p)}+w_t^{(p)}$$

with $A_0^{(p)}=\begin{pmatrix}A_{01} & A_{02} & \ldots & A_{0p}\\ I_r & 0 & \ldots & 0\\ \vdots & \vdots & & \vdots\\ 0 & 0 & \ldots & I_r\end{pmatrix}$ and $w_t^{(p)}=(w'_t,0,\ldots,0)'$.

If we denote $\Phi_0^{(p)}=E_{\Omega_0}\left[G_t^{(p)}G_t^{(p)\prime}\right]$ and $\Phi_1^{(p)}=E_{\Omega_0}\left[G_t^{(p)}G_{t-1}^{(p)\prime}\right]$, so that:

$$\Phi_0^{(p)}=\begin{pmatrix}I_r & \Phi_0(1) & \ldots & \Phi_0(p-1)\\ \Phi'_0(1) & I_r & \ldots & \Phi_0(p-2)\\ \vdots & \vdots & & \vdots\\ \Phi'_0(p-1) & \Phi'_0(p-2) & \ldots & I_r\end{pmatrix} \quad \Phi_1^{(p)}=\begin{pmatrix}\Phi_0(1) & \Phi_0(2) & \ldots & \Phi_0(p)\\ \Phi'_0(1) & I_r & \ldots & \Phi_0(p-1)\\ \vdots & \vdots & & \vdots\\ \Phi'_0(p-2) & \Phi'_0(p-3) & \ldots & \Phi_0(1)\end{pmatrix}$$

we have:

$$A_0^{(p)}=\Phi_1^{(p)}(\Phi_0^{(p)})^{-1}$$

We can define $\hat{\Phi}_0^{(p)}$ and $\hat{\Phi}_1^{(p)}$ having respectively the same form as $\Phi_0^{(p)}$ and $\Phi_1^{(p)}$, with $\Phi_{0,k}$ replaced by $\hat{\Gamma}_{\hat{G}}(k)$ for any value of k. Then, we also have:

$$\hat{A}^{(p)} = \hat{\Phi}_1^{(p)}(\hat{\Phi}_0^{(p)})^{-1}$$

where: $\hat{A}^{(p)} = \begin{pmatrix} \hat{A}_1 & \hat{A}_2 & \dots & \hat{A}_p \\ I_r & 0 & \dots & 0 \\ \vdots & \vdots & & \vdots \\ 0 & 0 & \dots & I_r \end{pmatrix}$.

It thus follows from (iii) that:

$$\|\Phi_0^{(p)} - \hat{\Phi}_0^{(p)}\| = O_P\left(\frac{1}{n}\right) + O_P\left(\frac{1}{\sqrt{T}}\right) \text{ and } \|\Phi_1^{(p)} - \hat{\Phi}_1^{(p)}\| = O_P\left(\frac{1}{n}\right) + O_P\left(\frac{1}{\sqrt{T}}\right)$$

We have: $\|A_0^{(p)} - \hat{A}^{(p)}\| \leq \|\Phi_1^{(p)} - \hat{\Phi}_1^{(p)}\|\|(\Phi_0^{(p)})^{-1}\| + \|\hat{\Phi}_1^{(p)}\|\|(\Phi_0^{(p)})^{-1} - (\hat{\Phi}_0^{(p)})^{-1}\|$. If we apply to the last term the relation (R) which has been introduced in the proof of proposition 1, we then get:

$$\|A_0^{(p)} - \hat{A}^{(p)}\| = O_P\left(\frac{1}{n}\right) + O_P\left(\frac{1}{\sqrt{T}}\right)$$

It then follows that: $\|A_{0s} - \hat{A}_s\| = O_P\left(\frac{1}{n}\right) + O_P\left(\frac{1}{\sqrt{T}}\right)$ for any s.

Proof of Corollary 1

i) As $\hat{\bar{\psi}}_{R1} = \frac{1}{n-r}\text{trace}(S - \hat{D}) = \frac{1}{n-r}\text{trace}\left(S - \hat{\Lambda}\hat{\Lambda}'\right)$, it follows from proposition 2 (ii) that:

$$\hat{\bar{\psi}}_{R1} = \frac{1}{n-r}\sum_{i=1}^{n}\hat{\psi}_{ii} = \frac{1}{n-r}\sum_{i=1}^{n}\left(\psi_{0,ii} + O_P\left(\frac{1}{n}\right) + O_P\left(\frac{1}{\sqrt{T}}\right)\right) = O_P(1)$$

The result then immediately follows from the fact that $\hat{D} = O_P(n)$ and:

$$\hat{G}_{t/T,R1} = \hat{D}^{-1}\left(\hat{D} - \hat{\bar{\psi}}_{R1}I_r\right)^{1/2}\hat{P}'X_t = \hat{D}^{-1}\left(\hat{D} - \hat{\bar{\psi}}_{R1}I_r\right)^{1/2}\hat{D}^{1/2}\hat{G}_t$$

ii) As $\hat{\Lambda}_{R1} = \hat{P}\left(\hat{D} - \hat{\bar{\psi}}_{R1}I_r\right)^{1/2} = \hat{\Lambda}\hat{D}^{-1/2}\left(\hat{D} - \hat{\bar{\psi}}_{R1}I_r\right)^{1/2}$, it then follows from (i) that $\hat{\Lambda}_{R1}$ is asymptotically equivalent to $\hat{\Lambda}$, and that all the properties which have been obtained for $\hat{\Lambda}$ are also true for $\hat{\Lambda}_{R1}$.

A.3   Consistency of weighted PCA ($\hat{\Omega}^{R2}$ framework)

From Proposition 1 (iii) we know that, for any i:

$$\hat{\psi}_{ii} = \psi_{0,ii} + O_P\left(\frac{1}{n}\right) + O_P\left(\frac{1}{\sqrt{T}}\right)$$

so that $\underline{\lambda} + O_P\left(\frac{1}{n}\right) + O_P\left(\frac{1}{\sqrt{T}}\right) < \hat{\psi}_{ii} < \overline{\lambda} + O_P\left(\frac{1}{n}\right) + O_P\left(\frac{1}{\sqrt{T}}\right)$. Equivalently, if $\underline{d}$ and $\overline{d}$ are two given numbers such that: $0 < \underline{d} < \underline{\lambda}$ and $\overline{\lambda} < \overline{d} < \infty$,

$$\underline{d}I_n < \hat{\Psi}_d < \overline{d}I_n$$

for any n and T, where $\hat{\Psi}_d := \mathrm{diag}(\hat{\psi}_{11}, ..., \hat{\psi}_{nn})$.

$$\mathrm{Proj}_{\hat{\Omega}^{R2}}[G_t|X_s, s \leq T] = \left(\hat{\Lambda}'\hat{\Psi}_d^{-1}\hat{\Lambda} + \hat{\Phi}^{-1}\right)^{-1}\hat{\Lambda}'\hat{\Psi}_d^{-1}X_t$$

This estimates of the common factors are proportional to PCA on the weighted data. More precisely, if we denote by:

- $X_t^w = \hat{\Psi}_d^{-1/2}X_t$ the vector of weighted data

- $S^w = \frac{1}{T}\sum_{t=1}^T X_t^w X_t^{w\prime}$ the associated empirical variance-covariance matrix

- $\Lambda_0^w = \hat{\Psi}_d^{-1/2}\Lambda_0$

- $\Psi_0^w = \hat{\Psi}_d^{-1/2}\Psi_0\hat{\Psi}_d^{-1/2}$

- $\Sigma_0^w = \Lambda_0^w\Lambda_0^{w\prime} + \Psi_0^w$

it is straightforward to extend the previous proofs to this new case. Actually, as: $\underline{d}I_n < \hat{\Psi}_d < \overline{d}I_n$, we get, for any symmetric matrix M:

$$\frac{1}{\sqrt{\overline{d}}}M < \hat{\Psi}_d^{-1/2}M\hat{\Psi}_d^{-1/2} < \frac{1}{\sqrt{\underline{d}}}M$$

Thus, the assumptions which have been made for the initial matrices Σ_0, Λ_0 and Ψ_0 are all still valid for the matrices Σ_0^w, Λ_0^w and Ψ_0^w. In the same way, the assumptions which have been made for the initial data X_t are still valid for X_t^w.

A.4   Consistency of Kalman Filtering: ($\hat{\Omega}^{R3}$ and $\hat{\Omega}^{R4}$ framework)

Lemma 5

i) $(\hat{P} - P_0)'\Psi_{0R}^{-1}P_0 = O_P\left(\frac{1}{n}\right) + O_P\left(\frac{1}{\sqrt{T}}\right)$

ii) $\hat{P}'\hat{\Psi}_R^{-1}\hat{P} - P_0'\Psi_{0R}^{-1}P_0 = O_P\left(\frac{1}{n}\right) + O_P\left(\frac{1}{\sqrt{T}}\right)$

iii) $\|\hat{P}'\hat{\Psi}_R^{-1} - P_0'\Psi_{0R}^{-1}\| = O_P\left(\frac{1}{n}\right) + O_P\left(\frac{1}{\sqrt{T}}\right)$

iv) $\|(\hat{P}'\hat{\Psi}_R^{-1}\hat{P})^{-1}\hat{P}'\hat{\Psi}_R^{-1} - (P_0'\Psi_{0R}^{-1}P_0)^{-1}P_0'\Psi_{0R}^{-1}\| = O_P\left(\frac{1}{n}\right) + O_P\left(\frac{1}{\sqrt{T}}\right)$

v) $\frac{1}{n}\|\hat{\Lambda}'\hat{\Psi}_R^{-1}\hat{\Lambda} - \Lambda_0'\Psi_{0R}^{-1}\Lambda_0\| = O_P\left(\frac{1}{n}\right) + O_P\left(\frac{1}{\sqrt{T}}\right)$

vi) $\|(\hat{\Lambda}'\hat{\Psi}_R^{-1}\hat{\Lambda})^{-1}\hat{\Lambda}'\hat{\Psi}_R^{-1} - (\Lambda_0'\Psi_{0R}^{-1}\Lambda_0)^{-1}\Lambda_0'\Psi_{0R}^{-1}\| = O_P\left(\frac{1}{n\sqrt{n}}\right) + O_P\left(\frac{1}{\sqrt{n}\sqrt{T}}\right)$

Proof

i) Defining $P_{0\perp}$ as we did in the proof of proposition 2, we can write:

$$(\hat{P} - P_0)'\Psi_{0R}^{-1}P_0 = (\hat{P} - P_0)'(P_0P_0' + P_{0\perp}P_{0\perp}')\Psi_{0R}^{-1}P_0$$
$$= \hat{P}'P_0P_0'\Psi_{0R}^{-1}P_0 - P_0'\Psi_{0R}^{-1}P_0 + \hat{P}'P_{0\perp}P_{0\perp}'\Psi_{0R}^{-1}P_0$$

We have seen before (see proof of proposition 2) that:

$$\|\hat{P}'P_{0\perp}\| = O_P\left(\frac{1}{n}\right) + O_P\left(\frac{1}{\sqrt{T}}\right)$$

As $P_0'\Psi_{0R}^{-1}P_0$ and $P_{0\perp}'\Psi_{0R}^{-1}P_0$ are $O(1)$, the result then follows from lemma 4 (i).

ii) $\hat{P}'\hat{\Psi}_R^{-1}\hat{P} - P_0'\Psi_{0R}^{-1}P_0 = \hat{P}'(\hat{\Psi}_R^{-1} - \Psi_{0R}^{-1})\hat{P} + \hat{P}'\Psi_{0R}^{-1}\hat{P} - P_0'\Psi_{0R}^{-1}P_0$.

As $\|\hat{P}'(\hat{\Psi}_R^{-1} - \Psi_{0R}^{-1})\hat{P}\| \leq \|\hat{P}\|^2\|\hat{\Psi}_R^{-1} - \Psi_{0R}^{-1}\| = \|\hat{\Psi}_R^{-1} - \Psi_{0R}^{-1}\|$,
and as $\|\hat{\Psi}_R^{-1} - \Psi_{0R}^{-1}\| = \text{Max}_{1\leq i\leq n}|\hat{\psi}_{ii}^{-1} - \psi_{0ii}^{-1}|$, it follows from proposition 3 (ii) that

$$\|\hat{P}'(\hat{\Psi}_R^{-1} - \Psi_{0R}^{-1})\hat{P}\| = O_P\left(\frac{1}{n}\right) + O_P\left(\frac{1}{\sqrt{T}}\right)$$

Further:

$$\|\hat{P}'\Psi_{0R}^{-1}\hat{P} - P_0'\Psi_{0R}^{-1}P_0\| = \|(\hat{P} - P_0)'\Psi_{0R}^{-1}P_0 + P_0'\Psi_{0R}^{-1}(\hat{P} - P_0) + (\hat{P} - P_0)'\Psi_{0R}^{-1}(\hat{P} - P_0)\|$$
$$\leq 2\|(\hat{P} - P_0)'\Psi_{0R}^{-1}P_0\| + \|\Psi_{0R}^{-1}\|\|\hat{P} - P_0\|^2$$

It then follows from lemma 4 (ii), assumption (CR2), and lemma 5 (i) that

$$\|\hat{P}'\Psi_{0R}^{-1}\hat{P} - P_0'\Psi_{0R}^{-1}P_0\| = O_P\left(\frac{1}{n}\right) + O_P\left(\frac{1}{\sqrt{T}}\right)$$

so that (ii) follows.

iii) In the same way: $\|\hat{P}'\hat{\Psi}_R^{-1} - P_0'\Psi_{0R}^{-1}\| \leq \|\hat{P}'(\hat{\Psi}_R^{-1} - \Psi_{0R}^{-1})\| + \|(\hat{P} - P_0)'\Psi_{0R}^{-1}\|$ with:

$$\|\hat{P}'(\hat{\Psi}_R^{-1} - \Psi_{0R}^{-1})\| \leq \|\hat{\Psi}_R^{-1} - \Psi_{0R}^{-1}\| = O_P\left(\frac{1}{n}\right) + O_P\left(\frac{1}{\sqrt{T}}\right)$$

and:

$$\|(\hat{P} - P_0)'\Psi_{0R}^{-1}\| = \|(\hat{P} - P_0)'(P_0P_0' + P_{0\perp}P_{0\perp}')\Psi_{0R}^{-1}\|$$
$$\leq \|(\hat{P}'P_0 - I_r)P_0'\Psi_{0R}^{-1}\| + \|\hat{P}'P_{0\perp}P_{0\perp}'\Psi_{0R}^{-1}\|$$
$$\leq \|\hat{P}'P_0 - I_r\|\|P_0'\Psi_{0R}^{-1}\| + \|\hat{P}'P_{0\perp}\|\|P_{0\perp}'\Psi_{0R}^{-1}\|$$
$$= O_P\left(\frac{1}{n}\right) + O_P\left(\frac{1}{\sqrt{T}}\right)$$

iv) As $\|\Psi_0^{-1}\| = O(1)$ by asssumption (A4), we know from proposition 3 (ii) that $\|\hat{\Psi}_R^{-1}\| = O_P(1)$ so that $(\hat{P}'\hat{\Psi}_R^{-1}\hat{P})^{-1} = O_P(1)$. We then can write:

$$\|(\hat{P}'\hat{\Psi}_R^{-1}\hat{P})^{-1}\hat{P}'\hat{\Psi}_R^{-1} - (P_0'\Psi_{0R}^{-1}P_0)^{-1}P_0'\Psi_R^{-1}\|$$
$$= \|(\hat{P}'\hat{\Psi}_R^{-1}\hat{P})^{-1}(\hat{P}'\hat{\Psi}_R^{-1} - P_0'\Psi_R^{-1})$$
$$+ ((\hat{P}'\hat{\Psi}_R^{-1}\hat{P})^{-1} - P_0'\Psi_R^{-1}P_0)^{-1})P_0'\Psi_R^{-1}\|$$
$$= \|(\hat{P}'\hat{\Psi}_R^{-1}\hat{P})^{-1}(\hat{P}'\hat{\Psi}_R^{-1} - P_0'\Psi_R^{-1})$$
$$+ (\hat{P}'\hat{\Psi}_R^{-1}\hat{P})^{-1}[P_0'\Psi_R^{-1}P_0 - \hat{P}'\hat{\Psi}_R^{-1}\hat{P}](P_0'\Psi_R^{-1}P_0)^{-1}P_0'\Psi_R^{-1}\|$$
$$\leq \|(\hat{P}'\hat{\Psi}_R^{-1}\hat{P})^{-1}\|\|(\hat{P}'\hat{\Psi}_R^{-1} - P_0'\Psi_R^{-1})\|$$
$$+ \|(\hat{P}'\hat{\Psi}_R^{-1}\hat{P})^{-1}\|\|P_0'\Psi_R^{-1}P_0 - \hat{P}'\hat{\Psi}_R^{-1}\hat{P}\|\|(P_0'\Psi_R^{-1}P_0)^{-1}\|\|P_0'\Psi_R^{-1}\|$$

The result then follows from (ii) and (iii).

v) $\frac{1}{n}\hat{\Lambda}'\hat{\Psi}_R^{-1}\hat{\Lambda} = \frac{1}{n}\hat{D}^{1/2}\hat{P}'\hat{\Psi}_R^{-1}\hat{P}\hat{D}^{1/2} = \frac{1}{n}\hat{D}^{1/2}D_0^{-1/2}D_0^{1/2}\hat{P}'\hat{\Psi}_R^{-1}\hat{P}D_0^{1/2}D_0^{-1/2}\hat{D}^{1/2}$.

The result then follows from lemma 2 (iv), lemma 5 (ii), and the fact that $D_0 = O_P\left(\frac{1}{n}\right)$.

vi) We can write:

$$(\hat{\Lambda}'\hat{\Psi}_R^{-1}\hat{\Lambda})^{-1}\hat{\Lambda}'\hat{\Psi}_R^{-1} - (\Lambda_0'\Psi_{0R}^{-1}\Lambda_0)^{-1}\Lambda_0'\Psi_{0R}^{-1}$$
$$= \hat{D}^{-1/2}(\hat{P}'\hat{\Psi}_R^{-1}\hat{P})^{-1}\hat{P}'\hat{\Psi}_R^{-1} - D_0^{-1/2}(P_0'\Psi_{0R}^{-1}P_0)^{-1}P_0'\Psi_{0R}^{-1}$$
$$= \hat{D}^{-1/2}\left((\hat{P}'\hat{\Psi}_R^{-1}\hat{P})^{-1}\hat{P}'\hat{\Psi}_R^{-1} - \hat{D}^{1/2}D_0^{-1/2}(P_0'\Psi_{0R}^{-1}P_0)^{-1}P_0'\Psi_{0R}^{-1}\right)$$

Lemma 6

Denote $\hat{\Sigma}_{\hat{G},R}$ is then the empirical counterpart $\Sigma_{G,R}$, so that $\hat{\Sigma}_{\hat{G},R}$ is the (rT, rT) matrix whose general (s,t) block entry is $\hat{\Gamma}_{\hat{G}}(s-t)$. The following properties hold:

i) $\|\hat{\Sigma}_{\hat{G},R} - \Sigma_{G,R}\| = O_P\left(\frac{1}{n}\right) + O_P\left(\frac{1}{\sqrt{T}}\right)$

ii) $\|\hat{\Sigma}_{\hat{G},R}\| = O_P(1)$ and $\|\hat{\Sigma}_{\hat{G},R}^{-1}\| = O_P(1)$

iii) $\|\hat{\Sigma}_{\hat{G},R}^{-1} - \Sigma_{G,R}^{-1}\| = O_P\left(\frac{1}{n}\right) + O_P\left(\frac{1}{\sqrt{T}}\right)$

Proof

(i) If $\mathbf{x} = (x_1', ...x_T')'$ is a non-random vector of $\mathbb{R}^{rT}$ such that: $\|x\|^2 = \sum_{t=1}^T \|x_t\|^2 = 1$, we can write:

$$\mathbf{x}'(\hat{\Sigma}_{\hat{G},R} - \Sigma_{G,R})\mathbf{x} = \sum_{t=1}^T \sum_{\tau=1}^T x_t'\left(\hat{\Gamma}_{\hat{G}}(t-\tau) - \Gamma_G(t-\tau)\right)x_\tau$$

so that:

$$\mathbf{x}'(\hat{\Sigma}_{\hat{G},R} - \Sigma_{G,R})\mathbf{x} \leq \sum_{t=1}^T \sum_{\tau=1}^T \|x_t\|\|\hat{\Gamma}_{\hat{G}}(t-\tau) - \Gamma_G(t-\tau)\|\|x_\tau\|$$
$$\leq \sum_{t=1}^T \sum_{\tau=1}^T \|x_t\|\|\hat{\Gamma}_{\hat{G}}(t-\tau) - \Gamma_G(t-\tau)\|\|x_\tau\|$$
$$\leq Max_{|h|\leq(T-1)}\|\hat{\Gamma}_{\hat{G}}(h) - \Gamma_G(h)\| \sum_{t=1}^T \sum_{\tau=1}^T \|x_t\|\|x_\tau\|$$
$$\leq Max_{|h|\leq(T-1)}\|\hat{\Gamma}_{\hat{G}}(h) - \Gamma_G(h)\| \sum_{t=1}^T \|x_t\|^2$$
$$= Max_{|h|\leq(T-1)}\|\hat{\Gamma}_{\hat{G}}(h) - \Gamma_G(h)\|$$

The result then follows from proposition 3 (iii).

(ii) Follows directly from (i) and lemma 1 (iii).

(iii) Applying relation (R) as in the proof of proposition 1, we get:

$$\|\hat{\Sigma}_{\hat{G},R}^{-1} - \Sigma_{G,R}^{-1}\| \leq \|\hat{\Sigma}_{\hat{G},R}^{-1}\|\|\hat{\Sigma}_{\hat{G},R} - \Sigma_{G,R}\|\|\Sigma_{G,R}^{-1}\|$$

The result the follows from (i) and (ii).

Proof of proposition 4

As $G_{t/t,R} = \text{Proj}_{\Omega^R}[G_t|X_t]$ and $\hat{G}_{t/t,R} = \text{Proj}_{\hat{\Omega}^R}[G_t|X_t]$, they are obtained through the same formulas so that, by construction:

$$\hat{G}_{t/T,R} = \mathbf{U}_t'(\hat{\Sigma}_{\hat{G},R}^{-1} + I_T \otimes \hat{\Lambda}'\hat{\Psi}_R^{-1}\hat{\Lambda})^{-1}(I_T \otimes \hat{\Lambda}'\hat{\Psi}_R^{-1})\mathbf{X}_T$$

Using relation (R) as in the proof of proposition 1 (Taylor expansion at order 0), we obtain the same kind of decomposition for $\hat{G}_{t/T,R}$ as the one we have used to study $G_{t/T,R}$. Thus, if we denote $\hat{M} = \hat{\Lambda}'\hat{\Psi}_R^{-1}\hat{\Lambda}$, we can write: $\hat{G}_{t/T,R} = \hat{G}_{t/T,R}^1 - \hat{G}_{t/T,R}^2 - \hat{G}_{t/T,R}^3$, with:

$$\hat{G}_{t/T,R}^1 = \mathbf{U}_t' \left(I_T \otimes \hat{M}^{-1} \right) \left(I_T \otimes \hat{\Lambda}'\hat{\Psi}_R^{-1} \right) \mathbf{X}_T = \hat{M}^{-1}\hat{\Lambda}'\hat{\Psi}_R^{-1}X_t$$

$$\hat{G}_{t/T,R}^2 = \mathbf{U}_t' \left(\hat{\Sigma}_{\hat{G},R}^{-1} + I_T \otimes \hat{M} \right)^{-1} \hat{\Sigma}_{\hat{G},R}^{-1} \left(I_T \otimes \hat{M}^{-1}\hat{\Lambda}'\hat{\Psi}_R^{-1} \right) (I_T \otimes \Lambda_0) \mathbf{G}_T$$

$$\hat{G}_{t/T,R}^3 = \mathbf{U}_t' \left(\hat{\Sigma}_{\hat{G},R}^{-1} + I_T \otimes \hat{M} \right)^{-1} \hat{\Sigma}_{\hat{G},R}^{-1} \left(I_T \otimes \hat{M}^{-1}\hat{\Lambda}'\hat{\Psi}_R^{-1} \right) \mathbf{Z}_T$$

Let us study separately these three terms.

If we compare the first term with $G_{t/T,R}^1$, we get:

$$\hat{G}_{t/T,R}^1 - G_{t/T,R}^1 = \left(\left(\hat{\Lambda}'\hat{\Psi}_R^{-1}\hat{\Lambda} \right)^{-1} \hat{\Lambda}'\hat{\Psi}_R^{-1} - \left(\Lambda_0'\Psi_{0R}^{-1}\Lambda_0 \right)^{-1} \Lambda_0'\Psi_{0R}^{-1} \right) X_t$$

with

$$\| \left(\left(\hat{\Lambda}'\hat{\Psi}_R^{-1}\hat{\Lambda} \right)^{-1} \hat{\Lambda}'\hat{\Psi}_R^{-1} - \left(\Lambda_0'\Psi_{0R}^{-1}\Lambda_0 \right)^{-1} \Lambda_0'\Psi_{0R}^{-1} \right) X_t\|$$
$$\leq \| \left(\hat{\Lambda}'\hat{\Psi}_R^{-1}\hat{\Lambda} \right)^{-1} \hat{\Lambda}'\hat{\Psi}_R^{-1} - \left(\Lambda_0'\Psi_{0R}^{-1}\Lambda_0 \right)^{-1} \Lambda_0'\Psi_{0R}^{-1}\|\|X_t\|$$

As $X_t = O_P(\sqrt{n})$, it then follows from lemma 5 (v) that:

$$\hat{G}_{t/T,R}^1 - G_{t/T,R}^1 = O_P(\frac{1}{n}) + O_P(\frac{1}{\sqrt{T}})$$

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Finally, as $G_{t/T,R}^1 = G_t + O_P(\frac{1}{\sqrt{n}})$, we get:

$$\hat{G}_{t/T,R}^1 = G_t + O_P(\frac{1}{\sqrt{n}}) + O_P(\frac{1}{\sqrt{T}})$$

In the same way, we can write:

$$\hat{G}_{t/T,R}^2 - G_{t/T,R}^2 = \mathbf{U}_t' \left(\hat{\Sigma}_{\hat{G},R}^{-1} + I_T \otimes \hat{M} \right)^{-1} \hat{\Sigma}_{\hat{G},R}^{-1} \left(I_T \otimes \hat{M}^{-1} \hat{\Lambda}' \hat{\Psi}_R^{-1} \right) (I_T \otimes \Lambda_0) \mathbf{G}_T$$
$$-\mathbf{U}_t' \left(\Sigma_{G,R}^{-1} + I_T \otimes M_0 \right)^{-1} \Sigma_{G,R}^{-1} \left(I_T \otimes M_0^{-1} \Lambda_0' \Psi_{0R}^{-1} \right) (I_T \otimes \Lambda_0) \mathbf{G}_T$$

and:

$$\hat{G}_{t/T,R}^3 - G_{t/T,R}^3 = \mathbf{U}_t' \left(\hat{\Sigma}_{\hat{G},R}^{-1} + I_T \otimes \hat{M} \right)^{-1} \hat{\Sigma}_{\hat{G},R}^{-1} \left(I_T \otimes \hat{M}^{-1} \hat{\Lambda}' \hat{\Psi}_R^{-1} \right) \mathbf{Z}_T$$
$$-\mathbf{U}_t' \left(\Sigma_{G,R}^{-1} + I_T \otimes M_0 \right)^{-1} \Sigma_{G,R}^{-1} \left(I_T \otimes M_0^{-1} \Lambda_0' \Psi_{0R}^{-1} \right) \mathbf{Z}_T$$

so that: $\hat{G}_{t/T,R}^2 - G_{t/T,R}^2 = \mathbf{U}_t' \hat{H} (I_T \otimes \Lambda_0) \mathbf{G}_T$ and $\hat{G}_{t/T,R}^3 - G_{t/T,R}^3 = \mathbf{U}_t' \hat{H} \mathbf{Z}_T$ with:

$$\hat{H} = (\hat{\Sigma}_{\hat{G},R}^{-1} + I_T \otimes \hat{M})^{-1} \hat{\Sigma}_{\hat{G},R}^{-1} (I_T \otimes \hat{M}^{-1} \hat{\Lambda}' \hat{\Psi}_R^{-1}) - (\Sigma_{G,R}^{-1} + I_T \otimes M_0)^{-1} \Sigma_{G,R}^{-1} (I_T \otimes M_0^{-1} \Lambda_0' \Psi_{0R}^{-1})$$

We can also decompose $\hat{H}$ as: $\hat{H} = \hat{H}_1 + \hat{H}_2 + \hat{H}_3$ with:

$$\hat{H}_1 = (\hat{\Sigma}_{\hat{G},R}^{-1} + I_T \otimes \hat{M})^{-1} \hat{\Sigma}_{\hat{G},R}^{-1} \left(I_T \otimes (\hat{M}^{-1} \hat{\Lambda}' \hat{\Psi}_R^{-1} - M_0^{-1} \Lambda_0' \Psi_{0R}^{-1}) \right)$$
$$\hat{H}_2 = (\hat{\Sigma}_{\hat{G},R}^{-1} + I_T \otimes \hat{M})^{-1} (\hat{\Sigma}_{\hat{G},R}^{-1} - \Sigma_{G,R}^{-1})(I_T \otimes M_0^{-1} \Lambda_0' \Psi_{0R}^{-1})$$
$$\hat{H}_3 = \left((\hat{\Sigma}_{\hat{G},R}^{-1} + I_T \otimes \hat{M})^{-1} - (\Sigma_{G,R}^{-1} + I_T \otimes M_0)^{-1} \right) \Sigma_{G,R}^{-1} (I_T \otimes M_0^{-1} \Lambda_0' \Psi_{0R}^{-1})$$

We then get:

$$\|\hat{H}_1\| \leq \|(\hat{\Sigma}_{\hat{G},R}^{-1} + I_T \otimes \hat{M})^{-1}\| \|\hat{\Sigma}_{\hat{G},R}^{-1}\| \|I_T \otimes (\hat{M}^{-1} \hat{\Lambda}' \hat{\Psi}_R^{-1} - M_0^{-1} \Lambda_0' \Psi_{0R}^{-1})\|$$
$$\leq \|I_T \otimes \hat{M}^{-1}\| \|\hat{\Sigma}_{\hat{G},R}^{-1}\| \|I_T \otimes (\hat{M}^{-1} \hat{\Lambda}' \hat{\Psi}_R^{-1} - M_0^{-1} \Lambda_0' \Psi_{0R}^{-1})\|$$
$$= \|\hat{M}^{-1}\| \|\hat{\Sigma}_{\hat{G},R}^{-1}\| \|\hat{M}^{-1} \hat{\Lambda}' \hat{\Psi}_R^{-1} - M_0^{-1} \Lambda_0' \Psi_{0R}^{-1}\|$$

$$\|\hat{H}_2\| \leq \|(\hat{\Sigma}_{\hat{G},R}^{-1} + I_T \otimes \hat{M})^{-1}\| \|\hat{\Sigma}_{\hat{G},R}^{-1} - \Sigma_{G,R}^{-1}\| \|I_T \otimes M_0^{-1} \Lambda_0' \Psi_{0R}^{-1}\|$$
$$\leq \|\hat{M}^{-1}\| \|\hat{\Sigma}_{\hat{G},R}^{-1} - \Sigma_{G,R}^{-1}\| \|M_0^{-1} \Lambda_0' \Psi_{0R}^{-1}\|$$

$$\|\hat{H}_3\| \leq \|(\hat{\Sigma}_{\hat{G},R}^{-1} + I_T \otimes \hat{M})^{-1} - (\Sigma_{G,R}^{-1} + I_T \otimes M_0)^{-1}\| \|\Sigma_{G,R}^{-1}\| \|I_T \otimes M_0^{-1} \Lambda_0' \Psi_{0R}^{-1}\|$$
$$\leq \|(\hat{\Sigma}_{\hat{G},R}^{-1} + I_T \otimes \hat{M})^{-1}\| \|\hat{\Sigma}_{\hat{G},R}^{-1} + I_T \otimes \hat{M} - \Sigma_{G,R}^{-1} - I_T \otimes M_0\|$$
$$\times \|(\Sigma_{G,R}^{-1} + I_T \otimes M_0)^{-1}\| \|\Sigma_{G,R}^{-1}\| \|M_0^{-1} \Lambda_0' \Psi_{0R}^{-1}\|$$
$$\leq \|\hat{M}^{-1}\| \left[\|\hat{\Sigma}_{\hat{G},R}^{-1} - \Sigma_{G,R}^{-1}\| + \|\hat{M} - M_0\| \right] \|M_0^{-1}\| \|\Sigma_{G,R}^{-1}\| \|M_0^{-1} \Lambda_0' \Psi_{0R}^{-1}\|$$

From lemma 5 (v), we get: $\hat{M}^{-1} = O_P\left(\frac{1}{n}\right)$. Thus, applying lemma 5 (v) and (vi), and lemma 6, we get that:

$$\|\hat{H}_i\| = O_P\left(\frac{1}{n^2\sqrt{n}}\right) + O_P\left(\frac{1}{n\sqrt{nT}}\right) \text{ for } i = 1 \text{ to } 3$$

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so that $\|\hat{H}\| = O_P \left(\frac{1}{n^2 \sqrt{n}} \right) + O_P \left(\frac{1}{n\sqrt{nT}} \right)$.

As $E \left(\|\mathbf{G}_T\| \right)^2 = E \left(\sum_{t=1}^{T} \|G_t\|^2 \right) = rT$, we have: $\|\mathbf{G}_T\| = O_P \left(\sqrt{T} \right)$ so that:

$$\|\hat{G}_{t/T,R}^2 - G_{t/T,R}^2\| \le \|\mathbf{U}_t\| \|\hat{H}\| \|I_T \otimes \Lambda_0\| \|\mathbf{G}_T\| = O_P \left(\frac{\sqrt{T}}{n^2} \right) + O_P \left(\frac{1}{n} \right)$$

Similarly, $E \left(\|\mathbf{Z}_T\| \right)^2 = E \left(\sum_{t=1}^{T} \|\xi_t\|^2 \right) = T tr(\Psi_0) = O(nT)$, so that:

$$\|\hat{G}_{t/T,R}^3 - G_{t/T,R}^3\| \le \|\mathbf{U}_t\| \|\hat{H}\| \|\mathbf{Z}_T\| = O_P \left(\frac{\sqrt{T}}{n^2} \right) + O_P \left(\frac{1}{n} \right)$$

Finally, as we know, from the proof of proposition 1 that:

$$G_{t/T,R}^2 = O_P \left(\frac{1}{n} \right) \text{ and } G_{t/T,R}^3 = O_P \left(\frac{1}{n\sqrt{n}} \right)$$

we get: $\hat{G}_{t/T,R}^2 + \hat{G}_{t/T,R}^3 = O_P \left(\frac{\sqrt{T}}{n^2} \right) + O_P \left(\frac{1}{n} \right)$, so that:

$$\hat{G}_{t/T,R} = G_t + O_P \left(\frac{1}{\sqrt{n}} \right) + O_P \left(\frac{1}{\sqrt{T}} \right) + O_P \left(\frac{\sqrt{T}}{n^2} \right)$$

If $\limsup \frac{T}{\sqrt{n^3}} = O(1)$, we then get:

$$\hat{G}_{t/T,R} = G_t + O_P \left(\frac{1}{\sqrt{n}} \right) + O_P \left(\frac{1}{\sqrt{T}} \right)$$