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# Testing for Instability in Covariance Structures\*

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## Abstract

We propose a test for the stability over time of the covariance matrix of multivariate time series. The analysis is extended to the eigensystem to ascertain changes due to instability in the eigenvalues and/or eigenvectors. Using strong Invariance Principles and Law of Large Numbers, we normalise the CUSUM-type statistics to calculate their supremum over the whole sample. The power properties of the test versus alternative hypotheses, including also the case of breaks close to the beginning/end of sample are investigated theoretically and via simulation. We extend our theory to test for the stability of the covariance matrix of a multivariate regression model. The testing procedures are illustrated by studying the stability of the principal components of the term structure of 18 US interest rates.

**Keywords:** Covariance Matrix, Eigensystem, Change point, CUSUM Statistic.

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# 1 Introduction

In this paper, we propose a testing procedure to evaluate the structural stability of the covariance matrix (and its eigensystem) of multivariate time series. A large amount of empirical evidence shows that the issue of changepoint detection in a covariance matrix is of great importance. A classical example is the application of Principal Component Analysis (PCA) to the term structure of interest rates, with the three main principal components interpreted as “slope”, “level” and “curvature” (Litterman and Scheinkman, 1991). Bliss (1997), Bliss and Smith (1997) and Perignon and Villa (2006) show that the principal components of the term structure change substantially over time. Similar findings, using a different methodology, are in Audrino et al. (2005). PCA is also widely used in macroeconometrics, for instance to forecast inflation (Stock and Watson, 1999, 2002, 2005). The importance of verifying the stability of a covariance matrix is also evident in the context of Vector AutoRegressive (VAR) models. In the context of forecasting, Castle et al. (2010) show that changes in the smallest eigenvalue of the covariance matrix of the error term have a large impact on predictive ability. Furthermore, the Choleski decomposition of the error covariance matrix is routinely employed in the context of variance decomposition analysis, when examining how much of the variance of the forecast error of each variable in a VAR is due to exogenous shocks to the other variables (see e.g. Pesaran and Shin, 1998).

Despite the relevance of the topic, most studies either assume stability as a working assumption without testing for it, or the testing is carried out by splitting the sample, thus assuming knowledge of the break date *a priori*. This calls for a rigorous testing procedure to estimate the location of the changepoint when breaks are detected.

The theoretical framework developed in this paper builds on a plethora of results for the changepoint problem available in statistics and in econometrics. Existing testing procedures (see e.g. the reviews by Csörgő and Horváth, 1997; Aue and Horváth, 2012; and Jandhyala et al., 2013) are typically based on taking the supremum (or some other metric - see Andrews and Ploberger, 1994) of a sequence of CUSUM-type statistics, thus not requiring prior knowledge of the breakdate. In particular, Aue et al. (2009) develop a test for the structural stability of a covariance matrix, based on minimal assumptions. However, a feature of this test is that, by construction, it has power versus breaks occurring at least (respectively, at most)  $O(\sqrt{T})$  time periods from the beginning (respectively to the end) of the sample. Lack of

power versus alternatives close to either end of the sample is a typical feature in this literature (see also Andrews, 1993), which somewhat limits the applicability of the test. Situations where breaks are due to recent events, like e.g. the 2008 recession, are left out of the analysis. Our contribution complements that of Aue et al. (2009) by proposing a test that has power versus breaks occurring close to the beginning/end of the sample.

The main contribution of this paper is twofold. First, testing for changepoints is extended to PCA. In addition, the extension to testing for the stability of principal components is useful for the purpose of dimension reduction. Our simulations show that tests for the stability of the whole covariance matrix have severe size distortions in finite samples. Contrary to this, testing for the stability of eigenvalues is found to have the correct size and good power even for relatively small samples. As a second contribution, our testing procedure is able to detect breaks occurring up to  $O(\ln \ln T)$  periods to the end of the sample. This is achieved by using a Strong Invariance Principle (SIP) and a Strong Law of Large Numbers (SLLN) for the partial sample estimators of the covariance matrix, and by using these results to normalize the CUSUM-type test statistic, using a Darling-Erdős limit theory (see Csörgő and Horváth, 1997; Horváth, 1993). In the Supplemental Material to this paper (henceforth referred to as Kao et al., 2015), we also extend our results to the case of testing for the stability of the covariance matrix of the error term in a multivariate regression setting.

The theory derived in our paper is illustrated through an application to the US term structure of interest rates, with a dataset spanning from the late nineties to the current date. We find (as expected) evidence of changes in the volatility and in the loading of the principal components of the term structure around the end of 2007/beginning of 2008. In the Supplemental Material (Kao et al., 2015) we also report another exercise, based on verifying the stability of the covariance matrix of the error term in a VAR model for exchange rates.

The paper is organized as follows. Section 2 contains the SIP and its extension to the eigensystem. The test statistic and its distribution under the null (as well as its behaviour under local-to-null alternatives) is in Section 3. Monte Carlo evidence is in Section 4, while the application to the term structure of interest rates is in Section 5. Section 6 concludes.

A word on notation. Limits are denoted as “ $\rightarrow$ ” (the ordinary limit); “ $\xrightarrow{p}$ ” (convergence in probability); “ $\xrightarrow{d}$ ” and (convergence in distribution). Orders of magnitude for an almost surely convergent sequence (say  $s_T$ ) are denoted as  $O_{a.s.}(T^\varsigma)$  and  $o_{a.s.}(T^\varsigma)$  when, for some  $\varepsilon > 0$  and

$\tilde{T} < \infty$ ,  $P \left[ |T^{-\varsigma} s_T| < \varepsilon \text{ for all } T \geq \tilde{T} \right] = 1$  and  $T^{-\varsigma} s_T \rightarrow 0$  almost surely respectively. Orders of magnitude for a sequence converging in probability (say  $s'_T$ ) are denoted as  $O_p(T^\varsigma)$  and  $o_p(T^\varsigma)$  when, for some  $\varepsilon > 0$ ,  $\Delta_\varepsilon > 0$  and  $\tilde{T}_\varepsilon < \infty$ ,  $P [|T^{-\varsigma} s'_T| > \Delta_\varepsilon] < \varepsilon$  for all  $T > \tilde{T}_\varepsilon$  and  $T^{-\varsigma} s'_T \rightarrow 0$  in probability respectively. Standard Wiener processes and Brownian bridges of dimension  $q$  are denoted as  $W_q(\cdot)$  and  $B_q(\cdot)$  respectively;  $\|v\|$  denotes the Euclidean norm of a vector  $v$  in  $\mathbb{R}^n$ ; similarly,  $\|A\|$  denotes the Euclidean norm of a matrix  $A$  in  $\mathbb{R}^{n \times n}$ , and  $|\cdot|_p$  the  $L_p$ -norm; the integer part of a real number  $x$  is denoted as  $\lfloor x \rfloor$ . Constants that do not depend on the sample size are denoted as  $M, M', M'', \text{ etc.}$

## 2 Theoretical framework

This section derives results on the convergence rate of the sample covariance matrix, its eigensystem, and an estimator of its asymptotic variance, assuming a covariance stationary time series with no breaks. These calculations are useful in Section 3, for deriving the null distribution of our test.

Let  $\{y_t\}_{t=1}^T$  be a time series of dimension  $n$ ; we assume that  $y_t$  has zero mean and covariance matrix  $\Sigma \equiv E(y_t y'_t)$ . This section contains the asymptotics of the partial sample estimates of  $\Sigma$ ; the results are used in Section 3 in order to construct the CUSUM-type test statistic to test for breaks in  $\Sigma$  and its eigensystem. Specifically, we report a SIP for the partial sample estimators of  $\Sigma$  and an estimator of the long run covariance matrix of the estimated  $\Sigma$ , say  $V_\Sigma$ ; and we extend the asymptotics to PCA.

### *Strong Invariance Principle and estimation of $V_\Sigma$*

Let  $\hat{\Sigma}$  be the sample covariance matrix, i.e.  $\hat{\Sigma} = T^{-1} \sum_{t=1}^T y_t y'_t$ . For a given  $\tau \in [0, 1]$ , we define a point in time  $\lfloor T\tau \rfloor$ , and we use the subscripts  $\tau$  and  $1 - \tau$  to denote quantities calculated using the subsamples  $t = 1, \dots, \lfloor T\tau \rfloor$  and  $t = \lfloor T\tau \rfloor + 1, \dots, T$  respectively. In particular, we consider the sequence of partial sample estimators  $\hat{\Sigma}_\tau = (T\tau)^{-1} \sum_{t=1}^{\lfloor T\tau \rfloor} y_t y'_t$ , and similarly  $\hat{\Sigma}_{1-\tau} = [T(1 - \tau)]^{-1} \sum_{t=\lfloor T\tau \rfloor + 1}^T y_t y'_t$ . Finally, henceforth we denote  $w_t = \text{vec}(y_t y'_t)$  and  $\bar{w}_t = \text{vec}(y_t y'_t - \Sigma)$ .

In the sequel, we need the following assumption.

**Assumption 1** (i)  $\sup_t E \|y_t\|^{2r} < \infty$  for some  $r > 2$ ; (ii)  $y_t$  is  $L_{2+\epsilon}$ -NED (Near Epoch Dependent) for some  $\epsilon > 0$ , of size  $\alpha \in (1, +\infty)$  on an *i.i.d.* basis  $\{v_t\}_{t=-\infty}^{+\infty}$ , with  $r > \frac{2\alpha-1}{\alpha-1} (1 + \frac{\epsilon}{2})$ ;

(iii) letting  $V_{\Sigma,T} = T^{-1}E \left[ \left( \sum_{t=1}^T \bar{w}_t \right) \left( \sum_{t=1}^T \bar{w}_t \right)' \right]$ ,  $V_{\Sigma,T}$  is positive definite uniformly in  $T$ , and as  $T \rightarrow \infty$ ,  $V_{\Sigma,T} \rightarrow V_{\Sigma}$  with  $\|V_{\Sigma}\| < \infty$ ; (iv) letting  $\bar{w}_{it}$  be the  $i$ -th element of  $\bar{w}_t$  and defining  $S_{iT,m} \equiv \sum_{t=m+1}^{m+T} \bar{w}_{it}$ , there exists a positive definite matrix  $\bar{\Omega} = \{\varpi_{ij}\}$  such that  $T^{-1} |E[S_{iT,m} S_{jT,m}] - \varpi_{ij}| \leq MT^{-\psi}$ , for all  $i$  and  $j$  and uniformly in  $m$ , with  $\psi > 0$ .

Assumption 1 specifies the moment conditions and the memory allowed in  $y_t$ ; no distributional assumptions are required. According to part (i), at least the 4-th moment of  $y_t$  is required to be finite, similarly to Aue et al. (2009). As far as serial dependence is concerned, the requirement that  $y_t$  be NED is typical in nonlinear time series analysis (see Gallant and White, 1988) and it implies that  $y_t$  is a mixingale (Davidson, 1994). Many of the DGPs considered in the literature generate NED series - examples include GARCH, bilinear and threshold models (see Davidson, 2002). Part (ii) illustrates the trade-off between the memory of  $y_t$  (i.e. its NED size  $\alpha$ ), and its largest existing moment: as  $\alpha$  (the memory of  $y_t$ ) approaches 1,  $r$  has to increase. Note that in our context, the data ( $y_t$ ) undergo a non-Lipschitz transformation (viz., they are squared), and therefore the relationship between moment conditions and memory is not the “standard” one (see e.g. the IP in Theorem 29.6 in Davidson, 1994). In principle, moment conditions such as the one in part (ii) could be tested for, e.g. using a test based on some tail-index estimator - J. B. Hill (2010, 2011) extends the well-known Hill’s estimator to the context of dependent data. Other types of dependence could be considered, e.g. assuming a linear process for  $y_t$  - an IP for the sample variance is in Phillips and Solo (1992, Theorem 3.8). Part (iv) is a bound on the growth rate of the variance of partial sums of  $\bar{w}_t$ , and it is the same as Assumption A.3 in Corradi (1999). Although it is not needed to prove the IP for the partial sum process of  $\bar{w}_t$ , it is a sufficient condition for the SIP; despite it being rather technical, it can be shown to hold e.g. for the case of a weakly stationary sequence (see Proposition 2.1 in Eberlein (1986)).

Theorem 1 contains the IP and the SIP for the partial sums of  $\bar{w}_t$ .

**Theorem 1** *Under Assumptions 1(i)-(iii), as  $T \rightarrow \infty$*

$$\frac{1}{\sqrt{T}} \sum_{t=1}^{\lfloor T\tau \rfloor} \bar{w}_t \xrightarrow{d} [V_{\Sigma}]^{1/2} W_{n^2}(\tau), \quad (1)$$

*uniformly in  $\tau$ . Redefining  $\bar{w}_t$  in a richer probability space, under Assumptions 1(i)-(iv), there*

exists a  $\delta > 0$  such that

$$\sum_{t=1}^{\lfloor T\tau \rfloor} \bar{w}_t = \sum_{t=1}^{\lfloor T\tau \rfloor} X_t + O_{a.s.} \left( \lfloor T\tau \rfloor^{\frac{1}{2}-\delta} \right), \quad (2)$$

uniformly in  $\tau$ , where  $X_t$  is a zero mean, *i.i.d.* Gaussian sequence with  $E(X_t X_t') = V_\Sigma$ .

### Remarks

T1.1 Equation (1) is an IP for  $\bar{w}_t$  (i.e. a weak convergence result), which is sufficient to use the test statistics discussed e.g. in Andrews (1993) and Andrews and Ploberger (1994).

T1.2 Equation (2) is an almost sure result, which also provides a rate of convergence. The practical consequence of (2) is that the dependent, heteroskedastic series  $\bar{w}_t$  can be replaced with a sequence of *i.i.d.* normally distributed random variables, with the same long run variance as  $\bar{w}_t$ . In both results - (1) and (2) - one difference with the literature is that we are dealing with a non-Lipschitz transformation of NED data (essentially,  $\bar{w}_t$  is the square of  $y_t$ ), which requires some intermediate results on the dependence in  $\bar{w}_t$  itself; we refer to the Supplemental Material (Kao et al., 2015) for the whole set of derivations.

We now turn to the estimation of  $V_\Sigma$ . If no serial dependence is present, a possible choice is the full sample estimator  $\hat{V}_\Sigma = \frac{1}{T} \sum_{t=1}^T w_t w_t' - \left[ \text{vec}(\hat{\Sigma}) \right] \left[ \text{vec}(\hat{\Sigma}) \right]'$ . Alternatively, one could use the sequence of partial sample estimators

$$\hat{V}_{\Sigma,\tau} = \frac{1}{T} \sum_{t=1}^T w_t w_t' - \left\{ \tau \left[ \text{vec}(\hat{\Sigma}_\tau) \right] \left[ \text{vec}(\hat{\Sigma}_\tau) \right]' + (1-\tau) \left[ \text{vec}(\hat{\Sigma}_{1-\tau}) \right] \left[ \text{vec}(\hat{\Sigma}_{1-\tau}) \right]' \right\}.$$

To accommodate for the case  $\Psi_l \equiv E(\bar{w}_t \bar{w}_{t-l}') \neq 0$  for some  $l$ , we propose a weighted sum-of-covariance estimator with bandwidth  $m$ :

$$\tilde{V}_\Sigma = \hat{\Psi}_0 + \sum_{l=1}^m \left( 1 - \frac{l}{m} \right) \left[ \hat{\Psi}_l + \hat{\Psi}_l' \right], \quad (3)$$

where  $\hat{\Psi}_l = \frac{1}{T} \sum_{t=l+1}^T \left[ w_t - \text{vec}(\hat{\Sigma}) \right] \left[ w_{t-l} - \text{vec}(\hat{\Sigma}) \right]'$ ; or  $\tilde{V}_{\Sigma,\tau} = \left( \hat{\Psi}_{0,\tau} + \hat{\Psi}_{0,1-\tau} \right) + \sum_{l=1}^m \left( 1 - \frac{l}{m} \right) \left[ \left( \hat{\Psi}_{l,\tau} + \hat{\Psi}_{l,\tau}' \right) + \left( \hat{\Psi}_{l,1-\tau} + \hat{\Psi}_{l,1-\tau}' \right) \right]$ , where  $\hat{\Psi}_{l,\tau} = \frac{1}{T} \sum_{t=l+1}^{\lfloor T\tau \rfloor} \left[ w_t - \text{vec}(\hat{\Sigma}_\tau) \right] \left[ w_{t-l} - \text{vec}(\hat{\Sigma}_\tau) \right]'$ , and similarly for  $\hat{\Psi}_{l,1-\tau}$ .

In order to derive the asymptotics of  $\hat{V}_{\Sigma,\tau}$  and  $\tilde{V}_{\Sigma,\tau}$ , consider the following assumption:

**Assumption 2.** (i) either (a)  $\Psi_l = 0$  for all  $l \neq 0$  or (b)  $\sum_{l=0}^{\infty} l^s \|\Psi_l\| < \infty$  for  $s = 1$ ; (ii)  $\sup_t E \|y_t\|^{4r} < \infty$  for some  $r > 2$ ; (iii) letting  $\Omega_T = T^{-1} E \left\{ \sum_{t=1}^T \text{vec} [\bar{w}_t \bar{w}_t' - E(\bar{w}_t \bar{w}_t')] \text{vec} [\bar{w}_t \bar{w}_t' - E(\bar{w}_t \bar{w}_t')]' \right\}$ ,  $\Omega_T$  is positive definite uniformly in  $T$ , and  $\Omega_T \rightarrow \Omega$  with  $\|\Omega\| < \infty$ .

Assumption 2 encompasses various possible cases. Part (i)(a) considers the basic, non autocorrelated case, for which both  $\hat{V}_\Sigma$  and  $\hat{V}_{\Sigma,\tau}$  are valid choices. Part (i)(b) considers the possibility of non-zero autocorrelations. Intuitively, the assumption that the 4-th moment of  $y_t$  exists, as in Assumption 1(i), entails, through a Law of Large Numbers (LLN), the consistency of  $\hat{V}_{\Sigma,\tau}$ . Part (ii) supersedes Assumption 1(i), by requiring the existence of moments up to the 8-th. Intuitively, this implies that an IP holds for the partial sums of  $\text{vec} [\bar{w}_t \bar{w}_t' - E(\bar{w}_t \bar{w}_t')]$ .

The consistency of  $\hat{V}_{\Sigma,\tau}$  and of  $\tilde{V}_{\Sigma,\tau}$  is in Theorem 2:

**Theorem 2** *Under no changes in  $\Sigma$ :*

*if Assumptions 1(i)-(iii) and 2(i)(a) hold, as  $T \rightarrow \infty$ , there exists a  $\delta' > 0$  such that*

$$\sup_{1 \leq \lfloor T\tau \rfloor \leq T} \left\| \hat{V}_{\Sigma,\tau} - V_\Sigma \right\| = o_p \left( \frac{1}{T^{\delta'}} \right); \quad (4)$$

*if Assumptions 1(i)-(iii) and 2(i)(b) hold, as  $(m, T) \rightarrow \infty$ , there exists a  $\delta' > 0$  such that*

$$\sup_{1 \leq \lfloor T\tau \rfloor \leq T} \left\| \tilde{V}_{\Sigma,\tau} - V_\Sigma \right\| = O_p \left( \frac{1}{m} \right) + O_p \left( m \frac{\ln T}{T^{\delta'}} \right); \quad (5)$$

*if Assumptions 1(i)-(iii) and 2(i)(b)-(ii)-(iii) hold, as  $(m, T) \rightarrow \infty$*

$$\sup_{1 \leq \lfloor T\tau \rfloor \leq T} \left\| \tilde{V}_{\Sigma,\tau} - V_\Sigma \right\| = O_p \left( \frac{1}{m} \right) + O_p \left( m \frac{\ln T}{\sqrt{T}} \right). \quad (6)$$

*The same rates hold for  $\hat{V}_\Sigma$  or  $\tilde{V}_\Sigma$ .*

## Remarks

T2.1 Equation (4) is based on a SLLN for the case of no autocorrelation in  $w_t$  - see also Ling (2007). Theorem 2 provides a uniform rate of convergence for  $\hat{V}_{\Sigma,\tau}$  and  $\tilde{V}_{\Sigma,\tau}$ , as it is usually required in this literature (e.g. Lemma 2.1.2 in Csörgő and Horváth, 1997, p. 76; see also the proof of Theorem 3 below). In case of serial dependence, (5) states that it is possible to construct an estimator of  $V_\Sigma$  with a rate of convergence. This can be refined as in (6).



T2.2 A word of warning on the the weighted-sum-of-covariance estimator  $\tilde{V}_{\Sigma,\tau}$  is in order. As well documented in several contributions (we refer to Muller, 2014, and the references therein, for an exposition of the issues),  $\tilde{V}_{\Sigma,\tau}$  can be expected to suffer from (possibly severe) finite sample bias, especially in the presence of large autoregressive roots. In Section 4, we assess the robustness of  $\tilde{V}_{\Sigma,\tau}$  to the case of strong serial correlation in the data.

### *Estimation of the eigensystem*

In this section, we extend the asymptotics for the partial sample estimates of  $\Sigma$  to its eigensystem.

Let the  $i$ -th eigenvalue/eigenvector couple be defined as  $(\lambda_i, x_i)$ ; the eigenvectors are defined as an orthonormal basis, i.e.  $x_i'x_j = \delta_{ij}$ , where  $\delta_{ij}$  is Kronecker's delta. Since  $\Sigma x_i = \lambda_i x_i$ , a natural estimator for  $(\lambda_i, x_i)$  is the solution to the system

$$\begin{cases} \hat{\Sigma} \hat{X} = \hat{X} \hat{\Lambda} \\ \hat{X}' \hat{X} = I \end{cases}, \quad (7)$$

where  $\hat{X} = [\hat{x}_1, \dots, \hat{x}_n]$ ,  $\hat{x}_i$  denotes the estimate of  $x_i$ , and  $\hat{\Lambda}$  is a diagonal matrix containing the estimated eigenvalues  $\hat{\lambda}_i$  in decreasing order. Estimation of  $\{(\lambda_i, x_i)\}_{i=1}^n$  based on (7) is known as Anderson's Principal Component (PC) estimator. Similarly, the partial sample estimators of the eigenvalues and eigenvectors are the solutions to  $\hat{\Sigma}_\tau \hat{x}_{i,\tau} = \hat{\lambda}_{i,\tau} \hat{x}_{i,\tau}$ .

As we mention below (see Remark P1.2), one disadvantage of Anderson's PC estimator is that the estimated eigenvectors have a singular asymptotic covariance matrix (see Kollo and Neudecker, 1997). In order to avoid this issue, an estimator based on a different normalisation can be proposed, known as the Pearson-Hotelling's PC estimator; in this case, the estimated eigenvalues are the same as from (7), but the eigenvectors  $\gamma_i$  are defined (and estimated) as an eigenvalue-normed basis, viz.  $\gamma_i' \gamma_j = \lambda_i \delta_{ij}$ . Thus,  $\gamma_i \equiv \lambda_i^{1/2} x_i$ . A typical interpretation of the  $\gamma_i$ s in the context of the term structure of interest rates (Litterman and Scheinkman, 1991; Perignon and Villa, 2006) is that  $\lambda_i$  is the "volatility" of  $\gamma_i$ , and  $x_i$  represents its "loading". The estimates of the eigensystem according to the Pearson-Hotelling approach are the solution

to the system

$$\begin{cases} \hat{\Sigma}\hat{X} = \hat{X}\hat{\Lambda} \\ \hat{X}'\hat{X} = \hat{\Lambda} \end{cases}. \quad (8)$$

Upon calculating the solutions of (8), it turns out that the eigenvectors are estimated by  $\hat{\gamma}_i = \hat{\lambda}_i^{1/2} \hat{x}_i$ , i.e. by the same estimator for the eigenvector as in (7) multiplied by the square root of the corresponding estimate of the eigenvalue. Similarly, we define the partial sample estimator of  $\gamma_i$  as  $\hat{\gamma}_{i,\tau} = \hat{\lambda}_{i,\tau}^{1/2} \hat{x}_{i,\tau}$ .

Consider the following assumption.

**Assumption 3.** It holds that  $\min_{1 \leq i \leq n-1} (\lambda_i - \lambda_{i+1}) > 0$  with  $\lambda_n > 0$ .

Assumption 3 requires that  $\Sigma$  has distinct, strictly positive eigenvalues, and it is typical of PCA, affording to use Matrix Perturbation Theory (MPT); the assumption could be relaxed at the price of a more complicated analysis, still based on MPT. In essence, the asymptotics of  $(\hat{\lambda}_{i,\tau}, \hat{x}_{i,\tau})$  is derived by treating  $\hat{\Sigma}_\tau$  as a perturbation of  $\Sigma$ , thus deriving the expressions for the estimation errors of  $\hat{\lambda}_{i,\tau}$  and  $\hat{x}_{i,\tau}$ . The way in which the assumption is formulated is the same as in Horváth and Rice (2015, see equation (1.11)). As a consequence of the requirement that eigenvalues be strictly positive, our set-up does not directly cover the case of exact factor models, where the covariance matrix of the data has reduced rank by construction - see Han and Inoue (2011), Stock and Watson (2012) and Cheng, Liao and Schorfheide (2014).

The extension of the IP and the SIP to the eigensystem of  $\Sigma$  is reported in Proposition 1:

**Proposition 1** *Under Assumptions 1 and 3, as  $T \rightarrow \infty$ , uniformly in  $\tau$*

$$\hat{\lambda}_{i,\tau} - \lambda_i = (x'_i \otimes x'_i) \text{vec}(\hat{\Sigma}_\tau - \Sigma) + O_p(T^{-1}), \quad (9)$$

$$\hat{x}_{i,\tau} - x_i = v_{x,i} \text{vec}(\hat{\Sigma}_\tau - \Sigma) + O_p(T^{-1}), \quad (10)$$

$$\hat{\gamma}_{i,\tau} - \gamma_i = v_{\gamma,i} \text{vec}(\hat{\Sigma}_\tau - \Sigma) + O_p(T^{-1}), \quad (11)$$

where  $v_{x,i} = \left[ \sum_{k \neq i} \frac{x_k}{\lambda_i - \lambda_k} (x'_k \otimes x'_i) \right]$  and  $v_{\gamma,i} = \frac{1}{2} \frac{x_i}{\lambda_i^{1/2}} (x'_i \otimes x'_i) + \sum_{k \neq i} \frac{\lambda_i^{1/2} x_k}{\lambda_i - \lambda_k} (x'_i \otimes x'_k)$ .

## Remarks

**P1.1** Proposition 1 is the central ingredient in order to apply the test for structural breaks to the eigensystem. It states that the estimation errors  $\hat{\lambda}_{i,\tau} - \lambda_i$ ,  $\hat{x}_{i,\tau} - x_i$  and  $\hat{\gamma}_{i,\tau} - \gamma_i$  are,

asymptotically, linear functions of  $\hat{\Sigma}_\tau - \Sigma$ ; thus, the IP and the SIP in Theorem 1 carry through to the estimated eigensystem. The results in Proposition 1, and the method of proof, can be compared to related results in Kollo and Neudecker (1997).

**P1.2** By (10), the asymptotic covariance matrix of  $\sqrt{T}(\hat{x}_{i,\tau} - x_i)$  is  $v_{x,i}V_\Sigma v'_{x,i}$ . It can be shown (see e.g. Kollo and Neudecker, 1997, p. 66) that  $v_{x,i}V_\Sigma v'_{x,i}$  is singular; given that there is no obvious way to calculate the rank of  $v_{x,i}V_\Sigma v'_{x,i}$ , it is difficult to prove the consistency of the Moore-Penrose inverse for  $v_{x,i}V_\Sigma v'_{x,i}$  (see Andrews, 1987). Thus, we recommend to carry out tests on the eigenvectors using the  $\gamma_i$ s.

**P1.3** Proposition 1 shows that  $\hat{\lambda}_{i,\tau} - \lambda_i$  is linear in  $\hat{\Sigma}_\tau - \Sigma$  to the order  $O_p(T^{-1})$ ; the proof of the proposition shows that the leading order term in the approximation error is  $T^{-1} \sum_{k \neq i} [\hat{x}'_i \otimes \hat{x}'_k] \frac{\tilde{V}_\Sigma}{\hat{\lambda}_i - \hat{\lambda}_k} [\hat{x}_k \otimes \hat{x}_i]$ , so finite sample improvements may be obtained using  $\tilde{\lambda}_{i,\tau} = \hat{\lambda}_{i,\tau} - T^{-1} \sum_{k \neq i} [\hat{x}'_i \otimes \hat{x}'_k] \frac{\tilde{V}_\Sigma}{\hat{\lambda}_i - \hat{\lambda}_k} [\hat{x}_k \otimes \hat{x}_i]$ . This result is of independent interest; it could be useful e.g. when measuring the percentage of the total variance of  $y_t$  explained by each of its principal components. Similarly, in equation (36) in appendix we provide a formula to estimate the expected value of the  $O_p(T^{-1})$  order terms of  $(\hat{x}_{i,\tau} - x_i)$ ; combining these results, a bias-correction for  $\hat{\gamma}_{i,\tau}$  can also be computed.

Define  $\lambda \equiv [\lambda_1, \dots, \lambda_n]'$  as the  $n$ -dimensional vector containing the eigenvalues sorted in descending order, and  $\Gamma \equiv [\gamma_1, \dots, \gamma_n]$ ;  $\hat{z} \equiv \left[ \hat{\lambda}', \text{vec}(\hat{\Gamma})' \right]'$  with  $\hat{z}_\tau - z = D_{\lambda\gamma} \text{vec}(\hat{\Sigma}_\tau - \Sigma) + O_p(T^{-1})$  and  $D_{\lambda\gamma} \equiv [x_1 \otimes x_1, \dots, x_n \otimes x_n, v'_{\gamma,1}, \dots, v'_{\gamma,n}]'$ . The matrix  $D_{\lambda\gamma}$  can be estimated as  $\hat{D}_{\lambda\gamma} = [\hat{x}_1 \otimes \hat{x}_1, \dots, \hat{x}_n \otimes \hat{x}_n, \hat{v}'_{\gamma,1}, \dots, \hat{v}'_{\gamma,n}]'$ , with  $\hat{v}_{\gamma,i} = \frac{1}{2} \frac{\hat{x}_i}{\hat{\lambda}_i^{1/2}} (\hat{x}'_i \otimes \hat{x}'_i) + \sum_{k \neq i} \frac{\hat{\lambda}_i^{1/2} \hat{x}_k}{\hat{\lambda}_i - \hat{\lambda}_k} (\hat{x}'_i \otimes \hat{x}'_k)$ .

The asymptotics of  $\hat{z}_\tau$  follows from Theorem 1 and Proposition 1, and we summarize it below.

**Corollary 1** *Under Assumptions 1 and 3, as  $T \rightarrow \infty$ , it holds that  $\sqrt{T}(\hat{z}_\tau - z) \xrightarrow{d} [V_z]^{1/2} W_{n(2n+1)}(\tau)$ . Also, there exists a  $\delta > 0$  such that  $T(\hat{z}_\tau - z) = \sum_{t=1}^{\lfloor T\tau \rfloor} \tilde{X}_t + O_{a.s.}(\lfloor T\tau \rfloor^{\frac{1}{2}-\delta})$ , uniformly in  $\tau$ , where  $V_z = D_{\lambda\gamma} V_\Sigma D'_{\lambda\gamma}$  and  $\tilde{X}_t$  is a zero mean, i.i.d. Gaussian sequence with  $E(\tilde{X}_t \tilde{X}'_t) = V_z$ .*

Corollary 1 entails that

$$\begin{aligned} \sqrt{T}(\hat{\lambda}_\tau - \lambda) &\xrightarrow{d} [V_\lambda]^{1/2} W_n(\tau), \\ \sqrt{T} \text{vec}(\hat{\Gamma}_\tau - \Gamma) &\xrightarrow{d} [V_\Gamma]^{1/2} W_{n^2}(\tau), \end{aligned}$$

with:  $V_\lambda$  a matrix with  $(i, j)$ -th element given by  $V_{ij}^\lambda = (x'_i \otimes x'_i) V_\Sigma (x_j \otimes x_j)$ , and  $V_\Gamma$  is an  $(n^2 \times n^2)$ -dimensional matrix whose  $(i, j)$ -th  $n \times n$  block is defined as  $V_{ij}^\Gamma = v_{\gamma,i} V_\Sigma v'_{\gamma,j}$ .

### 3 Testing

This section studies the null distribution and the consistency of tests based on CUSUM-type statistics.

Henceforth, we define the CUSUM process  $S(\tau) = \sum_{t=1}^{\lfloor T\tau \rfloor} \text{vec}(y_t y'_t)$ . In light of Corollary 1, test statistics for  $\Sigma$  and its eigensystem can be based on

$$\tilde{S}(\tau) = R \times D_{\lambda\gamma} \times \left[ S(\tau) - \frac{\lfloor T\tau \rfloor}{T} S(1) \right], \quad (12)$$

with  $\tilde{S}(\tau) = 0$  for  $\tau \leq \frac{1}{T}$  or  $\geq 1 - \frac{1}{T}$ , and  $R$  a  $p \times n(n+1)$  matrix. For example, when testing for the null of no changes in the largest eigenvalue,  $R$  is the matrix that extracts the first element of  $D_{\lambda\gamma} \times \left[ S(\tau) - \frac{\lfloor T\tau \rfloor}{T} S(1) \right]$ . Thence, testing is carried out by using (the supremum of)

$$\Lambda_T(\tau) = \sqrt{\frac{T}{\lfloor T\tau \rfloor \times \lfloor T(1-\tau) \rfloor}} \times \left[ \tilde{S}(\tau)' \tilde{V}_{z,\tau}^{-1} \tilde{S}(\tau) \right]^{1/2}, \quad (13)$$

with  $\tilde{V}_{z,\tau} = R D_{\lambda\gamma} \tilde{V}_{\Sigma,\tau} D'_{\lambda\gamma} R'$ . The test statistic defined in (13) can be compared with the one proposed by Aue et al. (2009), which, in our context, would be based on (the supremum of)

$$\Lambda_T^A(\tau) = \sqrt{\frac{1}{T}} \times \left[ \tilde{S}(\tau)' \tilde{V}_{z,\tau}^{-1} \tilde{S}(\tau) \right]^{1/2}. \quad (14)$$

Contrasting (13) with (14), it is clear that the only difference between the two test statistics is the norming factors,  $\sqrt{\frac{T}{\lfloor T\tau \rfloor \times \lfloor T(1-\tau) \rfloor}}$  versus  $\sqrt{\frac{1}{T}}$ . However, such difference is crucial: by virtue of the weighing scheme proposed in (13), we are able to detect the presence of breaks closer to either end of the sample than afforded by (14). More specific comments on the power properties of tests based on (13) versus tests based on (14) are in the remarks to Theorem 4; here we point out that the price to pay is that we are not able to study the limiting distribution of the supremum of (13) using the IP shown in Theorem 1, but conversely the SIP is needed.

Theorem 3 contains the asymptotics of  $\sup_{\lfloor T\tau \rfloor} \Lambda_T(\tau)$  under the null.

**Theorem 3** Under Assumptions 1-3, as  $(m, T) \rightarrow \infty$  with  $\frac{1}{m} + m\frac{\ln T}{\sqrt{T}} \rightarrow 0$ ,

$$\sup_{\lfloor T\tau_1 \rfloor \leq \lfloor T\tau \rfloor \leq \lfloor T\tau_2 \rfloor} \Lambda_T(\tau) \xrightarrow{d} \sup_{\tau_1 \leq \tau \leq \tau_2} \frac{\|B_p(\tau)\|}{\sqrt{\tau(1-\tau)}}, \quad (15)$$

where  $B_p(\tau)$  is a  $p$ -dimensional standard Brownian bridge and  $[\tau_1, \tau_2] \subset (0, 1)$ . Also, as  $(m, T) \rightarrow \infty$  with  $\frac{\sqrt{\ln \ln T}}{m} + m \ln T \sqrt{\frac{\ln \ln T}{T}} \rightarrow 0$ ,

$$P \left\{ a_T \left[ \sup_{n \leq \lfloor T\tau \rfloor \leq T-n} \Lambda_T(\tau) \right] \leq x + b_T \right\} \rightarrow e^{-2e^{-x}}, \quad (16)$$

where  $a_T = \sqrt{2 \ln \ln T}$  and  $b_T = 2 \ln \ln T + \frac{p}{2} \ln \ln T - \ln \Gamma\left(\frac{p}{2}\right)$ , with  $\Gamma(\cdot)$  the Gamma function.

### Remarks

T3.1 According to (15), the maximum is taken in a subset of  $[0, 1]$ , namely  $[\tau_1, \tau_2]$ . This approach requires an IP for  $S(\tau)$ , and the Continuous Mapping Theorem (CMT). As noted in Corollary 1 in Andrews (1993, p. 838),  $\Lambda_T(\tau)$  is not continuous at  $\{0, 1\}$  and  $\sup_{1 \leq \lfloor T\tau \rfloor \leq T} \Lambda_T(\tau) \xrightarrow{P} \infty$  under  $H_0$ . Thus, trimming is necessary in this case. Further, in this case it suffices to have a consistent estimator of the long-run covariance matrix  $V_\Sigma$  which, in light of equation (6) in Theorem 2, entails that  $m \rightarrow \infty$  with  $m = o(T)$ . The considerations in Remark T2.1 apply here.

T3.2 As an alternative approach, the SIP can be used: sums of  $\bar{w}_t$  can be replaced by sums of *i.i.d.* Gaussian variables, with an approximation error. Upon normalising  $\Lambda_T(\tau)$  with the appropriate norming constants, say  $a_T$  and  $b_T$ , an Extreme Value (EV henceforth) theorem can be employed. Tests based on  $\sup_{n \leq \lfloor T\tau \rfloor \leq T-n} [a_T \Lambda_T(\tau) - b_T]$  are designed to be able to detect breaks close to the end of the sample. Results like (16) have been derived by Horváth (1993), for *i.i.d.* Gaussian data, and extended to the case of dependence by Ling (2007), *inter alia*. As far as the long-run covariance matrix estimator is concerned, in this case the theory requires a consistent estimator at a rate (at least)  $o_p \left[ \left( \sqrt{\ln \ln T} \right)^{-1} \right]$ : therefore, from (6), we need the restrictions  $\frac{\sqrt{\ln \ln T}}{m} \rightarrow 0$  and  $m \ln T \sqrt{\frac{\ln \ln T}{T}} \rightarrow 0$ .

### Consistency of the test

We now turn to studying the behaviour of  $\sup_{n \leq \lfloor T\tau \rfloor \leq T-n} \Lambda_T(\tau)$  under alternatives. As a leading example, we consider the case of testing for no change in  $\Sigma$  in presence of one abrupt

change

$$H_a^{(T)} : \text{vech}(\Sigma_t) = \begin{cases} \text{vech}(\Sigma) & \text{for } t = 1, \dots, k_{0,T} \\ \text{vech}(\Sigma) + \Delta_T & \text{for } t = k_{0,T} + 1, \dots, T \end{cases}, \quad (17)$$

where both the changepoint  $(k_{0,T})$  and the size of the break  $(\Delta_T)$  could depend on  $T$ . More general alternatives could be considered (see e.g. Andrews, 1993; Csörgő and Horváth, 1997): these include epidemic alternatives, and also breaks that occur as a smooth transition over time as opposed to abruptly as in (17). Further, note that (17) does not rule out the possibility that only some series (i.e. only some of the coordinates of  $y_t$ ) actually have a break. This entails that tests based on  $\Lambda_T(\tau)$  are capable of detecting breaks that only affect some of the series, and possibly at different points in time.

Theorem 4 illustrates the dependence of the power on  $\Delta_T$  and  $k_{0,T}$ .

**Theorem 4** *Let Assumptions 1-3 hold, and define  $c_{\alpha,T}$  such that, under  $H_0$ ,  $P \left[ \sup_{n \leq \lfloor T\tau \rfloor \leq T-n} \Lambda_T(\tau) \leq c_{\alpha,T} \right] = 1 - \alpha$  for some  $\alpha \in [0, 1]$ . If, under  $H_a^{(T)}$ , as  $T \rightarrow \infty$*

$$\frac{1}{\ln \ln T} \left[ \frac{(T - k_{0,T}) k_{0,T}}{T} \|RD_{\lambda\gamma} \Delta_T\|^2 \right] \rightarrow \infty, \quad (18)$$

*it holds that*

$$P \left[ \sup_{n \leq \lfloor T\tau \rfloor \leq T-n} \Lambda_T(\tau) > c_{\alpha,T} \right] \rightarrow 1. \quad (19)$$

### Remarks

T4.1 Theorem 4 illustrates the impact of  $k_{0,T}$  and  $\Delta_T$  on the power of tests based on  $\sup_{n \leq \lfloor T\tau \rfloor \leq T-n} \Lambda_T(\tau)$ . Particularly, consider the two extreme cases:

T4.1.a  $\|\Delta_T\| = O(1)$ , i.e. finite break size. In this case, the test has power as long as  $k_{0,T}$  is strictly bigger than  $O(\ln \ln T)$ . This can be compared with tests based on  $\sup_{1 \leq \lfloor T\tau \rfloor \leq T} T^{-1} \tilde{S}(\tau)' \tilde{V}_{z,\tau}^{-1} \tilde{S}(\tau)$ , which can be shown to have nontrivial power in presence of finite breaks at most as close as  $O(\sqrt{T})$  to either end of the sample. Using similar algebra as in the proof of Theorem 4, it can be shown that the noncentrality parameter of  $\sup_{1 \leq \lfloor T\tau \rfloor \leq T} T^{-1} \tilde{S}(\tau)' \tilde{V}_{z,\tau}^{-1} \tilde{S}(\tau)$  is proportional to  $\|\Delta_T\|^2 \frac{k_{0,T}^2}{T}$ . Under  $\|\Delta_T\| = O(1)$ , this entails that nontrivial power is attained as long as  $k_{0,T} = O(\sqrt{T})$ .

T4.1.b  $k_{0,T} = O(T)$  - i.e. the break occurs in the middle of the sample. The test is powerful as long as the size of the break is strictly bigger than  $O\left(\sqrt{\frac{\ln \ln T}{T}}\right)$ . When using

trimmed statistics such as in (15), the test is powerful versus mid-sample alternatives of size  $O\left(\frac{1}{\sqrt{T}}\right)$ : when no trimming is used, there is some, limited loss of power versus mid-sample alternatives.

T4.2 Equation (18) also indicates that the test has no power when  $RD_{\lambda\gamma}\Delta_T = 0$  (or whenever it is “very small”) This could e.g. happen in the case of having a break, however massive, in the eigenvalue  $\lambda_i$ , and applying the test for a change in eigenvalue  $\lambda_j$ ,  $j \neq i$ ; such a test is bound to have no ability to detect a change in  $\lambda_i$ , by construction.

In the Supplemental Material (Kao et al., 2015) we also show that all the results developed above also hold when applied to residuals - that is, one can test for the stability of the covariance matrix (and its eigensystem) of the error term in the multivariate regression (including e.g. a VAR)

$$y_t = \beta x_t + \varepsilon_t, \quad (20)$$

where  $t = 1, \dots, T$  and  $y_t$  and  $\varepsilon_t$  are  $n \times 1$  vectors,  $x_t$  is of dimension  $q \times 1$  (and results can be extended to also include linear or polynomial trends in  $x_t$ ) and the matrix of regressors  $\beta$  has dimensions  $n \times q$ . As shown in the Supplemental Material (Kao et al., 2015), the extension to residuals only requires that  $x_t\varepsilon_t$  and  $\varepsilon_t$  satisfy similar assumptions to the one spelt out above.

#### *Computation of critical values*

Based on Theorem 3, there are two possible approaches to the computation of critical values: either using the EV distribution in (16) or using an approximation similar to that proposed in Csörgő and Horváth (1997, Section 1.3.2).

Direct computation of critical values  $c_{\alpha,T}$  for a test of level  $\alpha$  is based on  $c_{\alpha,T} = a_T^{-1} \{b_T - \ln[-\frac{1}{2} \ln(1 - \alpha)]\}$ . Thus, critical values only depend on  $p$  and  $T$ . It is well known that convergence to the EV distribution is usually very slow, which hampers the quality of  $c_{\alpha,T}$ . Alternatively, critical values can be simulated from

$$P \left\{ \sup_{h_{nT} \leq \tau \leq 1-h_{nT}} \left[ \sum_{i=1}^p \frac{B_{1,i}^2(\tau)}{\tau(1-\tau)} \right]^{1/2} \leq c'_{\alpha,T} \right\} = 1 - \alpha, \quad (21)$$

where the  $B_{1,i}(\tau)$ s are independent, univariate Brownian bridges, generated over a grid of dimension  $T$ . We set  $T \times h_{nT} = \max\{n, \ln^{3/2} T\}$ . The “time series” part of this bound (i.e. the  $\ln^{3/2} T$  part) is based on Csörgő and Horváth (1997, p. 25), who show that computing

the maxima over restricted intervals (specifically, by truncating at  $T \times h_{nT} = \ln^{3/2} T$ ) yields tests with good size properties; in our simulations, we have tried other solutions to restrict the interval over which the maximum is taken, but truncating at  $\ln^{3/2} T$  yielded the best size properties. In addition to this, due to the multivariate nature of the problem, we also need to truncate at  $n$ ; this is in order to have full rank estimated covariance matrices. In view of this, critical values  $c'_{\alpha,T}$  are to be simulated for a given combination of  $p$ ,  $n$  and  $T$ . For the purpose of comparison, critical values for the test statistic defined in (14), based on Aue et al. (2009), are computed by using the largest value taken by  $\sum_{i=1}^p B_{1,i}^2(\tau)$  across the whole grid for each simulation.

## 4 Monte Carlo evidence

We evaluate size and power through a Monte Carlo exercise. Data are generated according to the following DGP:

$$y_t = \rho y_{t-1} + e_t + \theta e_{t-1}. \quad (22)$$

Under the null, we simulate  $e_t$  as *i.i.d.*  $N(0, I_n)$ . Our experiments are conducted by setting  $(\rho, \theta) = \{(0, 0), (0.5, 0), (0, 0.5), (0, -0.5)\}$ ; as far as the sample size  $T$ , and the matrix dimension  $n$ , are concerned, experiments are reported for  $T = \{50, 200, 500\}$  and  $n = \{3, 10\}$ . Finally, in order to avoid dependence on initial conditions,  $T + 1000$  data are generated, discarding the first 1000 observations.

As far as the test is concerned, this is based on

$$\sup_{Th_{nT} \leq \lfloor T\tau \rfloor \leq T - Th_{nT}} \Lambda_T(\tau), \quad (23)$$

where  $h_{nT}$  is defined above as  $h_{nT} = \max \left\{ \frac{n}{T}, \frac{\ln^{3/2} T}{T} \right\}$ . In all experiments, we use the long run variance estimator in (3), based on full sample estimation of the autocovariance matrices with  $m = T^{2/5}$ .

### *Testing for changes in the largest eigenvalue*

In the first set of experiments, we test for the null of no changes in the largest eigenvalue of



$\Sigma$ . Under the alternative, breaks in  $E(e_t e'_t)$  are defined as

$$\begin{cases} I_n & \text{for } t = 1, \dots, k \\ I_n + \Delta & \text{for } t = k + 1, \dots, T \end{cases} . \quad (24)$$

Breaks are evaluated according to the following schemes

$$k = \left\lfloor \frac{T}{2} \right\rfloor \text{ and } \Delta = \sqrt{\frac{\ln \ln T}{T^{2/3}}} \times I_n, \quad (25)$$

$$k = \left\lfloor \frac{T}{2} \right\rfloor \text{ and } \Delta = \sqrt{\frac{\ln \ln T}{T^{1/2}}} \times I_n, \quad (26)$$

$$k = Th_{nT} + 1, k = \frac{1}{2} (\ln T)^2, k = \frac{1}{2} (\ln T)^{5/2} \text{ and } k = 3\sqrt{T}; \Delta = I_n \quad (27)$$

The first two alternatives consider power versus mid-sample breaks; the last set of alternatives considers breaks of finite magnitude that are close to the beginning of the sample.

**[Insert Tables 1-4 somewhere here]**

We note that:

1. as far as *size* is concerned, considering a 5% level, Table 1 shows that the test is, in general, undersized in small samples; this tends to disappear as  $T$  increases, with empirical rejection frequencies belonging, in general, to the interval  $[0.04, 0.06]$  with few exceptions. Interestingly, higher values of  $n$  have a slight tendency to reduce the size. Similar results are found with Aue et al.'s (2009) test based on (14);
2. as far as *power* is concerned:
  - a. mid-sample breaks are studied in Tables 2-3, which correspond to cases (25) and (26) respectively. The test has good power, with the power increasing as  $n$  increases. As predicted by the theory, the test by Aue et al.'s (2009) has higher power. Note the adverse impact of higher serial correlation on both tests;
  - b. breaks close to the beginning of the sample are considered in Table 4, corresponding to equation (27). The test has power versus finite alternatives that are close to the beginning of the sample, and the power increases with  $n$ ;

- as is natural, Aue et al.'s (2009) test has very little power versus beginning of sample alternatives; by construction, such tests do not have power versus changes that occur closer than  $O(\sqrt{T})$  periods to the beginning (or the end) of the sample; again, note the adverse effect of higher serial correlation on the power of both tests.

#### *Testing for changes in the covariance matrix*

We also carry out a second set of experiments to evaluate the performance of the test when applied to detect a change in  $E(y_t y_t')$ . The test is based on the null that all eigenvalues are constant - that is, it is an *omnibus* test for breaks in the trace of  $\Sigma$ . We consider one mid-sample break (based on equation (26)) and one end-of-sample break (based on equation (27)); the full set of results is in Tables A1-A3 in the Supplemental Material (Kao et al., 2015).

The main findings are as follows:

1. as far as *size* is concerned, as  $n$  increases, the test becomes increasingly conservative in finite samples; however, as  $T \rightarrow \infty$ , the empirical rejection frequencies tend towards their nominal values;
2. as far as *power* is concerned:
  - a. under mid-sample alternatives, the power increases monotonically with  $T$  as expected. As far as  $n$  is concerned, the power seems to have a mild tendency to increase with  $n$ . As expected, in this context our test is less powerful than the one proposed by Aue et al. (2009), and it has power higher than 50% when  $T \geq 200$ ;
  - b. under end-of-sample alternatives, as  $n$  increases, the power also increases. As expected, our test is decidedly more powerful than the test by Aue et al. (2009), at least for large samples ( $T \geq 200$ ). Neither test has satisfactory power when the sample size is small;
  - c. in both experiments, we also considered a break of equal magnitude and location as above, but only for the first element in the matrix, i.e. the volatility of the first series. We considered the case of *i.i.d.* data only. Results are comparable with the rest of the tables.

### Other experiments

We conducted some more limited experiments to assess how the test works in presence of “boundary” situations - such as a nearly singular covariance matrix (which nearly violates Assumption 3), or a highly persistent autoregressive process (which is bound to hamper the performance of the weighted-sum-of-covariances estimators of the long-run variance  $V_\Sigma$ ).

1. The case of a *nearly singular covariance matrix* (Table B1 in the Supplementary Material) has been simulated by using  $e_t \sim N(0, C_n)$  in the DGP defined in equation (22), where  $C_n$  is an  $n$ -dimensional diagonal matrix defined as

$$\{C_n\}_{ii} = \begin{cases} 1 & \text{for } i = 1 \\ U[0, 0.02] & \text{for } 2 \leq i \leq n \end{cases}. \quad (28)$$

This set-up, with one large eigenvalues and the others being very small, corresponds to the case of having a factor model. By way of comparison, we also carried out the same exercise, but with data generated by setting  $e_t \sim N(0, I_n)$ . We test for the stability of the first principal component, considering size and power versus mid-sample and end-of-sample alternatives: in presence of very small eigenvalues, the test still has good size and power properties, although power is better (especially as  $n$  grows) when eigenvalues are of comparable magnitude.

2. The case of *highly autocorrelated data* (Table B2 in the Supplementary Material) has been simulated using the following variant of the DGP defined in (22)

$$y_t = 0.9y_{t-1} + e_t. \quad (29)$$

Without pre-whitening, the test is so grossly oversized (empirical rejection frequencies, under the null, are well above 50%) that we do not even report the results: the basic message is that the test cannot be employed in presence of highly correlated data. This is essentially due to the poor performance of the long-run variance estimator; unreported experiments where the test is carried out using the population long-run variance reinforce this conjecture. As a solution, we suggest pre-whitening, which in our case we carry out by estimating a VAR(1) and using a short bandwidth chosen as  $m = T^{1/4}$ : in this case, results are very good in terms of power and size. By way of robustness check, we have

also tried to assess whether, in presence of a mis-specified pre-whitening, the test works well - to this end, we have simulated data as

$$y_t = 0.9y_{t-1} + e_t + 0.9e_{t-1},$$

with pre-whitening being carried out as before - i.e. by using a VAR(1). Results show that even when pre-whitening is not correctly specified, the test has the correct size, and good power versus mid-sample alternatives; however, the power versus breaks close to either end of the sample is significantly lower when the pre-whitening is not correctly specified.<sup>1</sup>

## 5 Application: the time stability of the covariance matrix of interest rates

In this section, we apply the theory developed above to test for the stability of the covariance matrix of the term structure of interest rates - returns, computed as log differences of zero-coupon bond prices are used, since preliminary analysis shows that the yields are highly persistent. Our analysis is motivated by the study in Perignon and Villa (2006), and follows similar steps.

As a first step, we investigate whether the “volatility curve” (i.e. the term structure of the volatility of interest rates) changes over time; this corresponds to testing for the stability of the main diagonal of the covariance matrix. Further, we verify whether the whole covariance matrix changes. This could be done by directly testing for the constancy of the matrix. Alternatively, in order to reduce the dimensionality of the problem, one could check whether the main three principal components (customarily known as level, slope and curvature) are stable through time. We choose the latter approach, verifying separately, for each principal component, whether sources of time variation are in the loadings (i.e. the eigenvectors) or in the volatility (i.e. the eigenvalues), or both.

Previous studies have found evidence of changes in the yield curve. Using a descriptive approach based on splitting the sample at some predetermined points in time, indicated by stylised facts, Bliss (1997) finds that the eigenvectors of the covariance matrix of interest rates are quite stable, although the eigenvalues differ across subsamples. Perignon and Villa (2006),

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<sup>1</sup>It should be noted that pre-whitening is only one possible approach - an alternative of increasing popularity in the econometric literature is to use a fixed bandwidth approach, setting  $m = cT$ ; we refer to Muller (2014) for an analysis of this approach that goes beyond the scope of the present paper.

under the assumption that data are *i.i.d.* Gaussian, find evidence of changes in the volatilities (eigenvalues) of the principal components across four different subperiods (chosen *a priori*) in the time interval January 1960 - December 1999.

We apply our test to US data, considering monthly and weekly frequencies, spanning from April 1997 to November 2010 (monthly - the sample size is  $T_m = 164$ ) and from the first week of April 1997 to the last week of November 2010 (weekly - the sample size is  $T_w = 713$ ); the use of different frequencies within the same endpoint may be helpful to show whether the properties of the data depend on their frequency or not. The number of maturities which we consider is  $n = 18$ , corresponding to (1m, 3m, 6m, 9m, 12m, 15m, 18m, 21m, 24m, 30m, 3y, 4y, 5y, 6y, 7y, 8y, 9y, 10y). Figure 1 reports the term structure in the period considered.

**[Insert Figure 1 somewhere here]**

In the Supplemental Material, we also report some descriptive statistics (Table C). Since there seems to be some serial correlation (at least with lower maturities), we pre-whiten the data using the VAR(1) scheme employed in the previous section. We let  $y_t$  denote, henceforth, the demeaned 18-dimensional vector of maturities. The first step of our analysis is an evaluation of the stability of the variances, i.e. of the elements on the main diagonal of  $\Sigma = E(y_t y_t')$ . Instead of checking for the stability of the whole main diagonal, we test the volatilities one by one; this approach should be more constructive if the null of no changes were to be rejected, in that it would indicate which maturity changes and when. In order to control for the size of this multiple comparison, we propose a Bonferroni correction, computing the critical values for each test as  $\alpha_I = \frac{\alpha_P}{n}$ , where  $\alpha_P$  is the size of the whole procedure. Using these critical values yields, approximately, a level  $\alpha_P$  not greater than 1%, 5% and 10% corresponds to conducting each test at levels  $\alpha_I = 0.056\%$ ,  $0.28\%$  and  $0.56\%$  respectively.

As a second step, we verify whether the first three principal components are constant over time. Particularly, we carry out separately the detection of changes in the volatility of the principal components (verifying the time stability of the three largest eigenvalues, say  $\lambda_1$ ,  $\lambda_2$  and  $\lambda_3$ ), and in their loading (verifying the stability of the eigenvalue-normed eigenvectors corresponding to the three largest eigenvalues, denoted as  $\gamma_1$ ,  $\gamma_2$  and  $\gamma_3$ ). As far as eigenvectors are concerned, (10) and (11) ensure that, when running the test, the CUSUM transformation of the estimated  $\gamma_i$ s has the same sign for all values of  $\tau$ , thus overcoming the issue of the eigenvectors being defined up to a sign.

Results for both experiments, at both frequencies, are reported in Table 5 (critical values are in Table D in the Supplemental Material (Kao et al., 2015)).

**[Insert Table 5 somewhere here]**

Interestingly, when using a 5% level, rejections occur for the same maturities, whether one uses the Bonferroni correction or not. The only exception is the test for the stability of the second eigenvector,  $\gamma_2$ , when using weekly data, where the null of no change is now rejected at 5%. A marginal discrepancy can be observed in Panel A of Table 5, when testing for the constancy of the diagonal elements of  $\Sigma$  with weekly data. When considering a single hypothesis testing approach, two maturities (the 30 months and the 3 years ones) now appear to have a break. The rest of the results (especially the absence of breaks in monthly data) is the same as when using a Bonferroni correction.

Table 5 shows some discrepancy between monthly and weekly data. Monthly data, as a whole, have a stabler covariance structure over time, with no changes in the volatilities of the maturities, or in any of the principal components. Indeed, the only instability is observed in the eigenvalue structure (Panel B):  $\lambda_3$ , the volatility of the curvature, has a break significant at 5%. The corresponding estimated breakdate, selected as the maximizer of the CUSUM statistic, is January 2008. As far as weekly data are concerned, there is evidence of instability in the covariance structure. At a “macro” level, the variances of longer maturities (from 5 years onwards) change, whilst the variances of shorter maturities are constant (see Panel A). For most maturities, the breakdate is around the first week of December 2007, which is generally associated with the deepening of the recent recession. It is interesting to note that the longest maturities (9 and 10-year ones) have a break at around the last week of August 2008. As far as principal components are concerned, Panel B of Table 5 shows that whilst the volatility of slope and curvature does not change over time, the loading of the level changes at the first week of December 2007, consistently with the findings for the variances. As Panel C of the table shows, the loadings of principal components are subject to change: the level and the curvature change significantly around the middle/end of March 2008 (possibly due to an “attraction” effect of the variance of the 10-year maturity); the slope has a significant break also, a few weeks later. The presence of significant changes in the loadings of each principal component as a result of the 2007-2009 recession is a different feature to what Perignon and Villa (2006) found in the time period they consider, when eigenvectors were not subject to changes over time.

Finally, we report the proportion of the total variance explained by each principal component before and after this date.

[Insert Table 6 somewhere here]

## 6 Conclusions

In this paper, we propose a test for the null of no breaks in the eigensystem of a covariance matrix. The assumptions under which we derive our results are sufficiently general to accommodate for a wide variety of datasets. We show that our test is powerful versus alternatives as close to the boundaries of the sample as  $O(\ln \ln T)$ . Results are extended to testing for the stability of the eigensystem. We also derive a correction for the finite sample bias when estimating eigenvalues and eigenvectors, which can be relatively severe for large  $n$  or small  $T$ . The theory is also extended to develop tests for the null of no change in the covariance matrix of the error term in a multivariate regression (including the case of VARs; see the Supplemental Material (Kao et al., 2015)). As shown in Section 4, the properties of the test are satisfactory: the correct size is attained under various degrees of serial dependence, and the test exhibits good power.

The results in this paper suggest several avenues for research. An important issue is the specification of the long-run variance estimator when implementing the test. Monte Carlo evidence suggests that employing the estimator with pre-whitening, subsequently choosing a small bandwidth, yields good results - this could be an initial guideline for the applied user. Also, the theory is derived under the minimal assumption that the 4-th moment exists. Aue et al. (2009) provide a discussion as to how to proceed if this is not the case, which involves fractional transformations of the series, viz.  $y_{it}^\Delta$  for some  $\Delta \in (0, 1)$ , although the optimal choice of  $\Delta$  is not straightforward. Also, the estimator of the long-run variance  $V_\Sigma$  proves to be crucial in affecting the properties of the test. These issues are currently under investigation by the authors.

Supplemental Material [doi]COMPLETED BY THE TYPESETTER .pdf We provide technical Lemmas, and further Monte Carlo output.

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## Appendix: Proofs of the main results

**Proof of Theorem 1.** The proof of (1) is essentially based on checking the validity of the assumptions in Theorem 29.6 in Davidson (1994, p. 481) for the normalized sequence  $\bar{w}_{T,t} = V_{\Sigma,T}^{-1/2} \bar{w}_t$ . In light of Lemma 2 in the Supplemental Material (Kao et al., 2005),  $\bar{w}_{T,t}$ , for given values of  $\alpha$  and  $r$  in Assumption 2, is  $L_2$ -NED on the strong mixing base  $\{v_t\}_{t=-\infty}^{+\infty}$  with size  $\alpha' > \frac{1}{2}$ , which entails the validity of Assumption (c) in Davidson (1994; Theorem 29.6). Assumption 1(ii) implies that  $E(\bar{w}_{T,t}) = V_{\Sigma,T}^{-1/2} E(\bar{w}_t) = 0$ . Assumption (b) in Theorem 29.6 in Davidson (1994, p. 482) follows from Assumption 1(ii) and from noting that, in light of Assumption 1(i),  $\sup_t E(\|\bar{w}_t\|^{r/2}) < \infty$ . Assumptions (d) and (f) in Theorem 29.6 in Davidson (1994) are implied by Assumption 1(iii). Finally, Assumption (e) follows from the LLN entailed by Assumptions 1(iii). Thus, (1) holds.

As far as (2) is concerned, its proof is based on Theorem 1 in Eberlein (1986, p. 263). Lemma 2 in the Supplemental Material (Kao et al., 2005) entails that  $\bar{w}_t$  is a zero-mean  $L_{2+\epsilon}$ -mixingale of size  $\alpha'' > \frac{1}{2}$ . Letting  $\mathfrak{S}_m = \{\bar{w}_1, \dots, \bar{w}_m\}$  and  $S_{Tm} \equiv \sum_{t=m+1}^{m+T} \bar{w}_t$ , (2) follows if  $|E[S_{Tm} | \mathfrak{S}_m]|_2 < \infty$  and  $|E[S_{iTm} S_{jTm} | \mathfrak{S}_m] - E[S_{iTm} S_{jTm}]| = O(T^{1-\theta})$  for  $\theta > 0$  and all  $i, j$ . Both conditions can be proved following the same passages as in Corradi (1999, p. 651-652).

■

**Proof of Theorem 2.** The proof is similar to the proof of Lemma 2.1.1 in Csörgő and Horváth (1997, p. 74-75). In view of Lemma 3 in the Supplemental Material (Kao et al., 2005), a SLLN holds (see Ling, 2007, Theorem 2.1), whereby for all  $l$

$$\frac{1}{[T\tau]} \sum_{t=1}^{[T\tau]} \text{vec} [\bar{w}_t \bar{w}'_{t-l} - E(\bar{w}_t \bar{w}'_{t-l})] = o_{a.s.} \left( \frac{1}{[T\tau]^{\delta'}} \right);$$

similarly,  $\hat{\Sigma}_\tau - \Sigma = o_{a.s.}([T\tau]^{-\delta'})$ , since  $w_t$  also satisfies the assumptions needed for Theorem 2.1 in Ling (2007). This entails that, for any  $\varepsilon > 0$  and  $\varepsilon' > 0$ , there is an integer  $g_T = g_T(\varepsilon, \varepsilon')$  such that

$$\begin{aligned} P \left[ \sup_{g_T \leq [T\tau] \leq T} [T\tau]^{\delta'} \left\| \hat{\Psi}_{l,\tau} - \Psi_l \right\| > \varepsilon \right] &\leq \varepsilon', \\ P \left[ \sup_{1 \leq [T\tau] \leq T - g_T} [T\tau]^{\delta'} \left\| \hat{\Psi}_{l,\tau} - \Psi_l \right\| > \varepsilon \right] &\leq \varepsilon'. \end{aligned}$$

These yield  $\sup_{1 \leq [T\tau] \leq T} \left\| \hat{\Psi}_{l,\tau} - \Psi_l \right\| = o_p \left( \frac{1}{T^{\delta'}} \right)$ . This proves (4).

In order to prove (5), note that

$$\begin{aligned}
\left\| \tilde{V}_{\Sigma, \tau} - V_{\Sigma} \right\| &\leq \left\| \left( \hat{\Psi}_{0, \tau} - \tau \Psi_0 \right) + \left( \hat{\Psi}_{0, 1-\tau} - (1-\tau) \Psi_0 \right) \right. \\
&\quad \left. + 2 \sum_{l=1}^m \left( 1 - \frac{l}{m} \right) \left[ \left( \hat{\Psi}_{l, \tau} - \tau \Psi_l \right) + \left( \hat{\Psi}_{l, 1-\tau} - (1-\tau) \Psi_l \right) \right] \right\| \\
&\quad + 2 \sum_{l=1}^m \frac{l}{m} \|\Psi_l\| + 2 \sum_{l=m+1}^{\infty} \|\Psi_l\| \\
&= I + II + III.
\end{aligned}$$

Note first that Assumption 2(i)(b) entails  $III = o(m^{-s})$ ; clearly, this holds uniformly in  $\tau$ . Also, again by Assumption 2(i)(b),  $II = 2m^{-1}O(1) = O(m^{-1})$ , again uniformly in  $\tau$ . We now study  $I$ ; in particular, we will consider the quantity  $\sum_{l=0}^m \left( 1 - \frac{l}{m} \right) \left( \hat{\Psi}_{l, \tau} - \tau \Psi_l \right)$ . Letting  $\hat{w}_t = w_t - \text{vec}(\hat{\Sigma}_{\tau})$ , we have

$$\begin{aligned}
&E \left\| \sum_{l=0}^m \left( 1 - \frac{l}{m} \right) \frac{1}{T} \sum_{t=1}^{\lfloor T\tau \rfloor} \left( \hat{w}_t \hat{w}'_{t-l} - \Psi_l \right) \right\|^2 \\
&\leq T^{-2} \sum_{l=0}^m \sum_{h=0}^m E \left[ \left\| \sum_{t=1}^{\lfloor T\tau \rfloor} \left( \hat{w}_t \hat{w}'_{t-l} - \Psi_l \right) \right\| \left\| \sum_{t=1}^{\lfloor T\tau \rfloor} \left( \hat{w}_t \hat{w}'_{t-h} - \Psi_h \right) \right\| \right] \\
&\leq T^{-2} \sum_{l=0}^m \sum_{h=0}^m E^{1/2} \left\| \sum_{t=1}^{\lfloor T\tau \rfloor} \left( \hat{w}_t \hat{w}'_{t-l} - \Psi_l \right) \right\|^2 E^{1/2} \left\| \sum_{t=1}^{\lfloor T\tau \rfloor} \left( \hat{w}_t \hat{w}'_{t-h} - \Psi_h \right) \right\|^2;
\end{aligned}$$

we know by the proof of Theorem 2.1 in Ling (2007) that there is a constant  $\delta' > 0$  such that  $E^{1/2} \left\| \sum_{t=1}^{\lfloor T\tau \rfloor} \left( \hat{w}_t \hat{w}'_{t-l} - \Psi_l \right) \right\|^2 = O(\lfloor T\tau \rfloor^{1-\delta'})$ ; therefore,

$$E \left\| \sum_{l=0}^m \left( 1 - \frac{l}{m} \right) \frac{1}{T} \sum_{t=1}^{\lfloor T\tau \rfloor} \left( \hat{w}_t \hat{w}'_{t-l} - \Psi_l \right) \right\|^2 = O\left(m^2 \lfloor T\tau \rfloor^{-2\delta'}\right),$$

which entails (see Moricz, 1983)

$$E \sup_{0 \leq m' \leq m} \sup_{1 \leq \lfloor T\tau \rfloor \leq T} \left\| \sum_{l=0}^{m'} \left( 1 - \frac{l}{m'} \right) \frac{1}{T} \sum_{t=1}^{\lfloor T\tau \rfloor} \left( \hat{w}_t \hat{w}'_{t-l} - \Psi_l \right) \right\|^2 = O\left(m^2 T^{-2\delta'} \ln m \ln T\right),$$

and note that  $\ln m \leq \ln T$ . Hence, it can be shown that

$$\begin{aligned} & \sup_{1 \leq \lfloor T\tau \rfloor \leq T} \left\{ \left[ \left( \hat{\Psi}_{0,\tau} - \tau \Psi_0 \right) + \left( \hat{\Psi}_{0,1-\tau} - (1-\tau) \Psi_0 \right) \right] + \right. \\ & \quad \left. 2 \sum_{l=1}^m \left( 1 - \frac{l}{m} \right) \left[ \left( \hat{\Psi}_{l,\tau} - \tau \Psi_l \right) + \left( \hat{\Psi}_{l,1-\tau} - (1-\tau) \Psi_l \right) \right] \right\} \\ & = O_p \left( m T^{-\delta'} \ln T \right). \end{aligned}$$

Thus, (5) follows.

Equation (6) follows from the same passages as above; however, Lemma 4 in the Supplemental Material (Kao et al., 2005) implies that

$$E \left\| \sum_{t=1}^{\lfloor T\tau \rfloor} \left( \hat{w}_t \hat{w}_{t-l}' - \Psi_l \right) \right\|^2 \leq K \sum_{t=1}^T M_t'' = O(T),$$

where  $K < \infty$  and  $M_t'' \leq \max \left\{ M_t', E \|y_t\|^{2r} \right\}$  - see Corollary 16.10 in Davidson (1994, p. 255).

■

**Proof of Proposition 1.** The estimation error in  $\hat{\Sigma}$  can be represented as a perturbation of  $\Sigma$ , with  $\hat{\Sigma}_\tau = \Sigma + (\hat{\Sigma}_\tau - \Sigma)$ . Recall that in light of Theorem 1,  $\sup_{\lfloor T\tau \rfloor} \left\| \hat{\Sigma}_\tau - \Sigma \right\| = O_p(T^{-1/2})$ . The eigenvalue problem for the perturbed matrix is

$$\left[ \Sigma + (\hat{\Sigma}_\tau - \Sigma) \right] [x_i + (\hat{x}_{i,\tau} - x_i)] = \left[ \lambda_i + (\hat{\lambda}_{i,\tau} - \lambda_i) \right] [x_i + (\hat{x}_{i,\tau} - x_i)]. \quad (30)$$

After expanding the product, consider the terms  $(\hat{\Sigma}_\tau - \Sigma)(\hat{x}_{i,\tau} - x_i)$  and  $(\hat{\lambda}_{i,\tau} - \lambda_i)(\hat{x}_{i,\tau} - x_i)$ . It holds that  $\hat{\lambda}_{i,\tau} - \lambda_i = O_p(T^{-1/2})$  uniformly in  $\tau$ . This is because  $\Sigma$  is symmetric, and therefore Corollary 6.3.4 in Horn and Johnson (1999, p. 367) entails that  $|\hat{\lambda}_{i,\tau} - \lambda_i| \leq \left\| \hat{\Sigma}_\tau - \Sigma \right\|$ . Equation (1) yields the result. Also, it holds that  $\hat{x}_{i,\tau} - x_i = O_p(T^{-1/2})$  uniformly in  $\tau$ .

This follows from the sin  $\Theta$  Theorem in Davis and Kahan (1970, p. 10). Letting  $\delta_\Theta = \min_{1 \leq i \leq k, 1 \leq j \leq n-k} |\lambda_i - \lambda_j|$ , this entails that

$$|\sin(\hat{x}_{i,\tau}, x_i)| \leq \delta_\Theta^{-1} \left\| (\hat{\Sigma}_\tau - \Sigma) \Gamma \right\|,$$

where  $\Gamma = [x_1 | \dots | x_n]$  and  $\sin(\hat{x}_{i,\tau}, x_i)$  is the sine of the angle between the spaces spanned by  $\hat{x}_{i,\tau}$  and  $x_i$ . Now, after some manipulations,  $\sin(\hat{x}_{i,\tau}, x_i) = \sqrt{1 - \hat{x}_{i,\tau}' x_i} = 2^{-1/2} \sqrt{(\hat{x}_{i,\tau} - x_i)' (\hat{x}_{i,\tau} - x_i)}$ ; also, by the Continuous Mapping Theorem,  $\delta_\Theta^{-1} = \min_{1 \leq i \leq k, 1 \leq j \leq n-k} |\lambda_i - \lambda_j| + o_p(1)$ . These

results entail that the order of magnitude of  $\hat{x}_{i,\tau} - x_i$  is the same as that of  $\hat{\Sigma}_\tau - \Sigma$ . Thus,  $\left(\hat{\Sigma}_\tau - \Sigma\right) (\hat{x}_{i,\tau} - x_i)$  and  $\left(\hat{\lambda}_{i,\tau} - \lambda_i\right) (\hat{x}_{i,\tau} - x_i)$  are  $O_p(T^{-1})$  uniformly in  $\tau$ ; hence (30) can be written as

$$\Sigma (\hat{x}_{i,\tau} - x_i) + \left(\hat{\Sigma}_\tau - \Sigma\right) x_i = \lambda_i (\hat{x}_{i,\tau} - x_i) + \left(\hat{\lambda}_{i,\tau} - \lambda_i\right) x_i + O_p(T^{-1}). \quad (31)$$

Consider (9). Premultiplying (31) by  $x'_i$ , we obtain  $x'_i \Sigma (\hat{x}_{i,\tau} - x_i) + x'_i \left(\hat{\Sigma}_\tau - \Sigma\right) x_i = \lambda_i x'_i (\hat{x}_{i,\tau} - x_i) + \left(\hat{\lambda}_{i,\tau} - \lambda_i\right) x'_i x_i$ . Recalling that  $x'_i \Sigma = \lambda_i x'_i$ , and that  $x'_i x_i = 1$ , we have  $x'_i \left(\hat{\Sigma}_\tau - \Sigma\right) x_i = \hat{\lambda}_{i,\tau} - \lambda_i$ , which entails (9). In order to prove (10), note that the  $x_i$ s are a complete (and orthonormal) basis. This entails that there exists a unique set of constants  $\{\phi_{i,j,\tau}\}_{j=1}^n$  such that

$$\hat{x}_{i,\tau} - x_i = \sum_{j=1}^n \phi_{i,j,\tau} x_j. \quad (32)$$

We now discuss the constants  $\phi_{i,j,\tau}$ . Let us start by premultiplying (31) by any  $x'_k$  for  $i \neq k$ ; using the identity  $x'_i \Sigma = \lambda_i x'_i$ , we obtain

$$x'_k \Sigma (\hat{x}_{i,\tau} - x_i) + x'_k \left(\hat{\Sigma}_\tau - \Sigma\right) x_i = x'_k \lambda_i (\hat{x}_{i,\tau} - x_i) + O_p(T^{-1}),$$

so that, using the identity  $x'_i \Sigma = \lambda_i x'_i$

$$\lambda_k x'_k (\hat{x}_{i,\tau} - x_i) + x'_k \left(\hat{\Sigma}_\tau - \Sigma\right) x_i = \lambda_i x'_k (\hat{x}_{i,\tau} - x_i) + O_p(T^{-1});$$

using (32)

$$\lambda_k x'_k \sum_{j=1}^n \phi_{i,j,\tau} x_j + x'_k \left(\hat{\Sigma}_\tau - \Sigma\right) x_i = \lambda_i x'_k \sum_{j=1}^n \phi_{i,j,\tau} x_j + O_p(T^{-1}),$$

which reduces to

$$\lambda_k \phi_{i,k,\tau} + x'_k \left(\hat{\Sigma}_\tau - \Sigma\right) x_i = \lambda_i \phi_{i,k,\tau} + O_p(T^{-1}),$$

which yields

$$\phi_{i,k,\tau} = \frac{x'_k \left(\hat{\Sigma}_\tau - \Sigma\right) x_i}{\lambda_i - \lambda_k} + O_p(T^{-1}). \quad (33)$$

Also, note that

$$\phi_{i,i,\tau} = x'_i (\hat{x}_{i,\tau} - x_i) = x'_i \hat{x}_{i,\tau} - 1 = -\frac{1}{2} (\hat{x}_{i,\tau} - x_i)' (\hat{x}_{i,\tau} - x_i) = O_p(T^{-1}). \quad (34)$$

Thus, by (32) and (34)

$$\hat{x}_{i,\tau} - x_i = \sum_{j \neq i} \frac{x'_k (\hat{\Sigma}_\tau - \Sigma) x_i}{\lambda_i - \lambda_k} x_j + O_p(T^{-1}),$$

which proves (10). Using the results above, it holds that

$$\begin{aligned} \hat{\gamma}_{i,\tau} &= \hat{\lambda}_{i,\tau}^{1/2} \hat{x}_{i,\tau} = \lambda_i^{1/2} \left[ 1 + \frac{\hat{\lambda}_{i,\tau} - \lambda_i}{2\lambda_i} + O_p\left(\left\|\hat{\lambda}_{i,\tau} - \lambda_i\right\|^2\right) \right] [x_i + (\hat{x}_{i,\tau} - x_i)] \\ &= \lambda_i^{1/2} x_i + \lambda_i^{1/2} (\hat{x}_{i,\tau} - x_i) + \frac{\hat{\lambda}_{i,\tau} - \lambda_i}{2\lambda_i^{1/2}} x_i + O_p(T^{-1}), \end{aligned}$$

which, combining (9) and (10), yields (11).

We now turn to deriving the bias estimator for  $\hat{\lambda}_{i,\tau} - \lambda_i$ . Expanding (30) and premultiplying by  $x'_i$  we obtain

$$(\hat{\lambda}_{i,\tau} - \lambda_i) [1 + x'_i (\hat{x}_{i,\tau} - x_i)] = x'_i (\hat{\Sigma}_\tau - \Sigma) x_i + x'_i (\hat{\Sigma}_\tau - \Sigma) (\hat{x}_{i,\tau} - x_i);$$

applying Taylor's expansion

$$\begin{aligned} \hat{\lambda}_{i,\tau} - \lambda_i &= x'_i (\hat{\Sigma}_\tau - \Sigma) x_i - [x'_i (\hat{x}_{i,\tau} - x_i)] [x'_i (\hat{\Sigma}_\tau - \Sigma) x_i] + x'_i (\hat{\Sigma}_\tau - \Sigma) (\hat{x}_{i,\tau} - x_i) \\ &\quad - [x'_i (\hat{x}_{i,\tau} - x_i)] [x'_i (\hat{\Sigma}_\tau - \Sigma) (\hat{x}_{i,\tau} - x_i)] + O_p(T^{-5/2}) \\ &= x'_i (\hat{\Sigma}_\tau - \Sigma) x_i + I + II + III. \end{aligned}$$

Given that  $\hat{\Sigma}_\tau - \Sigma = O_p(T^{-1/2})$ , and using (34), we get that  $I = O_p(T^{-3/2})$  and  $III = O_p(T^{-2})$ . As far as  $II$  is concerned, note that using (10),  $x'_i \sum_{k \neq i} \frac{x'_k (\hat{\Sigma}_\tau - \Sigma) x_i}{\lambda_i - \lambda_k} x_k = 0$ . Also

$$\begin{aligned} II &= [x'_i \otimes (\hat{x}_{i,\tau} - x_i)] \text{vec}(\hat{\Sigma}_\tau - \Sigma) \\ &= \left[ x'_i \otimes \sum_{k \neq i} \frac{x'_k}{\lambda_i - \lambda_k} x'_k (\hat{\Sigma}_\tau - \Sigma) x_i \right] \text{vec}(\hat{\Sigma}_\tau - \Sigma) + O_p(T^{-3/2}) \\ &= \sum_{k \neq i} \left[ x'_i \otimes \frac{x'_k}{\lambda_i - \lambda_k} \right] [\text{vec}(\hat{\Sigma}_\tau - \Sigma)] [\text{vec}(\hat{\Sigma}_\tau - \Sigma)]' [x_k \otimes x_i] + O_p(T^{-3/2}) \\ &= II_a + O_p(T^{-3/2}). \end{aligned}$$

We have  $II_a = O_p(T^{-1})$ , and this is the dominating term in the bias..

The higher order terms of  $\hat{x}_{i,\tau} - x_i$  can be studied from (10) following similar passages.

Using (30), and premultiplying both sides by  $x'_k$ , we have

$$\begin{aligned} & \lambda_k x'_k (\hat{x}_{i,\tau} - x_i) + x'_k (\hat{\Sigma}_\tau - \Sigma) x_i + x'_k (\hat{\Sigma}_\tau - \Sigma) (\hat{x}_{i,\tau} - x_i) \\ = & \lambda_i x'_k (\hat{x}_{i,\tau} - x_i) + (\hat{\lambda}_{i,\tau} - \lambda_i) x'_k (\hat{x}_{i,\tau} - x_i), \end{aligned}$$

whence

$$\lambda_k \phi_{i,k,\tau} + x'_k (\hat{\Sigma}_\tau - \Sigma) x_i + x'_k (\hat{\Sigma}_\tau - \Sigma) (\hat{x}_{i,\tau} - x_i) = \lambda_i \phi_{i,k,\tau} + (\hat{\lambda}_{i,\tau} - \lambda_i) \phi_{i,k,\tau},$$

so that (33) becomes

$$\phi_{i,k,\tau} = \frac{x'_k (\hat{\Sigma}_\tau - \Sigma) \hat{x}_i}{\lambda_i - \lambda_k + (\hat{\lambda}_i - \lambda_i)};$$

note also that, by (34),  $x'_i (\hat{x}_{i,\tau} - x_i) = \frac{1}{2} (\hat{x}_{i,\tau} - x_i)' (\hat{x}_{i,\tau} - x_i)$ . Hence we can write

$$\begin{aligned} \hat{x}_{i,\tau} - x_i &= \sum_{k \neq i} \frac{x'_k (\hat{\Sigma}_\tau - \Sigma) \hat{x}_i}{\lambda_i - \lambda_k + (\hat{\lambda}_i - \lambda_i)} x_k + x'_i (\hat{x}_{i,\tau} - x_i) x_i \\ &= \sum_{k \neq i} \frac{x'_k (\hat{\Sigma}_\tau - \Sigma) \hat{x}_i}{\lambda_i - \lambda_k} \left[ 1 - \frac{\hat{\lambda}_i - \lambda_i}{\lambda_i - \lambda_k} \right] x_k + x_i x'_i (\hat{x}_{i,\tau} - x_i) + O_p(T^{-3/2}) \\ &= \sum_{k \neq i} \frac{x'_k (\hat{\Sigma}_\tau - \Sigma) x_i}{\lambda_i - \lambda_k} x_k - \sum_{k \neq i} \frac{\hat{\lambda}_i - \lambda_i}{\lambda_i - \lambda_k} \frac{x'_k (\hat{\Sigma}_\tau - \Sigma) x_i}{\lambda_i - \lambda_k} x_k \\ &\quad \sum_{k \neq i} \frac{x'_k (\hat{\Sigma}_\tau - \Sigma) (\hat{x}_i - x_i)}{\lambda_i - \lambda_k} x_k - x_i \frac{1}{2} (\hat{x}_{i,\tau} - x_i)' (\hat{x}_{i,\tau} - x_i) + O_p(T^{-3/2}) \\ &= \sum_{k \neq i} \frac{x'_k (\hat{\Sigma}_\tau - \Sigma) x_i}{\lambda_i - \lambda_k} x_k + I + II + III + O_p(T^{-3/2}), \end{aligned} \tag{35}$$

so that  $I + II + III$  can be estimated by

$$-\sum_{k \neq i} \frac{(\hat{x}'_k \otimes \hat{x}'_i) \hat{V}_\Sigma (\hat{x}_i \otimes \hat{x}_i)}{(\hat{\lambda}_i - \hat{\lambda}_k)^2} \hat{x}_k + \sum_{k \neq i} \sum_{h \neq i} \frac{(\hat{x}'_k \otimes \hat{x}'_h) \hat{V}_\Sigma (\hat{x}_h \otimes \hat{x}_i)}{(\hat{\lambda}_i - \hat{\lambda}_k) (\hat{\lambda}_i - \hat{\lambda}_h)} \hat{x}_k - \frac{1}{2} \sum_{k \neq i} \frac{(\hat{x}'_k \otimes \hat{x}'_i) \hat{V}_\Sigma (\hat{x}_k \otimes \hat{x}_i)}{(\hat{\lambda}_i - \hat{\lambda}_k)^2} \hat{x}_i, \tag{36}$$

or with a different estimator for  $\hat{V}_\Sigma$  (e.g.  $\tilde{V}_{\Sigma,\tau}$ ), or using partial sample estimates of the eigen-system. ■

**Proof of Theorem 3.** The proof of (15) follows from (1), Theorem 2 and the CMT. As far as (16) is concerned, the proof is based on the proof of Theorem A.4.1 in Csörgő and



Horváth (1997, p. 368-370). Here we summarize the main steps, using, as a leading example,  $\dot{\Lambda}(\tau) = \frac{1}{\sqrt{T\tau(1-\tau)}} \left[ \bar{S}(\tau)' \tilde{V}_{\Sigma,\tau}^{-1} \bar{S}(\tau) \right]^{1/2}$ , where  $\bar{S}(\tau) = S(\tau) - \frac{\lfloor T\tau \rfloor}{T} S(1)$ . We also define  $\ddot{\Lambda}(\tau) = \frac{1}{\sqrt{T\tau(1-\tau)}} \left[ \bar{S}(\tau)' V_{\Sigma}^{-1} \bar{S}(\tau) \right]^{1/2}$ ; further, letting  $B_{1i}(\tau)$  be a sequence of standard, independent Brownian bridges for  $i = 1, \dots, n^2$ , we define  $M(\tau) = \left[ \sum_{i=1}^{n^2} \frac{B_{1i}^2(\tau)}{\tau(1-\tau)} \right]^{1/2}$ . The Darling-Erdos Theorem (see e.g. Corollary A.3.1 in Csörgő and Horváth, 1997, p. 366) states that  $P \left[ a_T \sup_{\frac{1}{T} \leq \tau \leq 1 - \frac{1}{T}} M(\tau) \leq x + b_T \right] = e^{-2e^{-x}}$ , where the norming constants  $a_T$  and  $b_T$  are defined in the Theorem. In order to prove (16), it is enough to show that  $\left| \sup_{\frac{1}{T} \leq \tau \leq 1 - \frac{1}{T}} \dot{\Lambda}(\tau) - \sup_{\frac{1}{T} \leq \tau \leq 1 - \frac{1}{T}} M(\tau) \right| = o_p \left[ (\ln \ln T)^{-1/2} \right]$ . By virtue of Theorem 2, this entails that, as far as the estimated long-run covariance matrix is concerned, we need to have  $\sup_{1 \leq \lfloor T\tau \rfloor \leq T} \left\| \tilde{V}_{\Sigma,\tau} - V_{\Sigma} \right\| = o_p \left[ (\ln \ln T)^{-1/2} \right]$ . This holds, by virtue of equation (6), if both  $\frac{\sqrt{\ln \ln T}}{m} \rightarrow 0$  and  $m \ln T \sqrt{\frac{\ln \ln T}{T}} \rightarrow 0$ , whence the restrictions on  $m$  in the statement of the Theorem. Under such restrictions, it suffices to prove that

$$\left| \sup_{\frac{1}{T} \leq \tau \leq 1 - \frac{1}{T}} \ddot{\Lambda}(\tau) - \sup_{\frac{1}{T} \leq \tau \leq 1 - \frac{1}{T}} M(\tau) \right| = o_p \left( \frac{1}{\sqrt{\ln \ln T}} \right). \quad (37)$$

In order to show (37), note first that (2) yields the (weak) result

$$\sup_{\frac{1}{T} \leq \tau \leq 1 - \frac{1}{T}} \left| \ddot{\Lambda}(\tau) - M(\tau) \right| = o_p \left( \sqrt{\ln \ln T} \right). \quad (38)$$

Indeed, (2) entails

$$\sup_{u(T,\varepsilon) \leq \tau \leq \frac{1}{2}} \left[ \lfloor T\tau \rfloor \right]^{\delta} \left| \ddot{\Lambda}(\tau) - M(\tau) \right| = o_p(1), \quad (39)$$

$$\sup_{\frac{1}{2} \leq \tau \leq 1 - u(T,\varepsilon)} \left[ \lfloor T(1-\tau) \rfloor \right]^{\delta} \left| \ddot{\Lambda}(\tau) - M(\tau) \right| = o_p(1), \quad (40)$$

for all sequences  $u(T, \varepsilon)$  such that  $u(T, \varepsilon) \rightarrow 0$  and  $Tu(T, \varepsilon) \rightarrow \infty$  as  $T \rightarrow \infty$ ; here,  $\varepsilon$  is a number between 0 and 1. Choosing  $Tu(T, \varepsilon) = e^{(\ln T)^{\varepsilon}}$ , and applying Theorem A.3.1 in Csörgő and Horváth (1997, p. 363) it holds that

$$\begin{aligned} \frac{1}{\sqrt{2 \ln \ln T}} \sup_{\frac{1}{T} \leq \tau \leq u(T,\varepsilon)} M(\tau) &\xrightarrow{p} \sqrt{\varepsilon}, \\ \frac{1}{\sqrt{2 \ln \ln T}} \sup_{1 - u(T,\varepsilon) \leq \tau \leq 1 - \frac{1}{T}} M(\tau) &\xrightarrow{p} \sqrt{\varepsilon}. \end{aligned} \quad (41)$$

Hence, from (38)

$$\frac{1}{\sqrt{2 \ln \ln T}} \sup_{\frac{1}{T} \leq \tau \leq u(T, \varepsilon)} \ddot{\Lambda}(\tau) \xrightarrow{p} \sqrt{\varepsilon},$$

$$\frac{1}{\sqrt{2 \ln \ln T}} \sup_{1-u(T, \varepsilon) \leq \tau \leq 1-\frac{1}{T}} \ddot{\Lambda}(\tau) \xrightarrow{p} \sqrt{\varepsilon}.$$

Defining  $\xi(T)$  and  $\eta(T)$  as  $\sup_{1 \leq \lfloor T\tau \rfloor \leq T} M(\tau) = M[\xi(T)]$  and  $\sup_{1 \leq \lfloor T\tau \rfloor \leq T} \ddot{\Lambda}(\tau) = \ddot{\Lambda}[\eta(T)]$ , the relationships above entail  $P[u(T, \varepsilon) \leq \xi(T), \eta(T) \leq 1 - u(T, \varepsilon)] = 1$  as  $T \rightarrow \infty$ . Indeed, using (41) as an illustrative example, as  $T \rightarrow \infty$  and  $\varepsilon \rightarrow 0$

$$P \left[ a_T \sup_{\frac{1}{T} \leq \tau \leq u(T, \varepsilon)} M(\tau) - b_T \geq -K \right] = P[(\sqrt{\varepsilon} - 1) \ln \ln T \geq -K] = 0,$$

for some  $K > 0$ . Hence, (39) and (40) entail

$$\sup_{\frac{1}{T} \leq \tau \leq 1-\frac{1}{T}} \left| \ddot{\Lambda}(\tau) - M(\tau) \right| = o_p \left( e^{-\delta \ln^\varepsilon T} \right),$$

and since  $\left| \sup_{\frac{1}{T} \leq \tau \leq 1-\frac{1}{T}} \ddot{\Lambda}(\tau) - \sup_{\frac{1}{T} \leq \tau \leq 1-\frac{1}{T}} M(\tau) \right| \leq \sup_{\frac{1}{T} \leq \tau \leq 1-\frac{1}{T}} \left| \ddot{\Lambda}(\tau) - M(\tau) \right|$ , (37) follows in view of  $\sqrt{\ln \ln T} e^{-\delta \ln^\varepsilon T} \rightarrow 0$ . ■

**Proof of Theorem 4.** In order to prove (19), we show that, under  $H_a^{(T)}$

$$P \left[ \sup_{n \leq \lfloor T\tau \rfloor \leq T-n} \Lambda_T(\tau) > c_{\alpha, T} \right] = P[\Lambda_0 > c_{\alpha, T} - NC_T],$$

where  $\Lambda_0$  is the distribution of  $\sup_{n \leq \lfloor T\tau \rfloor \leq T-n} \Lambda_T(\tau)$  under the null of no change and  $NC_T$  is a non-centrality parameter. Tests based on  $\sup_{n \leq \lfloor T\tau \rfloor \leq T-n} \Lambda_T(\tau)$  are consistent as long as  $c_{\alpha, T} - NC_T \rightarrow -\infty$  as  $T \rightarrow \infty$ .

To begin with, note that

$$\Lambda_T^2(\tau) = \frac{T}{\lfloor T\tau \rfloor [T(1-\tau)]} \tilde{S}(\tau)' \hat{V}_\Sigma^{-1} \tilde{S}(\tau),$$

where we consider Assumption 2(i), and the full sample estimator only, for simplicity. Consider

$\tilde{S}(\tau)$ . Under  $H_a^{(T)}$

$$\begin{aligned}\sqrt{\frac{T}{[T\tau][T(1-\tau)]}}\tilde{S}(\tau) &= \sqrt{\frac{T}{[T\tau][T(1-\tau)]}}RD_{\lambda\gamma}\left[\sum_{t=1}^{[T\tau]}\bar{w}_t - \frac{[T\tau]}{T}\sum_{t=1}^T\bar{w}_t\right] \\ &\quad + RD_{\lambda\gamma}\Delta_T\sqrt{\frac{T}{[T\tau][T(1-\tau)]}}\left[\sum_{t=1}^{[T\tau]}I(t \leq k_{0,T}) - \frac{[T\tau]}{T}\sum_{t=1}^TI(t \leq k_{0,T})\right] \\ &= \tilde{S}_1(\tau) + \tilde{S}_2(\tau),\end{aligned}$$

where  $I(\cdot)$  is the indicator function.

We show that under  $H_a^{(T)}$ ,  $\|\hat{V}_\Sigma - V_\Sigma\|$  is bounded in probability. Consider  $\hat{\Sigma}$ ; it holds that  $vec(\hat{\Sigma}) = vec(\Sigma_t) + \left[\frac{T-k_{0,T}}{T} - I(t \geq k_{0,T})\right]\Delta_T + o_p(1)$ , where the  $o_p(1)$  term comes from a LLN. Therefore

$$\begin{aligned}\hat{V}_\Sigma &= \frac{1}{T}\sum_{t=1}^T\bar{w}_t\bar{w}_t' - \frac{1}{T}\sum_{t=1}^T\bar{w}_t\left[\frac{T-k_{0,T}}{T} - I(t \geq k_{0,T})\right]\Delta_T' \\ &\quad - \frac{1}{T}\sum_{t=1}^T\left[\frac{T-k_{0,T}}{T} - I(t \geq k_{0,T})\right]\Delta_T\bar{w}_t' \\ &\quad + \frac{1}{T}\sum_{t=1}^T\left[\frac{T-k_{0,T}}{T} - I(t \geq k_{0,T})\right]^2\Delta_T\Delta_T' \\ &= I + II + III + IV.\end{aligned}$$

The LLN entails that  $I \xrightarrow{p} V_\Sigma$ ;  $II$  and  $III$  have the same order of magnitude as each other.

Particularly, since  $\sum_{t=1}^T\bar{w}_t\left[\frac{T-k_{0,T}}{T} - I(t \geq k_{0,T})\right] = O_p(\sqrt{T})$ ,  $II = O_p\left(\frac{\|\Delta_T\|}{\sqrt{T}}\right)$ . Finally

$$\begin{aligned}&\frac{1}{T}\sum_{t=1}^T\left[\frac{T-k_{0,T}}{T} - I(t \geq k_{0,T})\right]^2 \\ &= \frac{1}{T}\sum_{t=1}^T\left(\frac{T-k_{0,T}}{T}\right)^2 - 2\left(\frac{T-k_{0,T}}{T}\right) + \frac{1}{T}\sum_{t=1}^TI(t \geq k_{0,T}) \\ &= \frac{k_{0,T}}{T}\frac{T-k_{0,T}}{T},\end{aligned}$$

thus,  $IV = O_p\left(\frac{k_{0,T}}{T}\frac{T-k_{0,T}}{T}\|\Delta_T\|^2\right)$ , which is  $O_p(1)$  under  $H_a^{(T)}$ . This entails that  $\|\hat{V}_\Sigma - V_\Sigma\| = O_p(1)$  under  $H_a^{(T)}$ . Applying Taylor's expansion, we can write  $\hat{V}_\Sigma^{-1} = V_\Sigma^{-1} + \dot{V}_\Sigma^{-1}(\hat{V}_\Sigma - V_\Sigma)\dot{V}_\Sigma^{-1}$ ,

for some invertible matrix  $\hat{V}_\Sigma$ . Further, consider the following intermediate result; since

$$\begin{aligned}\tilde{S}_2(\tau) &= RD_{\lambda\gamma}\Delta_T \sqrt{\frac{T}{\lfloor T\tau \rfloor \lfloor T(1-\tau) \rfloor}} \left[ \left( \frac{\lfloor T(1-\tau) \rfloor}{T} k_{0,T} \right) I(k_{0,T} < \lfloor T\tau \rfloor) \right. \\ &\quad \left. + \left( \frac{T - k_{0,T}}{T} \lfloor T\tau \rfloor \right) I(k_{0,T} \geq \lfloor T\tau \rfloor) \right],\end{aligned}$$

after some algebra we have

$$\sup_{1 \leq \lfloor T\tau \rfloor \leq T} \left\| \tilde{S}_2(\tau) \right\| = \|RD_{\lambda\gamma}\Delta_T\| \sqrt{k_{0,T} \left( \frac{T - k_{0,T}}{T} \right)}. \quad (42)$$

We now prove the theorem. It holds that

$$\begin{aligned}\Lambda_T(\tau) &= \tilde{S}_1(\tau)' V_\Sigma^{-1} \tilde{S}_1(\tau) + \tilde{S}_2(\tau)' \hat{V}_\Sigma^{-1} \tilde{S}_2(\tau) \\ &\quad + 2\tilde{S}_1(\tau)' \hat{V}_\Sigma^{-1} \tilde{S}_2(\tau) + \tilde{S}_1(\tau)' \hat{V}_\Sigma^{-1} (\hat{V}_\Sigma - V_\Sigma) \hat{V}_\Sigma^{-1} \tilde{S}_1(\tau) \\ &= I + II + III + IV.\end{aligned}$$

Consider  $I$ ; the sequence  $\bar{w}_t$  is zero mean, and it satisfies the assumptions of Theorem 1, and therefore  $\tilde{S}_1(\tau)$  follows the null distribution as  $T \rightarrow \infty$ . Further, given that  $\hat{V}_\Sigma$  is  $O_p(1)$  under  $H_a^{(T)}$ , term  $II$  has the same order as  $\sup_{n \leq \lfloor T\tau \rfloor \leq T-n} \left\| \tilde{S}_2(\tau) \right\|^2$ , which is  $O\left(k_{0,T} \frac{T - k_{0,T}}{T} \|RD_{\lambda\gamma}\Delta_T\|^2\right)$  in view of (42). Terms  $III$  and  $IV$  are of smaller order of magnitude than  $II$ : e.g. as far as  $III$  is concerned, it holds that  $E\left[\tilde{S}_1(\tau)' \hat{V}_\Sigma^{-1} \tilde{S}_2(\tau)\right] \leq \left(E\left\|\tilde{S}_1(\tau)\right\|^2\right)^{1/2} \left(E\left\|\tilde{S}_2(\tau)\right\|^2\right)^{1/2}$ , since  $\hat{V}_\Sigma^{-1}$  is  $O_p(1)$ ; thus,  $\sup_{n \leq \lfloor T\tau \rfloor \leq T-n} \tilde{S}_1(\tau)' \hat{V}_\Sigma^{-1} \tilde{S}_2(\tau) = O\left(\sqrt{\ln \ln T} \sqrt{k_{0,T} \frac{T - k_{0,T}}{T}} \|RD_{\lambda\gamma}\Delta_T\|\right)$ , which is smaller than  $II$ , as  $T \rightarrow \infty$ , when (18) holds. Therefore, under  $H_a^{(T)}$ ,  $P\left[\sup_{n \leq \lfloor T\tau \rfloor \leq T-n} \Lambda_T(\tau) > c_{\alpha,T}\right] = P[\Lambda_0 > c_{\alpha,T} - NC_T]$ , with

$$NC_T = \|RD_{\lambda\gamma}\Delta_T\| \sqrt{k_{0,T} \left( \frac{T - k_{0,T}}{T} \right)} + o\left[\|RD_{\lambda\gamma}\Delta_T\| \sqrt{k_{0,T} \left( \frac{T - k_{0,T}}{T} \right)}\right].$$

In view of  $c_{\alpha,T}$  being  $O\left(\sqrt{\ln \ln T}\right)$  and of (18), it holds that  $c_{\alpha,T} - NC_T \rightarrow -\infty$  as  $T \rightarrow \infty$ , whence (19) follows. ■

$n$	$T$		$(\rho, \theta)$				
			$(\rho, \theta)$	$(0, 0)$	$(0.5, 0)$	$(0, 0.5)$	$(0, -0.5)$
3	50	Kao et al.		0.013	0.015	0.009	0.010
		Aue et al.		0.006	0.002	0.003	0.005
	200	Kao et al.		0.041	0.048	0.041	0.040
		Aue et al.		0.029	0.023	0.025	0.029
	500	Kao et al.		0.044	0.058	0.043	0.045
		Aue et al.		0.034	0.027	0.029	0.033
10	50	Kao et al.		0.004	0.006	0.004	0.002
		Aue et al.		0.005	0.003	0.005	0.004
	200	Kao et al.		0.030	0.053	0.040	0.029
		Aue et al.		0.023	0.033	0.028	0.026
	500	Kao et al.		0.044	0.063	0.057	0.051
		Aue et al.		0.036	0.033	0.035	0.040

**Table 1. Empirical rejection frequencies for the null of no changes in the largest eigenvalue of  $\Sigma$ . Data are generated according to equation (22).**

$n$	$T$		$\Delta = \sqrt{\frac{\ln \ln(T)}{T^{2/3}}}$				
			$(\rho, \theta)$	$(0, 0)$	$(0.5, 0)$	$(0, 0.5)$	$(0, -0.5)$
3	50	Kao et al.		0.035	0.021	0.027	0.034
		Aue et al.		0.014	0.001	0.010	0.006
	200	Kao et al.		0.235	0.180	0.211	0.191
		Aue et al.		0.293	0.185	0.228	0.232
	500	Kao et al.		0.427	0.302	0.371	0.335
		Aue et al.		0.533	0.350	0.424	0.410
10	50	Kao et al.		0.032	0.026	0.033	0.021
		Aue et al.		0.020	0.012	0.021	0.015
	200	Kao et al.		0.356	0.247	0.273	0.309
		Aue et al.		0.440	0.248	0.328	0.336
	500	Kao et al.		0.528	0.364	0.449	0.434
		Aue et al.		0.661	0.430	0.536	0.537

**Table 2. Power of the test for the null of no changes in the largest eigenvalue of  $\Sigma$ . Data are generated according to equation (22) and under the alternative hypothesis specified in equation (25).**

$n$	$T$		$\Delta = \sqrt{\frac{\ln \ln(T)}{T^{1/2}}}$				
			$(\rho, \theta)$	$(0, 0)$	$(0.5, 0)$	$(0, 0.5)$	$(0, -0.5)$
3	50	Kao et al.		0.055	0.030	0.045	0.050
		Aue et al.		0.020	0.004	0.021	0.013
	200	Kao et al.		0.514	0.380	0.450	0.415
		Aue et al.		0.593	0.413	0.510	0.486
	500	Kao et al.		0.874	0.652	0.753	0.760
		Aue et al.		0.934	0.762	0.850	0.836
10	50	Kao et al.		0.056	0.040	0.053	0.039
		Aue et al.		0.037	0.017	0.030	0.027
	200	Kao et al.		0.711	0.465	0.566	0.598
		Aue et al.		0.796	0.524	0.648	0.674
	500	Kao et al.		0.954	0.768	0.889	0.866
		Aue et al.		0.979	0.856	0.943	0.920

**Table 3.** Power of the test for the null of no changes in the largest eigenvalue of  $\Sigma$ . Data are generated according to equation (22) and under the alternative hypothesis specified in equation (26).

$n$	$T$		$k = Th_{nT} + 1$ $k = \frac{1}{2} [\ln(T)]^2$ $k = \frac{1}{2} [\ln(T)]^{5/2}$ $k = 3\sqrt{T}$			
3	50	Kao et al.	0.071	0.017	0.017	0.059
		Aue et al.	0.054	0.037	0.037	0.162
	200	Kao et al.	0.485	0.488	0.304	0.609
		Aue et al.	0.224	0.171	0.249	0.614
	500	Kao et al.	0.886	0.834	0.904	0.997
		Aue et al.	0.156	0.157	0.306	0.847
10	50	Kao et al.	0.012	0.010	0.010	0.061
		Aue et al.	0.071	0.057	0.057	0.234
	200	Kao et al.	0.516	0.561	0.375	0.709
		Aue et al.	0.262	0.202	0.292	0.718
	500	Kao et al.	0.915	0.936	0.942	1.000
		Aue et al.	0.179	0.195	0.356	0.905

**Table 4.** Power of the test for the null of no changes in the largest eigenvalue of  $\Sigma$ . Data are generated as i.i.d., under the alternative specified in equation (27).

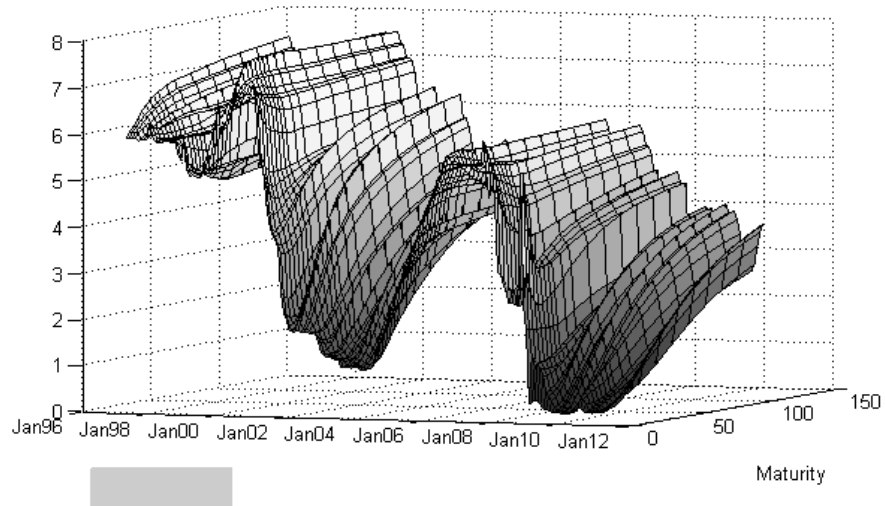
$i$	Panel A $H_0 : \Sigma_{ii} \text{ constant}$			Panel B $H_0 : \lambda_i \text{ constant}$			Panel C $H_0 : \gamma_i \text{ constant}$	
	<i>monthly</i>	<i>weekly</i>		<i>monthly</i>	<i>weekly</i>		<i>monthly</i>	<i>weekly</i>
1m	2.6989	2.8421						
3m	2.7656	3.5461	$\lambda_1$	1.6921	3.5798** [1st week, 12/2007]	$x_1$	3.9142	6.957** [3rd week, 03/2008]
6m	2.7394	3.0854						
9m	2.3924	2.1531	$\lambda_2$	2.5513	2.7488	$x_2$	4.3898	7.098** [3rd week, 04/2008]
12m	1.5350	2.9454						
15m	1.4991	2.6190	$\lambda_3$	3.4328** [01/2008]	2.7726	$x_3$	4.2340	7.261*** [2nd week, 03/2008]
18m	1.6467	2.4979						
21m	1.8065	2.6907						
24m	1.9827	2.9462						
30m	2.0718	3.1947						
3y	2.0815	3.4064						
4y	1.9314	3.7837						
5y	1.8964	3.8836* [1st week, 12/2007]						
6y	1.8369	4.0432** [1st week, 12/2007]						
7y	1.7677	4.0488** [1st week, 12/2007]						
8y	1.9601	4.1446** [1st week, 12/2007]						
9y	2.1046	4.2285** [last week, 08/2008]						
10y	2.1967	4.3417** [last week, 08/2008]						

**Table 5.** Tests for changes in the variances of the term structure; in the volatilities of each principal component; and in the eigenvalue-normed eigenvectors. Rejection at 10%, 5% and 1% levels are denoted with \*, \*\* and \*\*\* respectively. Where present, numbers in square brackets are the estimated breakdates, defined as  $T \times \arg \max \Lambda_T(\tau)$ .

<i>monthly data</i>			<i>weekly data</i>		
	<i>1st subsample</i>	<i>2nd subsample</i>		<i>1st subsample</i>	<i>2nd subsample</i>
$\lambda_1$	0.790	0.729	$\lambda_1$	0.737	0.784
$\lambda_2$	0.163	0.214	$\lambda_2$	0.163	0.138
$\lambda_3$	0.029	0.047	$\lambda_3$	0.056	0.056

**Table 6.** Proportion of the total variance explained by principal components ( $\lambda_1$ ,  $\lambda_2$  and  $\lambda_3$  refer to the level, slope and curvature respectively) for each subsample. The samples are split based on the results in Table 2. When considering monthly data, the sample was split at January 2008; when using weekly data, at the first week of December 2007.





**Figure 1. Term structure of the US interest rates. Maturities correspond to 1m, 3m, 6m, 9m, 12m, 15m, 18m, 21m, 24m, 30m, 3y, 4y, 5y, 6y, 7y, 8y, 9y, 10y over the period April 1997-November 2010.**

# Supplement to: Testing for instability in covariance structures

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This supplement contains some further material for the paper “Testing for instability in covariance structures”.

In particular, in this note we consider: the extension of results derived in the main paper to residuals (Section 1); a related empirical exercise, where we study the stability of the covariance matrix of the error term of a VAR model for leading currency exchange rates (Section 2); a data appendix, where we describe the data employed in both the application in the main paper and the one in Section 2 (Section 3); the statements and proofs of several lemmas, either used in the main paper or in this supplement, and of the results derived in here (Section 4); all the tables that refer to simulations in the main paper, the critical values employed for the empirical applications in this supplement and in the main paper, and the empirical exercise in this supplement (Section 5).

## 1 Applications to the residuals of a multivariate regression

In this section, we extend the results developed in the main paper by proposing a test for the time stability of the covariance matrix (and its eigensystem) of the error term in a multivariate regression, such as a VAR.

Consider the following multivariate regression model

$$y_t = \beta x_t + \varepsilon_t, \tag{1}$$

where  $t = 1, \dots, T$  and  $y_t$  and  $\varepsilon_t$  are  $n \times 1$  vectors,  $x_t$  is of dimension  $q \times 1$  and the matrix of

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regressors  $\beta$  has dimensions  $n \times q$ . Of course,  $n$  can be equal to 1, thereby recovering the usual univariate regression; alternatively, the regressors  $x_t$  can be the lags of  $y_t$ , thus having a VAR. Bai (2000) considers a test for structural changes in a VAR, extending it to test for changes in the covariance matrix of the error term. The theory developed in this section also extends naturally to the case of linear or polynomial trends among the regressors - see also Aue et al. (2012).

Let  $\Sigma_\varepsilon = E(\varepsilon_t \varepsilon_t')$  be the covariance matrix of the error term  $\varepsilon_t$ ; using the same notation as above, we denote its eigenvalues and eigenvectors as  $\lambda_i^\varepsilon$  and  $x_i^\varepsilon$  respectively, satisfying the relationship  $\Sigma_\varepsilon x_i^\varepsilon = \lambda_i^\varepsilon x_i^\varepsilon$  for  $i = 1, \dots, n$ .

Tests for the time stability of  $\Sigma_\varepsilon$  and of its eigensystem are based on the residuals  $\hat{\varepsilon}_t = y_t - \hat{\beta}x_t$ , where  $\hat{\beta} = \left[ \sum_{t=1}^T y_t x_t' \right] \left[ \sum_{t=1}^T x_t x_t' \right]^{-1}$ . Hence, the full sample estimator of  $\Sigma_\varepsilon$  is defined as  $\hat{\Sigma}_\varepsilon = T^{-1} \sum_{t=1}^T \hat{\varepsilon}_t \hat{\varepsilon}_t'$ . As above, for a given point in time  $[T\tau]$  with  $\tau \in [0, 1]$ , we use the subscript  $\tau$  to denote quantities calculated using the subsamples  $t = 1, \dots, [T\tau]$ . In particular, we extensively use the sequence of partial sample estimators  $\hat{\Sigma}_{\varepsilon, \tau} = (T\tau)^{-1} \sum_{t=1}^{[T\tau]} \hat{\varepsilon}_t \hat{\varepsilon}_t'$ . We also employ the following definitions:  $w_t^\varepsilon = \text{vec}(\varepsilon_t \varepsilon_t')$  and  $\bar{w}_t^\varepsilon = \text{vec}(\varepsilon_t \varepsilon_t' - \Sigma_\varepsilon)$ .

We need the following Assumptions, which are closely related to Assumptions 1-3 in the main paper.

**Assumption R1.** (i) Assumptions 1(i)-1(ii) hold for  $x_t$ ; (ii) Assumptions 1(i)-1(ii) hold for  $\varepsilon_t$ ; (iii) let  $w_t^{x\varepsilon} = \text{vec}(\varepsilon_t x_t')$ ; it holds that: (a)  $E(w_t^{x\varepsilon}) = 0$  for all  $t$ ; (b) defining  $V_T^{x\varepsilon} = T^{-1} E \left[ \left( \sum_{t=1}^T w_t^{x\varepsilon} \right) \left( \sum_{t=1}^T w_t^{x\varepsilon} \right)' \right]$ ,  $V_T^{x\varepsilon}$  is positive definite uniformly in  $T$ , and as  $T \rightarrow \infty$ ,  $V_T^{x\varepsilon} \rightarrow V^{x\varepsilon}$  with  $\|V^{x\varepsilon}\| < \infty$ ; (c) letting  $w_{it}^{x\varepsilon}$  be the  $i$ -th element of  $w_t^{x\varepsilon}$  and defining  $S_{iT, m}^{x\varepsilon} \equiv \sum_{t=m+1}^{m+T} w_{it}^{x\varepsilon}$ , there exists a positive definite matrix  $\bar{\Omega}^{x\varepsilon} = \{\varpi_{ij}^{x\varepsilon}\}$  such that  $T^{-1} \left| E \left[ S_{iT, m}^{x\varepsilon} S_{jT, m}^{x\varepsilon} \right] - \varpi_{ij}^{x\varepsilon} \right| \leq MT^{-\psi}$ , for all  $i$  and  $j$  and uniformly in  $m$ , with  $M$  a constant and  $\psi > 0$ ; (iv) (a) letting  $V_T^\varepsilon = T^{-1} E \left[ \left( \sum_{t=1}^T \bar{w}_t^\varepsilon \right) \left( \sum_{t=1}^T \bar{w}_t^\varepsilon \right)' \right]$ ,  $V_T^\varepsilon$  is positive definite uniformly in  $T$ , and as  $T \rightarrow \infty$ ,  $V_T^\varepsilon \rightarrow V^\varepsilon$  with  $\|V^\varepsilon\| < \infty$ ; (b) letting  $\bar{w}_{it}^\varepsilon$  be the  $i$ -th element of  $\bar{w}_t^\varepsilon$  and defining  $S_{iT, m}^\varepsilon \equiv \sum_{t=m+1}^{m+T} \bar{w}_{it}^\varepsilon$ , there exists a positive definite matrix  $\bar{\Omega}^\varepsilon = \{\varpi_{ij}^\varepsilon\}$  such that  $T^{-1} \left| E \left[ S_{iT, m}^\varepsilon S_{jT, m}^\varepsilon \right] - \varpi_{ij}^\varepsilon \right| \leq M'T^{-\psi'}$ , for all  $i$  and  $j$  and uniformly in  $m$ , with  $\psi' > 0$ .

**Assumption R2.** Let  $\Psi_l^\varepsilon = E(\bar{w}_t^\varepsilon \bar{w}_{t-l}^{\varepsilon'})$ ; it holds that (i)  $\sum_{l=0}^\infty l^s \|\Psi_l^\varepsilon\| < \infty$  for some  $s \geq 1$ ; (ii)  $\sup_t E \|\varepsilon_t\|^{4r} < \infty$  for some  $r > 2$ ; (iii) letting  $\Omega_T^{l\varepsilon} = T^{-1} E \left\{ \sum_{t=1}^T \text{vec} [\bar{w}_t^\varepsilon \bar{w}_{t-l}^{\varepsilon'} - E(\bar{w}_t^\varepsilon \bar{w}_{t-l}^{\varepsilon'})] \text{vec} [\bar{w}_t^\varepsilon \bar{w}_{t-l}^{\varepsilon'} - E(\bar{w}_t^\varepsilon \bar{w}_{t-l}^{\varepsilon'})]' \right\}$ ,  $\Omega_T^{l\varepsilon}$  is positive definite uniformly in  $T$ , and  $\Omega_T^{l\varepsilon} \rightarrow \Omega^{l\varepsilon}$  with  $\|\Omega^{l\varepsilon}\| < \infty$  for every  $l$ .

**Assumption R3.** Assumption 3 in the main paper holds for the matrix  $\Sigma_\varepsilon$ .

Let  $\hat{w}_t^\varepsilon = \text{vec}(\hat{\varepsilon}_t \hat{\varepsilon}_t')$ . In order to estimate the long run covariance matrix  $V^\varepsilon$  defined in Assumption R1(iv)(b), we propose the weighted sum-of-covariance estimator  $\tilde{V}_\tau^\varepsilon = \left( \hat{\Psi}_{0, [T\tau]}^\varepsilon + \hat{\Psi}_{0, 1-[T\tau]}^\varepsilon \right)$

$+ \sum_{l=1}^m (1 - \frac{l}{m}) \left[ \left( \hat{\Psi}_{l, \lfloor T\tau \rfloor}^\varepsilon + \hat{\Psi}_{l, \lfloor T\tau \rfloor}^{\varepsilon'} \right) + \left( \hat{\Psi}_{l, 1-\lfloor T\tau \rfloor}^\varepsilon + \hat{\Psi}_{l, 1-\lfloor T\tau \rfloor}^{\varepsilon'} \right) \right]$ , where  $\hat{\Psi}_{l, \lfloor T\tau \rfloor}^\varepsilon = \frac{1}{T - (l+1)} \sum_{t=l+1}^{\lfloor T\tau \rfloor} \left[ \hat{w}_t^\varepsilon - \text{vec} \left( \hat{\Sigma}_\varepsilon \right) \right] \left[ \hat{w}_{t-l}^\varepsilon - \text{vec} \left( \hat{\Sigma}_\varepsilon \right) \right]'$ , and similarly  $\hat{\Psi}_{l, 1-\lfloor T\tau \rfloor}^\varepsilon$ .

In order to set up the test statistic, consider the following notation. When estimating the eigensystem of  $\Sigma_\varepsilon$ , we define the estimators of  $(\lambda_i^\varepsilon, x_i^\varepsilon)$  as the solution to  $\hat{\Sigma}_\varepsilon \hat{x}_i^\varepsilon = \hat{\lambda}_i^\varepsilon \hat{x}_i^\varepsilon$ , where  $\hat{\lambda}_i^\varepsilon$  and  $\hat{x}_i^\varepsilon$  denote the estimates of  $\lambda_i^\varepsilon$  and  $x_i^\varepsilon$  respectively; the Pearson-Hotelling estimator of the eigenvectors is defined as  $\hat{\gamma}_i^\varepsilon = \sqrt{\hat{\lambda}_i^\varepsilon} \hat{x}_i^\varepsilon$ . Similarly, the partial sample estimators of the eigenvalues and eigenvectors are the solutions to  $\hat{\Sigma}_{\varepsilon, \tau} \hat{x}_{i, \tau}^\varepsilon = \hat{\lambda}_{i, \tau}^\varepsilon \hat{x}_{i, \tau}^\varepsilon$ , and we define  $\hat{\gamma}_{i, \tau}^\varepsilon = \sqrt{\hat{\lambda}_{i, \tau}^\varepsilon} \hat{x}_{i, \tau}^\varepsilon$ . Thence, let  $\lambda^\varepsilon \equiv [\lambda_1^\varepsilon, \dots, \lambda_n^\varepsilon]'$  be the  $n$ -dimensional vector containing the eigenvalues sorted in descending order;  $\Gamma^\varepsilon \equiv [\gamma_1^\varepsilon, \dots, \gamma_n^\varepsilon]$ ; and let  $D_{\lambda\gamma}^\varepsilon \equiv [x_1^\varepsilon \otimes x_1^\varepsilon, \dots, x_n^\varepsilon \otimes x_n^\varepsilon, v_{\gamma, 1}^{\varepsilon'}, \dots, v_{\gamma, n}^{\varepsilon'}]'$ , with  $v_{\gamma, i}^\varepsilon = \frac{1}{2} \frac{x_i^\varepsilon}{\sqrt{\lambda_i^\varepsilon}} (x_i^{\varepsilon'} \otimes x_i^{\varepsilon'}) + \sum_{k \neq i} \frac{\sqrt{\lambda_i^\varepsilon x_k^\varepsilon}}{\lambda_i^\varepsilon - \lambda_k^\varepsilon} (x_i^{\varepsilon'} \otimes x_k^{\varepsilon'})$ . Estimation of  $D_{\lambda\gamma}^\varepsilon$  is based on  $\hat{\lambda}_i^\varepsilon$  and  $\hat{x}_i^\varepsilon$ .

Based on the notation spelt out above, we define the test statistic for the time stability of  $\Sigma_\varepsilon$  and its eigensystem. Let

$$\tilde{S}^\varepsilon(\tau) = R \times D_{\lambda\gamma}^\varepsilon \times \left[ S^\varepsilon(\tau) - \frac{\lfloor T\tau \rfloor}{T} S^\varepsilon(1) \right],$$

where  $S^\varepsilon(\tau)$  is the CUSUM process defined as  $\sum_{t=1}^{\lfloor T\tau \rfloor} \text{vec}(\hat{\varepsilon}_t \hat{\varepsilon}_t')$ ,  $\tilde{S}^\varepsilon(\tau) = 0$  for  $\tau \leq \frac{1}{T}$  or  $\geq 1 - \frac{1}{T}$ , and  $R$  is defined in equation (13) in the main paper. The test statistic is defined as

$$\Lambda_T^\varepsilon(\tau) = \sqrt{\frac{T}{\lfloor T\tau \rfloor \times \lfloor T(1-\tau) \rfloor}} \times \left[ \tilde{S}^\varepsilon(\tau)' \left( \tilde{V}_{\varepsilon, \tau} \right)^{-1} \tilde{S}^\varepsilon(\tau) \right]^{1/2},$$

with  $\tilde{V}_{\varepsilon, \tau} = R D_{\lambda\gamma}^\varepsilon \tilde{V}_\tau^\varepsilon D_{\lambda\gamma}^{\varepsilon'} R'$ .

Let  $\Lambda_T^{\varepsilon*} \equiv \sup_{nq \leq \lfloor T\tau \rfloor \leq T-nq} \Lambda_T^\varepsilon(\tau)$ . The following theorem summarizes the asymptotics of  $\Lambda_T^{\varepsilon*}$  under the null and under the alternative. Similarly to Theorem 4 in the main paper, we define the critical value  $c_{\alpha, T}$  to be such that, under  $H_0$ ,  $P[\Lambda_T^{\varepsilon*} \leq c_{\alpha, T}] = 1 - \alpha$  for some  $\alpha \in [0, 1]$ .

**Theorem 1** *Let Assumptions R1-R3 hold, and let  $p$  be defined as in Theorem 3 in the main paper. Under the null, as  $(m, T) \rightarrow \infty$  with  $\frac{\sqrt{\ln \ln T}}{m} + \frac{m}{\sqrt{T}} \ln T \ln \ln T \rightarrow 0$ , it holds that  $P\{a_T \Lambda_T^{\varepsilon*} \leq x + b_T\} \rightarrow e^{-2e^{-x}}$ , where  $a_T = \sqrt{2 \ln \ln T}$  and  $b_T = 2 \ln \ln T + \frac{p}{2} \ln \ln \ln T - \ln \Gamma\left(\frac{p}{2}\right)$ , with  $\Gamma(\cdot)$  the Gamma function. Under the alternative hypothesis defined in equation (18) in the main paper for  $\Sigma_\varepsilon$ , if equation (19) in the main paper holds, then  $P[\Lambda_T^{\varepsilon*} > c_{\alpha, T}] = 1$ .*

## Remarks

T5.1 The main feature of Theorem 1 is that the same theory applies to residual-based tests as to observable data. This is due to the fact that, under our assumptions, an almost sure rate of convergence can be derived for the OLS estimator  $\hat{\beta}$ , with  $\hat{\beta} - \beta = O_{a.s.} \left( \sqrt{\frac{\ln \ln T}{T}} \right)$ . By virtue of this result, a SIP and a SLLN can be shown for the partial sums of the squared

residuals, so that the asymptotics of tests based on  $\Lambda_T^\varepsilon(\tau)$  can be derived in the same way as for tests based on  $\Lambda_T(\tau)$ .

T5.2 Theorem 1 can be readily extended to accommodate for the presence of linear or polynomial trends among the  $x_t$ s; the only thing that changes is the rate of convergence of  $\hat{\beta}$ , which in presence of linear trends is actually faster than  $O_{a.s.}\left(\sqrt{\frac{\ln \ln T}{T}}\right)$ . A consequence of Theorem 1 is that the test developed in the main paper can be applied to data with nonzero mean, trends or even time dummies: it suffices to filter out such deterministics, and then apply the test to the resulting residuals.

## 2 A VAR model for exchange rates: the stability of the covariance matrix of the error term

In this section, we apply the test developed in Section 1 to a VAR, checking whether the covariance matrix of the error term is stable over time. Inspired by Carriero et al. (2009), we consider a VAR consisting of 4 exchange rates vis-a-vis the US Dollar. Specifically, we consider monthly averages of the following currencies: Euro, British Pound, Yen, Canadian Dollar; data are collected from January 1999 to April 2013, which entails a sample size (modulo the first entry) of  $T = 172$ .

Two tests are carried out in parallel. We verify whether the covariance matrix of the error term as a whole is subject to any changes. Furthermore, we also test for the stability of each of the eigenvalues of the covariance matrix. Since we find very weak evidence of breaks (if any) in the whole matrix and in its spectrum, we do not carry out any analysis on the eigenvectors: the key finding in this section is that, despite a major event such as the 2007-2009 recession, the covariance matrix of the VAR model of exchange rates does not change over time.

Carriero et al. (2009) propose a Bayesian approach to model the exchange rates directly, imposing a prior on the VAR coefficients in order to shrink them towards a unit root representation. Preliminary analysis shows that, as expected, our series have a unit root. We therefore fit a VAR to the returns of the exchange rates. We report descriptive statistics for the returns, and a plot of the original series, in Table E1 and Figure A respectively.

**[Insert Table E1 and Figure A somewhere here]**

As the evidence in Carriero et al. (2009) shows, applying VAR models to exchange rates could yield a good level of forecasting accuracy, also in light of the fact that exchange rates have a tendency to co-move. However, Castle et al. (2010) find that when the covariance matrix of the error term in a regression changes, this could have pernicious effects on the forecasting ability of the model; thus, applying our test as discussed in Section 1 to the covariance matrix

of the error term (and to its eigenvalues) could shed some light as to the forecasting ability of the VAR. Estimation results and mis-specification tests are reported in Table E2: the best specification was found to be a VAR of order 5, using lags 1, 2 and 5. As can be seen from the Table, the model is correctly specified, save for the lack of normality in the error term; in view of Assumption R1, this is not an issue when applying our test. We also note that, as can be expected as regards the residuals of a correctly specified model, the error term is found to be serially uncorrelated, and we therefore estimate the long run variance without adjusting for serial correlation.

**[Insert Table E2 somewhere here]**

The test is applied, preliminarily, to the (lower triangular part of the) whole covariance matrix: the results in the Monte Carlo exercise suggest that the dimension of the matrix ( $n = 4$  in our context) affords the test to have the correct size and power versus the alternative of one or more breaks. Given that we consider only the lower triangular part of the matrix, this entails applying our test with  $p = 10$  constraints under the null, as opposed to  $p = 16$  constraints which is the case in the Monte Carlo exercise.

We also verify the stability of each of the four eigenvalues of the matrix. As in Section the Monte Carlo exercise, this procedure is implemented by controlling for the family-wise rejection rate though a Bonferroni correction. Critical values for individual tests of levels  $\alpha_I = 1\%$ ,  $5\%$  and  $10\%$ , and for procedure level  $\alpha_P = 1\%$ ,  $5\%$  and  $10\%$ , with the Bonferroni correction  $\alpha_I = \frac{\alpha_P}{4}$  are in Table D.

Results are reported in Table E3.

**[Insert Table E3 somewhere here]**

As can be seen from the first panel of the Table, the matrix, as a whole, is stable over time: as pointed out above, despite the presence of a major event such as the 2007-2009 recession, the covariance structure of the error term of the VAR does not change. This could be interpreted in light of the stylised fact that exchange rates tend to co-move (see the discussion in Carriero et al., 2009). This finding is also reinforced by the second panel of Table E3, where we find that the first three largest eigenvalues (denoted as  $\lambda_1 - \lambda_3$ ) do not change over time. As far as the smallest one is concerned ( $\lambda_4$ ), some evidence of a change is indeed found. However, the test rejects the null of a stable eigenvalue only when applied to each eigenvalue separately. Conversely, when adjusting for the family-wise error, no break is found at 5% level (the null of no break is rejected at 10% level for  $\lambda_4$ ); as mentioned in Section 4 in the main paper, this is a consequence of the Bonferroni correction yielding relatively conservative results. However, this suggests very weak evidence of a break in the smallest eigenvalue; the estimated breakdate is found to be around June 2008, which has the natural interpretation of being a consequence of the latest recession.

### 3 Data description

The data for the first exercise (time stability of the principal components of the term structure) have been obtained from Datastream (source: ICAP), and they refer to the zero-coupon yield curve obtained from Treasury bond yields. We have employed the following tickers: US00Y01; US00Y03; US00Y06; US00Y09; US01Y00; US01Y03; US01Y06; US01Y09; US02Y00; US02Y06; US03Y00; US04Y00; US05Y00; US06Y00; US07Y00; US08Y00; US09Y00; US10Y00.

The data for the second exercise (time stability of the covariance matrix of a VAR fitted to major currencies) has been obtained from the FRED archive of the Federal Reserve Bank of St Louis, <https://research.stlouisfed.org/fred2/release?rid=15>.

### 4 Useful lemmas and proofs

In this section, we report the statements and proofs of Lemmas 1-4, frequently cited in the paper. We also report the necessary Lemmas for the results in Section 5, and the proof of Theorem 1.

We note that many of the result in this section are extensions of existing results in the literature - indeed, most of the results derived here are in Davidson (1994). Most results in Davidson (1994), however, are reported for the  $L_1$ -norm, whilst, when proving the SIP, we need essentially the same results but for the  $L_{1+\epsilon}$ -norm.

**Lemma 1** *Let  $B$  and  $\rho$  be non-negative random variables, with  $|B|_{q^*/(q-1)} < \infty$  and  $|\rho|_{q^*} < \infty$ , where  $q^* = q(1 + \frac{\delta}{2})$ ,  $q \geq 1$  and  $\delta > 0$ . Assume  $|B\rho|_r < \infty$  for  $r > 2 + \delta$ . Then*

$$|B\rho|_{2+\delta} \leq \left[ |\rho|_{q^*} |B|_{q^*/(q-1)} \right]^{\frac{r-(2+\delta)}{2r-(2+\delta)}} \times \left[ |B\rho|_r^{\frac{r}{2+\delta} - \frac{r(2+\delta-r)}{(2+\delta)-r}} + |B\rho|_r^{\frac{(2+\delta)r}{2r-(2+\delta)}} \right]. \quad (2)$$

**Proof.** The proof is fairly similar to Gallant and White (1988; see Davidson, 1994, p. 271), who prove it for the  $L_2$ -norm. Let  $C = \left[ |\rho|_{q^*} |B|_{q^*/(q-1)} |B\rho|_r^{-r} \right]^{\frac{2+\delta}{2+\delta-2r}}$ , and define  $B_1 = I_{\{B\rho \leq C\}} B$ . By construction,  $|B\rho|_{2+\delta} \leq |B_1\rho|_{2+\delta} + |(B - B_1)\rho|_{2+\delta}$ . We have

$$\begin{aligned} |B_1\rho|_{2+\delta} &= \left[ \int_{B\rho \leq C} (B\rho)^{2+\delta} dP \right]^{\frac{1}{2+\delta}} \\ &\leq C^{1/2} \left[ \int_{B\rho \leq C} (B\rho)^{\frac{2+\delta}{2}} dP \right]^{\frac{2}{2+\delta} \frac{1}{2}} \\ &\leq C^{1/2} \left[ |\rho|_{q^*} |B|_{q^*/(q-1)} \right]^{1/2}, \end{aligned}$$

where the last passage follows from Holder's inequality. Also, since  $r > 2$  and, in this case,

$$(B\rho)^r > C^r,$$

$$\begin{aligned} |(B - B_1)\rho|_{2+\delta} &= \left[ \int_{B\rho > C} (B\rho)^{2+\delta} dP \right]^{\frac{1}{2+\delta}} \\ &\leq C^{1-\frac{r}{2+\delta}} \left[ \int_{B\rho > C} (B\rho)^r dP \right]^{\frac{r}{2+\delta} \frac{1}{r}} \\ &\leq C^{1-\frac{r}{2+\delta}} |B\rho|_r^{\frac{r}{2+\delta}}. \end{aligned}$$

Substituting for  $C$  gives (2). ■

**Lemma 2** Under Assumption 1,  $w_t$  is  $L_{2+\epsilon}$ -NED of size  $\alpha' > \frac{1}{2}$  on  $\{v_t\}_{t=-\infty}^{+\infty}$ .

**Proof.** The proof follows similar passages as in Example 17.17 in Davidson (1994, p. 273); for simplicity, assume  $n = 1$ . Let  $x_t^a = w_t$  and  $x_t^b = E[w_t | w_{t-m}, \dots, w_{t+m}]$ ; and define, in a similar fashion,  $y_t^a = y_t$  and  $y_t^b = E[y_t | y_{t-m}, \dots, y_{t+m}]$ . Then

$$\begin{aligned} |x_t^a - x_t^b|_{2+\epsilon} &\leq \left( |y_t^a| + |y_t^b| \right) \left( |y_t^a - y_t^b| \right)_{2+\epsilon} \\ &= |B(y_t^a, y_t^b) \rho(y_t^a, y_t^b)|_{2+\epsilon} = |B\rho|_{2+\epsilon}. \end{aligned}$$

We begin with considering some preliminary results. First, we show that  $|B\rho|_r < \infty$ , uniformly in  $m$ . We have  $|B\rho|_r < |B|_{2r} |\rho|_{2r}$ , with  $|B|_{2r} \leq |y_t^a|_{2r} + |y_t^b|_{2r} \leq 2|y_t^a|_{2r}$ , having used Minkowski's inequality and conditional modulus inequalities; this is bounded due to Assumption 1(i), and the bound holds uniformly in  $m$ . Also,  $|\rho|_{2r} \leq 2|y_t^a|_{2r}$  as a consequence of equation (17.4) in Davidson (1994, p. 262), again uniformly in  $m$ . Further, it holds that  $|\rho|_{q^*} < \infty$  for  $q^* \leq 2r$ , and thus for  $q < 2r$ ; also,  $|B|_{q^*/(q-1)} < \infty$  if  $q^* \geq \frac{4}{3}$ , i.e.  $q > \frac{4}{3}$ . Finally, note that for  $q^* \leq 2$ ,  $|\rho|_{q^*} \leq |\rho|_{2+\epsilon}$ ; by Assumption 1(i),  $|\rho|_{2+\epsilon} \leq M_t v_m$ , where  $M_t$  can be set so that  $M_t \leq 2E\|y_t\|^{2+\epsilon}$ , and  $v_m$  is a non-negative sequence of size  $v_m = O(m^{-\alpha})$ . Putting all these results together, and using Lemma 1, we have  $|x_t^a - x_t^b|_{2+\epsilon} = |x_t^a - E[w_t | w_{t-m}, \dots, w_{t+m}]|_{2+\epsilon} \leq M'_t v_m^{\frac{r-(2+\epsilon)}{2r-(2+\epsilon)}} = M'_t v'_m$ ; note that  $M'_t = M_t \left[ |B\rho|_r^{\frac{r}{2+\epsilon} - \frac{r(2+\epsilon-r)}{(2+\epsilon)-r}} + |B\rho|_r^{\frac{(2+\epsilon)r}{2r-(2+\epsilon)}} \right] |B|_{q^*/(q-1)}^{\frac{r-(2+\epsilon)}{2r-(2+\epsilon)}}$  is a sequence of positive, finite numbers. The size of  $v'_m$  is  $v'_m = O\left(m^{-\alpha \frac{r-(2+\epsilon)}{2r-(2+\epsilon)}}\right) = O(m^{-\alpha'})$ . Assumption 1(ii) entails that  $\alpha' > \frac{1}{2}$ . ■

**Lemma 3** Under Assumption 1,  $\text{vec}[\bar{w}_t \bar{w}'_{t-l} - E(\bar{w}_t \bar{w}'_{t-l})]$  is  $L_{1+\epsilon/2}$ -NED of size  $\alpha'$  on  $\{v_t\}_{t=-\infty}^{+\infty}$ , for every  $l$ .

**Proof.** The Lemma is an application of Theorem 17.9 in Davidson (1994, p. 268), where  $L_1$ - and  $L_2$ -norms are replaced, respectively, by  $L_{1+\epsilon/2}$ - and  $L_{2+\epsilon}$ -norms. ■

**Lemma 4** Under Assumption 1 and 2(i)(b)-(ii),  $\text{vec}[\bar{w}_t \bar{w}'_t - E(\bar{w}_t \bar{w}'_t)]$  is  $L_{2+\epsilon}$ -NED of size  $\alpha'$  on  $\{v_t\}_{t=-\infty}^{+\infty}$ .



**Proof.** The proof is very similar to that of Lemma 2. Assuming  $n = 1$  and letting  $\omega_t = w_t^2$

$$\begin{aligned} \left| \omega_t^a - \omega_t^b \right|_{2+\epsilon} &\leq \left| \left( x_t^a + x_t^b \right) \left( x_t^a - x_t^b \right) \right|_{2+\epsilon} \\ &= \left| \left( |y_t^a|^3 + |y_t^b|^3 + |y_t^a|^2 |y_t^b| + |y_t^a| |y_t^b|^2 \right) \left( |y_t^a - y_t^b| \right) \right|_{2+\epsilon} \\ &= \left| B \left( y_t^a, y_t^b \right) \rho \left( y_t^a, y_t^b \right) \right|_{2+\epsilon} = |B\rho|_{2+\epsilon}, \end{aligned}$$

so that the Lemma follows from Assumption 2(ii) and Lemma 1. ■

**Lemma 5** *Under Assumption R1, it holds that:*

- (i)  $\bar{w}_t^\varepsilon$  is  $L_{2+\epsilon}$ -NED of size  $\alpha' > \frac{1}{2}$  on  $\{v_t\}_{t=-\infty}^{+\infty}$ ;
- (ii) redefining  $\bar{w}_t^\varepsilon$  in a richer probability space,  $\sum_{t=1}^{\lfloor T\tau \rfloor} \bar{w}_t^\varepsilon = \sum_{t=1}^{\lfloor T\tau \rfloor} X_t^\varepsilon + O_{a.s.} \left( \lfloor T\tau \rfloor^{\frac{1}{2}-\delta} \right)$ , uniformly in  $\tau$ , where  $X_t^\varepsilon$  is a zero mean, i.i.d. Gaussian sequence with  $E(X_t^\varepsilon X_t^{\varepsilon'}) = V^\varepsilon$  and  $\delta > 0$ ;
- (iii)  $\hat{\beta} - \beta = O_{a.s.} \left( \sqrt{\frac{\ln \ln T}{T}} \right)$ ;
- (iv) defining  $\bar{\bar{w}}_t^\varepsilon = \hat{w}_t^\varepsilon - \text{vec}(\Sigma_\varepsilon)$ , it holds that  $\sum_{t=1}^{\lfloor T\tau \rfloor} \bar{\bar{w}}_t^\varepsilon = \sum_{t=1}^{\lfloor T\tau \rfloor} X_t^{\varepsilon*} + O_{a.s.} \left( \lfloor T\tau \rfloor^{\frac{1}{2}-\delta} \right)$ , uniformly in  $\tau$ , where  $X_t^{\varepsilon*}$  is a zero mean, i.i.d. Gaussian sequence with  $E(X_t^{\varepsilon*} X_t^{\varepsilon*'}) = V^\varepsilon$  and  $\delta > 0$ ;
- (v)  $\text{vec} [\bar{w}_t^\varepsilon \bar{w}_{t-l}^{\varepsilon'} - E(\bar{w}_t^\varepsilon \bar{w}_{t-l}^{\varepsilon'})]$  is  $L_{1+\epsilon/2}$ -NED of size  $\alpha'$  on  $\{v_t\}_{t=-\infty}^{+\infty}$ , for every  $l$ .

Further, under Assumptions R1(ii) and R2(i)-(ii), it holds that:

- (vi)  $\text{vec} [\bar{w}_t^\varepsilon \bar{w}_t^{\varepsilon'} - E(\bar{w}_t^\varepsilon \bar{w}_t^{\varepsilon'})]$  is  $L_{2+\epsilon}$ -NED of size  $\alpha'$  on  $\{v_t\}_{t=-\infty}^{+\infty}$ .

**Proof.** Parts (i) and (v) of the Lemma are the same as Lemmas 2 and 3 respectively. Part (ii) is the same as equation (2) in the main paper, and we refer to the proof of Theorem 1 in the main paper for details. Similarly, part (vi) of the Lemma is the same as Lemma 4.

As far as part (iii) is concerned, let  $n = q = 1$  for the sake of notational simplicity, and consider  $\hat{\beta} - \beta = \left( \sum_{t=1}^T x_t^2 \right)^{-1} \left( \sum_{t=1}^T x_t \varepsilon_t \right)$ . We start from the denominator. By virtue of Assumption R1(i), Lemma 2 entails that  $x_t^2$  is  $L_{2+\epsilon}$ -NED of size  $\alpha' > \frac{1}{2}$  on the strong mixing basis  $\{v_t\}_{t=-\infty}^{+\infty}$ . This result allows to use a SLLN, since the assumptions of Theorem 2.1 in Ling (2007) are satisfied; thus,  $\sum_{t=1}^T x_t^2 = O_{a.s.}(T)$ . Consider now the numerator; Assumptions R1(i)-(ii) entail that  $x_t \varepsilon_t$  is  $L_{2+\epsilon}$ -NED of size  $\alpha' > \frac{1}{2}$  on  $\{v_t\}_{t=-\infty}^{+\infty}$ ; this is a direct application of Theorem 17.17 in Davidson (1994, p. 273), and it follows the same lines as the proof of Lemma 2. Using this result and Assumption R1(iii), a SIP holds, based on the same passages as in the proof of Theorem 1 in the main paper, so that  $\sum_{t=1}^T x_t \varepsilon_t = \sum_{t=1}^T N_t^{x\varepsilon} + O_{a.s.} \left( T^{\frac{1}{2}-\delta} \right)$  for some  $\delta > 0$ , where  $N_t^{x\varepsilon}$  is a sequence of i.i.d. zero mean, Gaussian random variables with

covariance matrix  $V^{x\varepsilon}$ . Therefore, by the Law of the Iterated Logarithm (LIL),  $\sum_{t=1}^T x_t \varepsilon_t = O_{a.s.}(\sqrt{T \ln \ln T})$ . Part (iii) follows by combining the two results.

Finally, we show part (iv). We can write  $\sum_{t=1}^{\lfloor T\tau \rfloor} \bar{w}_t^\varepsilon = \sum_{t=1}^{\lfloor T\tau \rfloor} \bar{w}_t^\varepsilon - 2\text{vec} \left[ \left( \hat{\beta} - \beta \right) \left( \sum_{t=1}^{\lfloor T\tau \rfloor} x_t \varepsilon_t' \right) \right] + \text{vec} \left[ \left( \hat{\beta} - \beta \right) \left( \sum_{t=1}^{\lfloor T\tau \rfloor} x_t x_t' \right) \left( \hat{\beta} - \beta \right)' \right] = \sum_{t=1}^{\lfloor T\tau \rfloor} \bar{w}_t^\varepsilon + I + II$ . By virtue of the passages above,  $I$  and  $II$  are both  $O_{a.s.}(\ln \ln T)$ . Part (iv) then follows from part (ii). ■

**Lemma 6** *Under Assumptions R1(ii)-(iv)(a) and R2, it holds that  $\sup_{1 \leq \lfloor T\tau \rfloor \leq T} \left\| \tilde{V}_\tau^\varepsilon - V^\varepsilon \right\| = O_p\left(\frac{1}{m}\right) + O_p\left(m \ln T \sqrt{\frac{\ln \ln T}{T}}\right)$ .*

**Proof.** Let, preliminarily,  $\tilde{V}_\tau^{\varepsilon\varepsilon} = \left( \hat{\Psi}_{0,\lfloor T\tau \rfloor}^{\varepsilon\varepsilon} + \hat{\Psi}_{0,1-\lfloor T\tau \rfloor}^{\varepsilon\varepsilon} \right) + \sum_{l=1}^m \left( 1 - \frac{l}{m} \right) \left[ \left( \hat{\Psi}_{l,\lfloor T\tau \rfloor}^{\varepsilon\varepsilon} + \hat{\Psi}_{l,1-\lfloor T\tau \rfloor}^{\varepsilon\varepsilon} \right) + \left( \hat{\Psi}_{l,1-\lfloor T\tau \rfloor}^{\varepsilon\varepsilon} + \hat{\Psi}_{l,1-\lfloor T\tau \rfloor}^{\varepsilon\varepsilon} \right) \right]$ , where  $\hat{\Psi}_{l,\lfloor T\tau \rfloor}^{\varepsilon\varepsilon} = \frac{1}{T-(l+1)} \sum_{t=l+1}^{\lfloor T\tau \rfloor} \left[ w_t^\varepsilon - \text{vec} \left( \hat{\Sigma}_\varepsilon \right) \right] \left[ w_{t-l}^\varepsilon - \text{vec} \left( \hat{\Sigma}_\varepsilon \right) \right]'$ , and similarly for  $\hat{\Psi}_{l,1-\lfloor T\tau \rfloor}^{\varepsilon\varepsilon}$ . We start by showing that  $\sup_{1 \leq \lfloor T\tau \rfloor \leq T} \left\| \tilde{V}_\tau^{\varepsilon\varepsilon} - V^\varepsilon \right\| = O_p\left(\frac{1}{m}\right) + O_p\left(\frac{m}{\sqrt{T}}\right)$ ; the proof is very similar to that of Theorem 2 in the main paper, and therefore only the most relevant passages are reported. Let  $\hat{\Sigma}_{\varepsilon\varepsilon} = T^{-1} \sum_{t=1}^T \varepsilon_t \varepsilon_t'$ , and note that  $\hat{\Sigma}_\varepsilon - \hat{\Sigma}_{\varepsilon\varepsilon} = \left( \hat{\beta} - \beta \right) \left( T^{-1} \sum_{t=1}^T x_t x_t' \right) \left( \hat{\beta} - \beta \right)' - \left( T^{-1} \sum_{t=1}^T \varepsilon_t x_t' \right) \left( \hat{\beta} - \beta \right)' - \left( \hat{\beta} - \beta \right) \left( T^{-1} \sum_{t=1}^T x_t \varepsilon_t' \right) = I + II + II'$ . By Lemma 5, we have  $I = O_{a.s.}\left(\frac{\ln \ln T}{T}\right)$ , and similarly  $II = O_{a.s.}\left(\frac{\ln \ln T}{T}\right)$ . Thus, for every  $l$ ,  $\hat{\Psi}_{l,\lfloor T\tau \rfloor}^{\varepsilon\varepsilon} = [T - (l+1)]^{-1} \sum_{t=l+1}^{\lfloor T\tau \rfloor} \left[ w_t^\varepsilon - \text{vec} \left( \hat{\Sigma}_{\varepsilon\varepsilon} \right) \right] \left[ w_{t-l}^\varepsilon - \text{vec} \left( \hat{\Sigma}_{\varepsilon\varepsilon} \right) \right]' + O_{a.s.}\left(\frac{\ln \ln T}{T}\right)$ . By using part (ii) of Lemma 5, it follows that  $\hat{\Sigma}_{\varepsilon\varepsilon} - \Sigma_\varepsilon = O_{a.s.}\left(T^{-\frac{1}{2}-\delta}\right)$ , so that  $\hat{\Psi}_{l,\lfloor T\tau \rfloor}^{\varepsilon\varepsilon} = [T - (l+1)]^{-1} \sum_{t=l+1}^{\lfloor T\tau \rfloor} \bar{w}_t^\varepsilon \bar{w}_{t-l}^{\varepsilon'} + O_{a.s.}\left(T^{-\frac{1}{2}-\delta}\right)$ ; the same holds for  $\hat{\Psi}_{l,1-\lfloor T\tau \rfloor}^{\varepsilon\varepsilon}$ . Therefore,  $\hat{\Psi}_{l,\lfloor T\tau \rfloor}^{\varepsilon\varepsilon} + \hat{\Psi}_{l,1-\lfloor T\tau \rfloor}^{\varepsilon\varepsilon} - \Psi_l^{\varepsilon\varepsilon} = [T - (l+1)]^{-1} \sum_{t=l+1}^T \left[ \bar{w}_t^\varepsilon \bar{w}_{t-l}^{\varepsilon'} - E(\bar{w}_t^\varepsilon \bar{w}_{t-l}^{\varepsilon'}) \right] + O_{a.s.}\left(T^{-\frac{1}{2}-\delta}\right)$ . Henceforth, the proof is the same as for equation (6) in the main paper, by virtue of part (vi) of Lemma 5 and since the assumptions on the autocovariances  $\Psi_l^\varepsilon$  are the same as in Assumption 2. Finally, we show that  $\sup_{1 \leq \lfloor T\tau \rfloor \leq T} \left\| \tilde{V}_\tau^{\varepsilon\varepsilon} - \tilde{V}_\tau^\varepsilon \right\| = O_p\left(m \sqrt{\frac{\ln \ln T}{T}}\right)$ . Consider the quantities  $I = [T - (l+1)]^{-1} \sum_{t=l+1}^{\lfloor T\tau \rfloor} 2\text{vec} \left[ \left( \hat{\beta} - \beta \right)' x_t \varepsilon_t' \left[ w_{t-l}^\varepsilon - \text{vec} \left( \hat{\Sigma}_\varepsilon \right) \right]' \right]$  and  $II = [T - (l+1)]^{-1} \sum_{t=l+1}^{\lfloor T\tau \rfloor} \text{vec} \left[ \left( \hat{\beta} - \beta \right) \left( x_t x_t' \right) \left( \hat{\beta} - \beta \right)' \left[ w_{t-l}^\varepsilon - \text{vec} \left( \hat{\Sigma}_\varepsilon \right) \right]' \right]$ . We have  $I \leq \left\| \hat{\beta} - \beta \right\| \left\{ [T - (l+1)]^{-1} \sum_{t=l+1}^{\lfloor T\tau \rfloor} \|x_t \varepsilon_t'\|^2 \right\}^{1/2} \left\{ [T - (l+1)]^{-1} \sum_{t=l+1}^{\lfloor T\tau \rfloor} \left\| w_{t-l}^\varepsilon - \text{vec} \left( \hat{\Sigma}_\varepsilon \right) \right\|^2 \right\}^{1/2}$ . By applying Theorem 17.9 in Davidson (1994, p. 268) following similar lines as in the proof of Lemma 3,  $\|x_t \varepsilon_t'\|^2$  can be shown to be  $L_{1+\epsilon/2}$ -NED of size  $\alpha' > \frac{1}{2}$  on  $\{v_t\}_{t=-\infty}^{+\infty}$ ; thus, a SLLN (see e.g. Theorem 2.1 in Ling, 2007) can be applied with  $[T\tau - (l+1)]^{-1} \sum_{t=l+1}^{\lfloor T\tau \rfloor} \|x_t \varepsilon_t'\|^2 = O_{a.s.}(1)$ ; similarly, part (v) of Lemma 5 entails a SLLN whereby  $[T\tau - (l+1)]^{-1} \sum_{t=l+1}^{\lfloor T\tau \rfloor} \left\| w_{t-l}^\varepsilon - \text{vec} \left( \hat{\Sigma}_\varepsilon \right) \right\|^2 = O_{a.s.}(1)$ ; hence,  $I = O_{a.s.}\left(\sqrt{\frac{\ln \ln T}{T}}\right)$ . By the same logic,  $II = O_{a.s.}\left(\frac{\ln \ln T}{T}\right)$ . After some algebra,  $\sup_{1 \leq \lfloor T\tau \rfloor \leq T} \left\| \tilde{V}_\tau^{\varepsilon\varepsilon} - \tilde{V}_\tau^\varepsilon \right\| = O_p\left(m \sqrt{\frac{\ln \ln T}{T}}\right)$  follows. ■

**Lemma 7** Under Assumptions R1 and R3, it holds that as  $T \rightarrow \infty$ , uniformly in  $\tau$ ,  $\hat{\lambda}_{i,\tau}^\varepsilon - \lambda_i^\varepsilon = (x_i^{\varepsilon'} \otimes x_i^{\varepsilon'}) \text{vec}(\hat{\Sigma}_{\varepsilon,\tau} - \Sigma_\varepsilon) + O_p(T^{-1})$ ,  $\hat{x}_{i,\tau}^\varepsilon - x_{\varepsilon,i} = v_{x,i}^\varepsilon \text{vec}(\hat{\Sigma}_{\varepsilon,\tau} - \Sigma_\varepsilon) + O_p(T^{-1})$  and  $\hat{\gamma}_{i,\tau}^\varepsilon - \gamma_i^\varepsilon = v_{\gamma,i}^\varepsilon \text{vec}(\hat{\Sigma}_{\varepsilon,\tau} - \Sigma_\varepsilon) + O_p(T^{-1})$ .

**Proof.** By part (i) of Lemma 5, an FCLT holds for  $\text{vec}(\hat{\Sigma}_{\varepsilon,\tau} - \Sigma_\varepsilon)$ , so that  $\sup_{\lfloor T\tau \rfloor} \|\hat{\Sigma}_{\varepsilon,\tau} - \Sigma_\varepsilon\| = O_p(T^{-1/2})$  - see the proof of equation (1) in Theorem 1 in the main paper. Henceforth, the proof is the same as for Proposition 1 in the main paper. ■

**Proof of Theorem 1.** By Lemma 6, it suffices to show that  $\left| \sup_{\frac{1}{T} \leq \tau \leq 1 - \frac{1}{T}} \hat{\Lambda}_{iT}^\varepsilon(\tau) - \sup_{\frac{1}{T} \leq \tau \leq 1 - \frac{1}{T}} M^\varepsilon(\tau) \right| = o_p\left[(\ln \ln T)^{-1/2}\right]$ , where  $M^\varepsilon(\tau) = \left[ \sum_{i=1}^p \frac{B_{1i}^2(\tau)}{\tau(1-\tau)} \right]^{1/2}$ ,  $B_{1i}(\tau)$  is a sequence of standard, independent Brownian bridges for  $i = 1, \dots, p$ , and

$$\hat{\Lambda}_{iT}^\varepsilon(\tau) = \sqrt{\frac{T}{\lfloor T\tau \rfloor \times \lfloor T(1-\tau) \rfloor}} \times \left[ \tilde{S}^\varepsilon(\tau)' (V_{\varepsilon,\tau})^{-1} \tilde{S}^\varepsilon(\tau) \right]^{1/2}.$$

Indeed, we can show that  $\left| \sup_{\frac{1}{T} \leq \tau \leq 1 - \frac{1}{T}} \ddot{\Lambda}_{iT}^\varepsilon(\tau) - \sup_{\frac{1}{T} \leq \tau \leq 1 - \frac{1}{T}} \hat{\Lambda}_{iT}^\varepsilon(\tau) \right| = O_p\left(\frac{\ln \ln T}{T}\right)$ , where

$$\ddot{\Lambda}_{iT}^\varepsilon(\tau) = \sqrt{\frac{T}{\lfloor T\tau \rfloor \times \lfloor T(1-\tau) \rfloor}} \times \left[ \ddot{S}^\varepsilon(\tau)' (V_{\varepsilon,\tau})^{-1} \ddot{S}^\varepsilon(\tau) \right]^{1/2},$$

and  $\ddot{S}^\varepsilon(\tau) = \sum_{t=1}^{\lfloor T\tau \rfloor} \text{vec}(\varepsilon_t \varepsilon_t')$  with  $\ddot{S}^\varepsilon(\tau) = 0$  for  $\tau \leq \frac{1}{T}$  or  $\geq 1 - \frac{1}{T}$ . Since  $\left| \sup_{\frac{1}{T} \leq \tau \leq 1 - \frac{1}{T}} \ddot{\Lambda}_{iT}^\varepsilon(\tau) - \sup_{\frac{1}{T} \leq \tau \leq 1 - \frac{1}{T}} \hat{\Lambda}_{iT}^\varepsilon(\tau) \right| \leq \sup_{\frac{1}{T} \leq \tau \leq 1 - \frac{1}{T}} \left| \ddot{\Lambda}_{iT}^\varepsilon(\tau) - \hat{\Lambda}_{iT}^\varepsilon(\tau) \right|$ , the passages in the proof of Lemma 6 and the same logic as in the proof of Theorem 2 in the main paper yield  $\sup_{\frac{1}{T} \leq \tau \leq 1 - \frac{1}{T}} \left| \ddot{\Lambda}_{iT}^\varepsilon(\tau) - \hat{\Lambda}_{iT}^\varepsilon(\tau) \right| = O_p\left(\frac{\ln \ln T}{T}\right)$ . Thus, it suffices to show that  $\left| \sup_{\frac{1}{T} \leq \tau \leq 1 - \frac{1}{T}} \ddot{\Lambda}_{iT}^\varepsilon(\tau) - \sup_{\frac{1}{T} \leq \tau \leq 1 - \frac{1}{T}} M^\varepsilon(\tau) \right| = o_p\left[(\ln \ln T)^{-1/2}\right]$ . This can be shown following the passages in the proof of Theorem 3 in the main paper, using the SIP in part (ii) of Lemma 5 and Lemma 7. The behaviour of the test statistic under the alternative can be shown following the same lines as the proof of Theorem 4 in the main paper. ■

## 5 Tables

$n$	$T$		$(\rho, \theta)$	$(0, 0)$	$(0.5, 0)$	$(0, 0.5)$	$(0, -0.5)$
3	50	Kao et al.		0.001	0.004	0.004	0.004
		Aue et al.		0.002	0.007	0.004	0.006
	200	Kao et al.		0.022	0.049	0.042	0.052
		Aue et al.		0.030	0.035	0.035	0.043
	500	Kao et al.		0.046	0.052	0.053	0.063
		Aue et al.		0.037	0.051	0.056	0.058
5	50	Kao et al.		0.002	0.000	0.000	0.000
		Aue et al.		0.001	0.000	0.000	0.000
	200	Kao et al.		0.018	0.032	0.037	0.045
		Aue et al.		0.018	0.018	0.023	0.025
	500	Kao et al.		0.032	0.053	0.060	0.064
		Aue et al.		0.026	0.038	0.052	0.054
7	50	Kao et al.		0.000	0.000	0.000	0.000
		Aue et al.		0.000	0.000	0.000	0.000
	200	Kao et al.		0.010	0.014	0.017	0.035
		Aue et al.		0.008	0.015	0.011	0.016
	500	Kao et al.		0.029	0.040	0.062	0.062
		Aue et al.		0.027	0.034	0.031	0.046

**Table A1.** Empirical rejection frequencies for the null of no change in  $\Sigma$ , based on a test for the stability of all eigenvalues. Data are generated according to equation (22) in the main paper.

$n$	$T$		$\Delta = \sqrt{\frac{\ln \ln T}{T^{1/2}}}$				first element only	
			$(\rho, \theta)$	$(0, 0)$	$(0.5, 0)$	$(0, 0.5)$	$(0, -0.5)$	
3	50	Kao et al.		0.027	0.034	0.037	0.034	0.077
		Aue et al.		0.176	0.155	0.138	0.145	0.128
	200	Kao et al.		0.834	0.740	0.716	0.670	0.903
		Aue et al.		0.968	0.902	0.879	0.854	0.966
	500	Kao et al.		0.998	0.987	0.985	0.961	0.990
		Aue et al.		1.000	1.000	1.000	1.000	1.000
5	50	Kao et al.		0.000	0.000	0.001	0.000	0.000
		Aue et al.		0.032	0.026	0.029	0.027	0.044
	200	Kao et al.		0.969	0.884	0.875	0.832	0.805
		Aue et al.		0.998	0.980	0.975	0.951	0.921
	500	Kao et al.		1.000	0.999	1.000	0.998	0.996
		Aue et al.		1.000	1.000	1.000	0.999	1.000
7	50	Kao et al.		0.000	0.000	0.000	0.000	0.000
		Aue et al.		0.000	0.000	0.000	0.000	0.007
	200	Kao et al.		0.973	0.922	0.905	0.878	0.699
		Aue et al.		1.000	0.993	0.992	0.987	0.850
	500	Kao et al.		1.000	1.000	1.000	1.000	0.994
		Aue et al.		1.000	1.000	1.000	1.000	1.000

**Table A2.** Power of the test for the null of no changes in  $\Sigma$ , based on a test for the stability of all eigenvalues. Data are generated according to equation (22) in the main paper and under the alternative hypothesis specified in equation (25) in the main paper. The last column refers to a break of the same size, but only for the first element of  $\Sigma$ .

$n$	$T$		first element only				
			$(\rho, \theta)$	$(0, 0)$	$(0.5, 0)$	$(0, 0.5)$	$(0, -0.5)$
3	50	Kao et al.		0.003	0.010	0.011	0.012
		Aue et al.		0.038	0.042	0.047	0.043
	200	Kao et al.		0.502	0.230	0.233	0.149
		Aue et al.		0.076	0.078	0.090	0.089
	500	Kao et al.		0.963	0.606	0.589	0.325
		Aue et al.		0.075	0.074	0.082	0.084
5	50	Kao et al.		0.000	0.000	0.001	0.000
		Aue et al.		0.001	0.005	0.000	0.003
	200	Kao et al.		0.816	0.525	0.514	0.352
		Aue et al.		0.068	0.062	0.067	0.070
	500	Kao et al.		1.000	0.973	0.963	0.813
		Aue et al.		0.055	0.062	0.058	0.080
7	50	Kao et al.		0.000	0.000	0.000	0.000
		Aue et al.		0.000	0.000	0.000	0.000
	200	Kao et al.		0.855	0.602	0.572	0.484
		Aue et al.		0.032	0.050	0.040	0.050
	500	Kao et al.		1.000	0.997	0.997	0.967
		Aue et al.		0.057	0.067	0.058	0.085

**Table A3.** Power of the test for the null of no changes in  $\Sigma$ , based on a test for the stability of all eigenvalues. Data are generated according to equation (22) in the main paper and under the alternative hypothesis specified in equation (27) in the main paper, with  $k = Th_{nT} + 1$  and  $\Delta = I_n$ . The last column refers to a case where only the first element of  $\Sigma$  undergoes the break specified before.

$n$	$T$	size	power, mid-sample	power, non mid-sample	size	power, mid-sample	power, non mid-sample
3	50	0.053	0.240	0.356	0.050	0.309	0.568
	100	0.046	0.395	0.347	0.035	0.471	0.776
	200	0.065	0.773	0.469	0.063	0.838	0.993
	500	0.061	0.941	0.651	0.050	0.958	1.000
5	50	0.046	0.193	0.292	0.037	0.260	0.331
	100	0.062	0.334	0.383	0.055	0.472	0.464
	200	0.057	0.631	0.556	0.061	0.775	0.648
	500	0.066	0.877	0.935	0.065	0.955	0.954
7	50	0.041	0.180	0.204	0.046	0.274	0.261
	100	0.053	0.284	0.237	0.050	0.457	0.263
	200	0.062	0.628	0.268	0.067	0.739	0.325
	500	0.059	0.844	0.518	0.065	0.945	0.632

**Table B1. Size and power versus a break in the first principal component. Data are simulated as indicated in the main paper; particularly, the first eigenvalue of the covariance matrix is set to 1; the others are simulated from a Uniform distribution with support  $[0, 0.02]$ , thereby having mean 0.01. The last three columns contain empirical rejection frequencies under the same set-up, but all eigenvalues are of comparable magnitude, having mean 1.**

$n$	$T$	size	power, mid-sample	power, non mid-sample	size	power, mid-sample	power, non mid-sample
3	50	0.048	0.308	0.386	0.041	0.190	0.056
	100	0.053	0.559	0.673	0.050	0.431	0.106
	200	0.043	0.862	0.915	0.064	0.759	0.233
	500	0.055	0.982	0.999	0.067	0.917	0.510
5	50	0.040	0.446	0.369	0.039	0.189	0.031
	100	0.056	0.752	0.674	0.061	0.494	0.112
	200	0.047	0.944	0.930	0.069	0.818	0.229
	500	0.061	0.995	0.995	0.050	0.961	0.520
7	50	0.048	0.479	0.386	0.040	0.210	0.049
	100	0.045	0.768	0.654	0.063	0.489	0.122
	200	0.053	0.949	0.918	0.064	0.816	0.221
	500	0.061	0.996	1.000	0.060	0.964	0.497

**Table B2.** Size and power for the test for the null of no changes in the largest eigenvalue of  $\Sigma$ , in presence of a large autoregressive root, after pre-whitening. Data are simulated as indicated in the main paper.



	monthly data						weekly data					
	mean	std. dev	skew	kurt	AR(1)	ARCH(7)	mean	std. dev	skew	kurt	AR(1)	ARCH(7)
1m	-0.035	0.438	-0.481	25.266	-0.129	0.000***	-0.007	0.159	0.119	66.052	0.151	0.000***
3m	-0.035	0.366	-0.030	24.291	0.097	0.000***	-0.008	0.121	-0.664	46.650	0.324	0.000***
6m	-0.036	0.295	-1.591	15.137	0.209	0.011**	-0.008	0.108	-1.525	21.640	0.215	0.000***
9m	-0.037	0.274	-1.624	10.921	0.257	0.668	-0.008	0.121	-1.301	14.454	0.035	0.000***
12m	-0.037	0.262	-1.188	8.251	0.286	0.989	-0.008	0.136	-1.252	16.110	-0.098	0.000***
15m	-0.038	0.267	-0.851	6.835	0.268	0.972	-0.008	0.130	-0.883	9.984	-0.049	0.000***
18m	-0.038	0.273	-0.549	5.589	0.242	0.957	-0.008	0.127	-0.474	6.827	-0.014	0.000***
21m	-0.038	0.282	-0.323	4.695	0.210	0.962	-0.008	0.129	-0.138	5.693	-0.003	0.000***
24m	-0.038	0.294	-0.160	4.086	0.172	0.974	-0.008	0.134	0.026	5.298	-0.012	0.000***
30m	-0.038	0.303	-0.043	3.937	0.146	0.984	-0.008	0.139	0.080	4.930	-0.022	0.000***
3y	-0.038	0.314	0.052	3.982	0.118	0.980	-0.007	0.144	0.101	4.741	-0.029	0.000***
4y	-0.036	0.319	0.060	4.198	0.073	0.970	-0.007	0.149	0.041	4.560	-0.045	0.000***
5y	-0.035	0.323	0.116	4.706	0.037	0.988	-0.006	0.152	0.007	4.486	-0.051	0.000***
6y	-0.033	0.321	0.140	5.151	0.033	0.994	-0.006	0.152	-0.029	4.555	-0.052	0.000***
7y	-0.032	0.320	0.144	5.560	0.023	0.995	-0.005	0.152	-0.028	4.596	-0.049	0.000***
8y	-0.031	0.318	0.138	5.868	0.017	0.994	-0.005	0.152	-0.033	4.776	-0.044	0.000***
9y	-0.030	0.318	0.112	6.251	0.009	0.992	-0.005	0.151	-0.063	4.968	-0.047	0.000***
10y	-0.030	0.318	0.048	6.653	-0.001	0.990	-0.005	0.151	-0.059	4.946	-0.046	0.000***

Table C. Descriptive statistics for monthly and weekly data. We report the mean of the returns, the standard deviation, the skewness and the kurtosis, the AR(1) coefficient and the  $p$ -value of the ARCH(7) test for the returns - \*, \*\*, and \*\*\* denote rejection of the null of no ARCH effects at 10%, 5% and 1% levels respectively.

$(T, p)$		$level$	10%	5%	1%
(163, 1)	$cv_1$		2.8044	3.0716	3.5758
	$cv_3$		3.1987	3.4222	3.9344
	$cv_{18}$		3.7843	3.9690	4.1273
(712, 1)	$cv_1$		2.9344	3.1861	3.6524
	$cv_3$		3.3197	3.5131	3.9321
	$cv_{18}$		3.8242	3.9655	4.3805
(163, 18)	$cv_1$		6.1897	6.4075	6.7747
	$cv_3$		6.5068	6.6486	7.0053
(712, 18)	$cv_1$		6.2903	6.5016	6.9779
	$cv_3$		6.6279	6.8287	7.1674
(166, 1)	$cv_1$		2.7952	3.0336	3.5951
	$cv_4$		3.2688	3.5213	3.9874
(166, 10)	$cv_{10}$		5.0741	5.3172	5.7916

Table D. Critical values for the applications in Section 4 in the main paper, and in Section 2 in this supplement. In the Table, the notation  $cv_N$  refers to the critical value to be used when  $N$  hypotheses are being tested for, in order to have a procedure-wise level of 10%, 5% and 1% respectively. The panels with  $(T, p) = (163, 1)$  and  $(732, 1)$  contain critical values for unidimensional tests (monthly and weekly frequencies respectively), and therefore are used to test for changes in eigenvalues or when verifying the stability of the diagonal elements of  $\Sigma$  one at a time. Panels where  $(T, p) = (163, 1)$  and  $(732, 1)$  contain critical values for tests with 18 hypotheses under the null, and thus are designed for tests for the stability of one eigenvector. Finally, panels with  $(T, p) = (166, 1)$  and  $(166, 10)$  contain critical values for the application in Section 2: the former are for changes in the eigenvalues, both individual and family-wise tests unidimensional tests (monthly and weekly frequencies respectively); the latter are for testing the stability of the whole matrix.

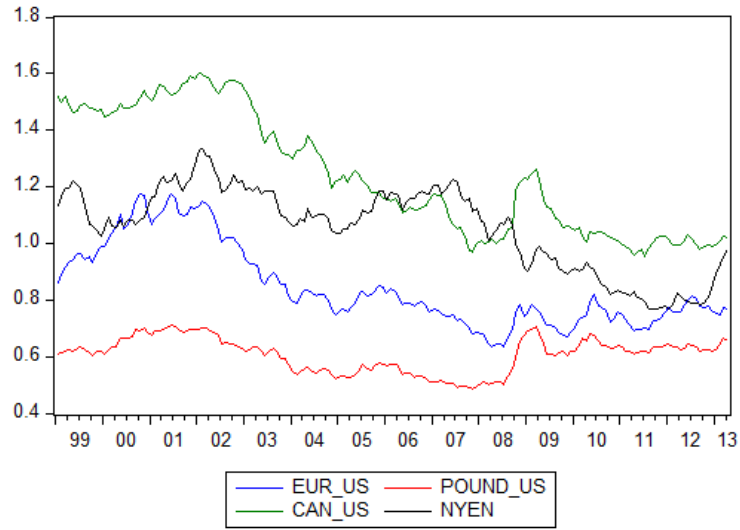


Figure A. Exchange rates vis-a-vis the US Dollar. Data are monthly, recorded between January 1999 and April 2013; as far as the Yen/US Dollar exchange rate is concerned, the original data have been divided by 100 to make figures comparable with the other exchange rates.

	mean	std. dev	skew	kurt	AR(1)	ARCH(7)
EUR	−0.0007	0.0252	−0.0026	2.990	0.297	0.113
GBP	0.0004	0.0219	0.4460	4.841	0.294	0.000***
CAN	−0.0023	0.0192	0.9417	9.774	0.272	0.505
YEN	−0.0009	0.0238	−0.1859	2.983	0.263	0.385

Table E1. Descriptive statistics for the returns on (vis-a-vis the US Dollar): the Euro (EUR), the British Pound (GBP), the Canadian Dollar (CAN) and the Yen (YEN). See Table 4 for an explanation of the output.

		<i>Estimation Output</i>			
		EUR	GBP	CAN	YEN
<i>Constant</i>		−0.0015	0.0004	−0.0020	−0.0004
	EUR	0.3760*	0.2080*	0.1604	−0.0408*
<i>Lag 1</i>	GBP	−0.1642	0.0442	−0.1053	−0.0650
	CAN	0.2966*	0.4057*	0.2786*	0.0667
	YEN	−0.0389	−0.0139	−0.0396	0.2045
<i>Lag 2</i>	EUR	−0.2868*	−0.2044*	−0.1499	−0.0674
	GBP	0.2425	0.2680*	0.1300	−0.0614
	CAN	−0.1387	−0.1891	0.0099	0.0169
	YEN	0.1462	−0.0405	0.0394	0.1195
<i>Lag 5</i>	EUR	0.0710	0.0165	0.2031*	0.0779
	GBP	0.0500	−0.1051	−0.0851	−0.1171
	CAN	−0.1385	0.0119	−0.1180	0.2350*
	YEN	−0.1603	−0.0684	−0.1311*	−0.2036*

		<i>Mis-specification tests</i>	
Serial Correlation LM test		134.378	[0.904]
Normality		100.308	[0.000]

**Table E2.** Estimation results and misspecification analysis for the VAR; the components of the vector are the following (vis-a-vis the US Dollar): the Euro (EUR), the British Pound (GBP), the Canadian Dollar (CAN) and the Yen (YEN). In the top half of the Table we report the estimation results: the symbol “\*” denotes rejection of the null that the corresponding coefficient is zero at 5% level. The second half of the Table contains misspecification tests: we report (in order of appearance) the portmanteu serial correlation test (up to lag 12) and the Jarque-Bera test.

$H_0 : \Sigma \text{ constant}$	4.8049	$H_0 : \lambda_1 \text{ constant}$	1.5457
		$H_0 : \lambda_2 \text{ constant}$	2.2527
		$H_0 : \lambda_3 \text{ constant}$	1.7077
		$H_0 : \lambda_4 \text{ constant}$	3.4154* [June 2008]

**Table E3.** Tests for changes in the covariance matrix of the error term of the VAR estimated in Table E2. Rejection at 10%, 5% and 1% levels are denoted with \*, \*\* and \*\*\* respectively. Where present, numbers in square brackets are the estimated breakdates, defined as  $T \times \arg \max \Lambda_T^\varepsilon(\tau)$ .

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