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Tests of risk premia in linear factor models

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Tests of Risk Premia in Linear Factor Models

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Abstract

We show that inference on risk premia in linear factor models that is based on the Fama-MacBeth and GLS risk premia estimators is misleading when the β 's are small and/or the number of assets is large. We propose novel statistics that remain trustworthy in these cases. The inadequacy of Fama-MacBeth and GLS based Wald statistics is highlighted in a power comparison and using daily portfolio returns from Jagannathan and Wang (1996). The power comparison shows that the Fama-MacBeth and GLS Wald statistics can be severely size distorted. The daily portfolio returns from Jagannathan and Wang (1996) reveal a large discrepancy between the 95% confidence sets for the risk premia that result from the different statistics. The Fama-MacBeth and GLS Wald statistics imply tight 95% confidence sets for the risk premia on the yield premium and labor income growth that only contain small values while the 95% confidence sets that result from statistics that remain trustworthy in case of small β 's imply much larger 95% confidence sets at considerably higher values of the risk premia.

1 Introduction

Linear factor models are amongst the most commonly used statistical models in finance, see *e.g.* Lintner (1965), Fama and MacBeth (1973), Gibbons (1982), Shanken (1992), Fama and French (1992,1993,1996) and Jagannathan and Wang (1996). This results as many well-known financial models, like, for example, the capital asset pricing model, imply a linear factor structure for the asset returns. Linear factor models imply risk premia for the different factors that result from a cross-sectional regression of the asset returns on the factor β 's where the β 's are obtained from a time-series regression of the asset returns on the factors. This two step procedure is typically referred to as the Fama-MacBeth (FM) regression, see Fama and MacBeth (1973). Linear regression estimates are sensitive to collinearity of the explanatory variables so the estimates of the risk premia that result from a FM regression are sensitive to collinearity of the β 's. Collinearity of the β 's occurs when they are close or equal to zero. Kan and Zhang (1999), for example, show that the estimates of the risk premia that result from a FM regression are erroneous when the β 's are zero and the expected asset returns are non-zero.

FM regressions lead to misleading risk premia estimates in more cases than just the one indicated by Kan and Zhang (1999). We show that FM risk premia estimates are spurious whenever the β 's are relatively small and this is further aggravated when the number of assets is large. Hence, even for

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size-able values of the β 's, the risk premia estimates can be misleading when the number of assets is large. The FM risk premia estimates that are obtained in such settings are typically too small and their 95% confidence sets indicate an erroneous high level of precision. The erroneous level of precision of the confidence sets shows that statistical inference based on standard t -statistics is misleading and that the large sample distribution of the t -statistic differs from the normal one. We therefore propose a set of statistics whose large sample distributions remain unaltered when the β 's are close to zero. Thus these statistics can be used in a trustworthy manner.

The paper is organized as follows. In the second section, we discuss the FM and Generalized Least Squares (GLS) risk premia estimators and construct their large sample distributions when the β 's are small. These large sample distributions are non-standard and imply that Wald statistics, like, for example, the t -statistic, are misleading when the β 's are small or when the number of assets is large. The third section proposes a number of novel statistics whose large sample distributions remain unaltered when the β 's are small and/or the number of assets is large. These statistics are based on the Generalized Method of Moment (GMM) and instrumental variable statistics that are proposed in Kleibergen (2002,2005a) and Moreira (2003). The fourth section conducts a size and power comparison. It illustrates the size distortions of the Wald t -statistics that result from the FM and GLS risk premia estimators. The fifth section generalizes the test statistics to allow for tests on some of the risk premia. The sixth section uses data from Jagannathan and Wang (1999) to estimate a linear factor model for the daily portfolio returns. It shows, that because of the large number of portfolios and the relatively small β 's of the yield premium and labor income growth, that the FM and GLS risk premia estimators for the risk premia on the yield premium and labor income growth are downward biased and lead to small 95% confidence sets at rather low values. The 95% confidence sets of the statistics that remain trustworthy in case of many portfolios and small β 's indicate much larger 95% confidence sets at considerably higher values of the risk premia. This shows the importance of usage of these statistics for inference on the risk premia.

We use the following notation throughout the paper: $E(a)$ is the expected value of the random variable a , $\text{vec}(A)$ stands for the column vectorization of the $T \times n$ dimensional matrix A , $\text{vec}(A) = (a'_1 \dots a'_n)'$ when $A = (a_1 \dots a_n)$, $\text{vecinv}_n((a'_1 \dots a'_n)') = A$, $P_A = A(A'A)^{-1}A'$ and $M_A = I_T - P_A$ for a full rank matrix A and the $T \times T$ identity matrix I_T , ι_n is a $n \times 1$ vector of ones, " \xrightarrow{p} " indicates convergence in probability, " \xrightarrow{d} " indicates convergence in distribution and " $\overset{a}{=}$ " implies equality in large samples.

2 Risk Premia estimators and small β 's

We analyze the FM estimator in a GMM setting, see *e.g.* Hansen (1982). We therefore specify the moments of the factor related (excess) assets returns¹ as, see *e.g.* Cochrane (2001),

$$\begin{aligned} E(R_t) &= \iota_n \lambda_1 + \beta \lambda_F \\ \text{cov}(R_t, F_t) &= \beta \text{var}(F_t) \\ E(F_t) &= \mu_F, \end{aligned} \tag{1}$$

with R_t : the $n \times 1$ vector of (excess) asset returns, F_t : the $k \times 1$ vector of factors, λ_1 : the zero- β return, λ_F : the $k \times 1$ vector of factor risk premia, β : the $n \times k$ matrix of asset β 's, μ_F : the $k \times 1$ vector of factor means, ι_n : a $n \times 1$ vector of ones and t : the time index of the observation, $t = 1, \dots, T$. The

¹All asset returns are in deviation from the riskless return.

moments conditions (1) result from a linear factor model for the asset returns and factors,

$$\begin{aligned} R_t &= \iota_n \lambda_1 + \beta(F_t - \mu_F + \lambda_F) + u_t \\ F_t &= \mu_F + v_t \end{aligned} \quad (2)$$

with u_t and $v_t : n \times 1$ and $k \times 1$ vectors of disturbances which we assume to be identically and independently distributed with mean zero and $n \times n$ and $k \times k$ dimensional covariance matrices Λ and V_{FF} . Because $\beta = \text{cov}(R_t, F_t) \text{var}(F_t)^{-1}$, the covariance between the disturbances u_t and v_t is equal to βV_{FF} . The linear factor model (2) can as well be specified as

$$\left. \begin{aligned} R_t - \iota_n \lambda_1 - \beta(F_t - \bar{F} + \bar{F} - \mu_F + \lambda_F) &= u_t \\ F_t - \mu_F &= v_t \end{aligned} \right\} \Leftrightarrow \left\{ \begin{aligned} R_t - \iota_n \lambda_1 - \beta(\bar{F}_t + \lambda_F) &= \varepsilon_t \\ F_t - \mu_F &= v_t \end{aligned} \right. \quad (3)$$

where $\bar{F}_t = F_t - \bar{F}$, $\bar{F} = \frac{1}{T} \sum_{t=1}^T F_t$, $\varepsilon_t = u_t + \beta \bar{v}$ and $\bar{v} = \frac{1}{T} \sum_{t=1}^T v_t$. Because of the specification of β , weighted averages of ε_t and v_t are uncorrelated in large samples. We therefore use the latter specification in (3) to make an assumption about the behavior of the disturbances of the linear factor model (2) in large samples.

Assumption 1. *When the number of time series observations T becomes large,*

$$\frac{1}{\sqrt{T}} \sum_{t=1}^T \left(\begin{pmatrix} 1 \\ F_t \end{pmatrix} \otimes (R_t - \iota_n \lambda_1 - \beta(\bar{F}_t + \lambda_F)) \right) \xrightarrow{d} \begin{pmatrix} \varphi_R \\ \varphi_\beta \\ \varphi_F \end{pmatrix}, \quad (4)$$

with $(\varphi'_R \vdots \varphi'_\beta \vdots \varphi'_F)' \sim N(0, V)$ and

$$V = \begin{pmatrix} (Q \otimes \Omega) & 0 \\ 0 & V_{FF} \end{pmatrix}, \quad (5)$$

with $Q = \begin{pmatrix} Q_{11} & Q_{1F} \\ Q_{F1} & Q_{FF} \end{pmatrix} = \begin{pmatrix} 1 & \mu'_F \\ \mu_F & V_{FF} + \mu_F \mu'_F \end{pmatrix} : (k+1) \times (k+1)$, $\Omega : n \times n$, $V_{FF} : k \times k$, $V_{FF} = \text{var}(F_t)$ and $\Omega = \Lambda - \beta V_{FF} \beta'$.

Assumption 1 is a central limit theorem for the disturbances of the linear factor model (2) interacted with a constant and the factors. It holds under rather mild conditions, like, for example, Assumptions 1 and 2 from Shanken (1992). Because of the independence of the disturbances over time and their finite variance, Assumption 1 holds under the linear factor model (2) with a constant covariance matrix as well. Appendix A shows the implications of Assumption 1 for a central limit theorem for the original specification of the linear factor model (2).

For time series observations of R_t and F_t , $t = 1, \dots, T$, the FM two-pass methodology estimates the risk premia using a least squares time-series regression estimate of β , $\hat{\beta} = \sum_{t=1}^T \bar{R}_t \bar{F}_t' \left[\sum_{j=1}^T \bar{F}_j \bar{F}_j' \right]^{-1}$, with $\bar{R}_t = R_t - \bar{R}$, $\bar{R} = \frac{1}{T} \sum_{t=1}^T R_t$. Given $\hat{\beta}$, estimators for λ_1 and λ_k are obtained by regressing the average value of R_t , \bar{R} , on $(\iota_n \vdots \hat{\beta})$,

$$\begin{pmatrix} \hat{\lambda}_1 \\ \hat{\lambda}_F \end{pmatrix} = \left[(\iota_n \vdots \hat{\beta})' (\iota_n \vdots \hat{\beta}) \right]^{-1} (\iota_n \vdots \hat{\beta})' \bar{R}. \quad (6)$$

Lemma 1. *When Assumption 1 holds, the average returns, \bar{R} , the least squares estimator of the assets β 's, $\hat{\beta}$, and the average factors, \bar{F} , converge in distribution as described by*

$$\sqrt{T} \begin{pmatrix} \bar{R} - \iota_n \lambda_1 - \beta \lambda_F \\ \text{vec}(\hat{\beta} - \beta) \\ \bar{F} - \mu_f \end{pmatrix} \xrightarrow{d} \begin{pmatrix} \psi_R \\ \psi_\beta \\ \psi_F \end{pmatrix}, \quad (7)$$

where ψ_R , ψ_β and ψ_F are $n \times 1$, $nk \times 1$ and $k \times 1$ independent random vectors that have normal distributions with mean zero and covariance matrices, Ω , $V_{FF}^{-1} \otimes \Omega$ and V_{FF} .

Proof. see Appendix B. ■

Lemma 1 corresponds with Lemma 1 of Shanken (1992). We use Lemma 1 to derive the properties of the FM risk premia estimator.

When the β 's have size-able values, the FM risk premia estimator has a normal limiting distribution so Wald t -tests perform adequately. Kan and Zhang (1999) show that the large sample distribution of the FM risk premia estimator is notably different when the β 's are equal to zero and $E(\bar{R})$ equals some $n \times 1$ vector of constants c . The behavior of the FM risk premia estimator differs in other cases as well. It results from the dependence of the FM risk premia estimator on the β 's. This dependence is similar to the dependence of the two stage least squares estimator on the quality of the instruments in econometrics. There is an extensive econometrics literature that discusses the behavior of the two stage least squares estimator under limit sequences where the quality of the instruments depends on the sample size, see *e.g.* Phillips (1989), Bekker (1994) and Staiger and Stock (1997). Theorem 1 states the behavior of the FM risk premia estimator in three similar cases for the β 's. The first case corresponds with Kan and Zhang (1999) and has a zero value of the β -vector. The second case has relatively small but non-zero elements of the β -vector. We therefore specify the β 's such that they decrease with the sample size which is analogous to the weak instruments of Staiger and Stock (1999) so we refer to this case as weak β 's. The third case has all β 's equal to zero except for a fixed fraction of the assets.

Theorem 1. *When Assumptions 1 and 2 hold, the behavior of the FM risk premia estimator (6) in large samples is characterized by:*

1. *When $\beta = 0$ and*

(a) $E(\bar{R}) = \iota_n \lambda_1$:

$$\begin{pmatrix} \sqrt{T}(\hat{\lambda}_1 - \lambda_1) \\ \hat{\lambda}_F \end{pmatrix} \xrightarrow{d} \begin{pmatrix} \frac{1}{\sqrt{T}} \frac{1}{a} b' \\ [\Psi'_\beta M_{\iota_n} \Psi_\beta]^{-1} \Psi'_\beta M_{\iota_n} \end{pmatrix} \psi_R, \quad (8)$$

and the order of finite moments of the limit behavior of $\hat{\lambda}_F$ is equal to $n - k$ while the mean of the limit behavior of $\hat{\lambda}_F$ equals zero when $n > k$.

(b) $E(\bar{R}) = c$:

$$\begin{pmatrix} \sqrt{T}(\hat{\lambda}_1 - \frac{1}{a} b' c) \\ \hat{\lambda}_F - \sqrt{T} [\Psi'_\beta M_{\iota_n} \Psi_\beta]^{-1} \Psi'_\beta M_{\iota_n} c \end{pmatrix} \xrightarrow{d} \begin{pmatrix} \frac{1}{\sqrt{T}} \frac{1}{a} b' \\ [\Psi'_\beta M_{\iota_n} \Psi_\beta]^{-1} \Psi'_\beta M_{\iota_n} \end{pmatrix} \psi_R \quad (9)$$

where $\Psi_\beta = \text{vecinv}_k(\psi_\beta)$, $a = \iota'_n M_{\Psi_\beta} \iota_n$, $b = M_{\Psi_\beta} \iota_n$.

2. *When β has a weak value such that $\beta = \beta_T = \frac{1}{\sqrt{T}} B$, with B a fixed full rank $n \times k$ matrix, and*

(a) $E(\bar{R}) = \iota_n \lambda_1 + \beta \lambda_F$:

$$\begin{aligned} & \begin{pmatrix} \sqrt{T}(\hat{\lambda}_1 - \lambda_1) \\ \hat{\lambda}_F - \lambda_F \end{pmatrix} \xrightarrow{d} \\ & \begin{pmatrix} \frac{1}{\sqrt{T}} \frac{1}{a} b' \\ [(B + \Psi_\beta)' M_{\iota_n} (B + \Psi_\beta)]^{-1} (B + \Psi_\beta)' M_{\iota_n} \end{pmatrix} (\psi_R - \Psi_\beta \lambda_F). \end{aligned} \quad (10)$$

and the order of finite moments of the limit behavior of $\hat{\lambda}_F$ is equal to $n - k$.

(b) $E(\bar{R}) = c$:

$$\begin{aligned} & \begin{pmatrix} \sqrt{T}(\hat{\lambda}_1 - \frac{1}{a} b' c) \\ \hat{\lambda}_F - \sqrt{T} [(B + \Psi_\beta)' M_{\iota_n} (B + \Psi_\beta)]^{-1} (B + \Psi_\beta)' M_{\iota_n} c \end{pmatrix} \\ & \xrightarrow{d} \begin{pmatrix} \frac{1}{\sqrt{T}} \frac{1}{a} b' \\ [(B + \Psi_\beta)' M_{\iota_n} (B + \Psi_\beta)]^{-1} (B + \Psi_\beta)' M_{\iota_n} \end{pmatrix} \psi_R, \end{aligned} \quad (11)$$

where $\Psi_\beta = \text{vecinv}_k(\psi_\beta)$, $a = \iota_n' M_{B + \Psi_\beta} \iota_n$, $b = M_{B + \Psi_\beta} \iota_n$.

3. When $\beta = \begin{pmatrix} \beta_1 \\ 0 \end{pmatrix}$, with $\beta_1 : n_1 \times k$, n_1 is fixed and $\frac{n}{T} \rightarrow c$, the bias of $\hat{\lambda}_F$ equals

$$(\beta_1' \beta_1 + W)^{-1} W (\mu_F - \lambda_F), \quad (12)$$

with $W = [\lim_{n \rightarrow \infty} \text{tr}(\frac{1}{n} \Omega)] V_{FF}^{-1}$.

Proof. see Appendix C. ■

Theorem 1 shows that the large sample distribution of the FM risk premia estimator differs substantially from normality in case of zero, weak or many β 's. First, Theorem 1 shows that zero or weak values of β alter the convergence rate of the risk premia estimator. When the moment conditions apply, so $E(\bar{R}) = \iota_n \lambda_1$ (1a) or $E(\bar{R}) = \iota_n \lambda_1 + \beta \lambda_F$ (2a), Theorem 1 shows that the risk premia estimator $\hat{\lambda}_F$ converges to a random variable instead of the convergence at rate \sqrt{T} to the true value λ_F in case of size-able values of β . When the moment conditions do not apply, so $E(\bar{R}) = c$ in (1b) and (2b), the risk premia estimator $\hat{\lambda}_F$ diverges at rate \sqrt{T} both in case of zero or weak β 's. This extends Kan and Zhang (1999) who show that the risk premia estimator $\hat{\lambda}_F$ diverges when β equals zero and $E(\bar{R}) = c$ which corresponds with (1b). Second, Theorem 1 shows that the random variable where the risk premia estimator $\hat{\lambda}_F$ converges to depends on the number of assets since the number of finite moments of this random variable depends on the number of assets. Hence, when we add assets with zero β 's, the density of the risk premia estimator becomes more concentrated. The dependence on the number of assets is further indicated in the third case which shows that the FM risk premia estimator is biased towards μ_F when the number of assets is large and the β 's are only non-zero for a relatively small number of assets.

Although we do not observe the stylized settings from Theorem 1 in practice, Theorem 1 shows that the standard limit theory, that is used to conduct tests on the risk premia using the FM estimator, becomes unreliable when the β 's are rather small and/or only a fraction of the β 's differs from zero. To illustrate the consequences of Theorem 1, we compute the empirical distribution of the risk premia estimator $\hat{\lambda}_F$ for some simulated data-sets. We therefore simulate data from the model,

$$R_t = c + \beta F_t + \varepsilon_t, \quad t = 1, \dots, T, \quad (13)$$

with $E(F_t) = 0$ and where the ε_t 's are independent realizations of $N(0, \Omega)$ distributed random variables. The parameter settings that we use for the simulated data-sets are obtained from Jagannathan and Wang

(1996). Unlike other factor models like those in, for example, Fama and French (1992,1993,1996), the factor model in Jagannathan and Wang (1996) involves factors that do not consist of portfolio returns, *i.e.* the yield premium and labor income growth. We can therefore estimate the risk premia on these factors which explains why we use Jagannathan and Wang (1996) to illustrate the convergence issues with the FM estimator and test statistics.

Panel 1: Shapes of densities of $\hat{\lambda}_F$ for small or zero values of β .

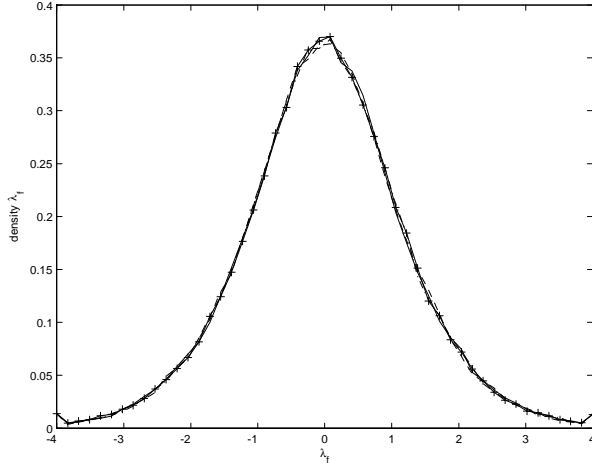


Figure 1.1: Density of $\hat{\lambda}_F$ when $\beta = 0$, $\lambda_1 = 1.24$ and $T = 330$ (solid line), 990 (dashed), 1880 (dashed-dotted), 3300 (solid with plusses)

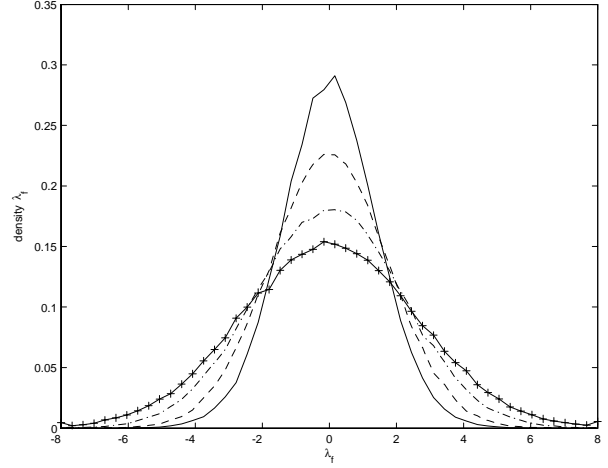


Figure 1.2: Density of $\hat{\lambda}_F$ when $\beta = 0$, $\lambda_1 = 1.24$ and $E(\bar{R}) = c \sim N(0, 0.25^2 I_n)$, $T = 330$ (solid line), 990 (dashed), 1880 (dashed-dotted), 3300 (solid with plusses)

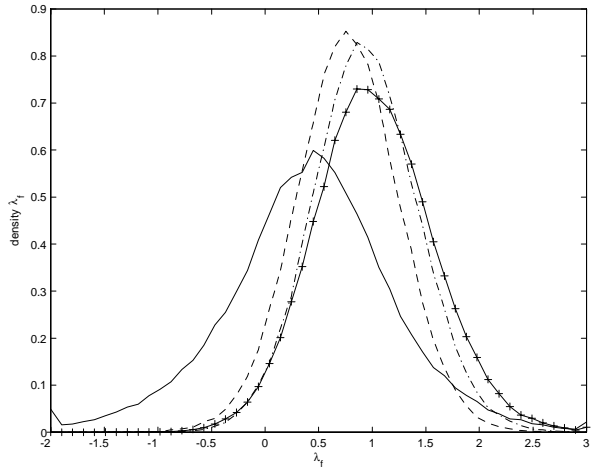


Figure 1.3: Density of $\hat{\lambda}_F$ with $\lambda_F = 1$, $\lambda_1 = 1.24$, $\beta = (0.3 \ 0 \ \dots \ 0)'$ and $n = 2$ (solid with plusses), 10 (dashed-dotted), 24 (dashed), 100 (solidline)

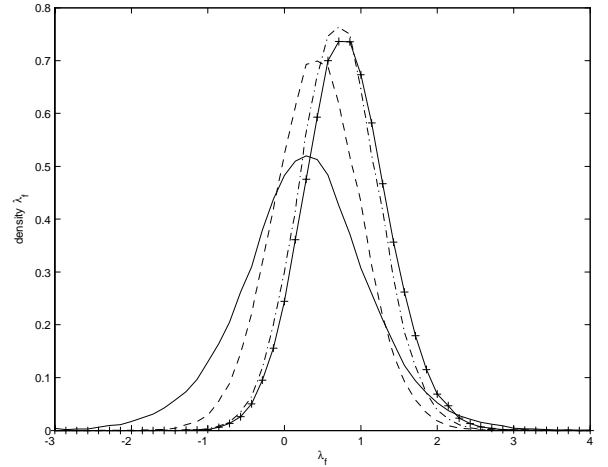


Figure 1.4: Density $\hat{\lambda}_F$ with $E(\bar{R}) = (t_n \lambda_1 + \beta, +c)$ with $\lambda_1 = 1.24$, $\beta = (0.3 \ 0 \ \dots \ 0)'$, $c \sim N(0.04 I_n)$, and $n = 2$ (solid with plusses), 10 (dashed-dotted), 24 (dashed), 100 (solidline)

The data-set of Jagannathan and Wang (1996) consists of monthly observations from July 1963-December 1990 of the return on hundred size and β sorted portfolios so $T = 330$. In this simulation experiment, the (demeaned) return on the value weighted portfolio is the only factor. The covariance matrix Ω results from regressing the size and β -sorted portfolio returns on a constant and the return on the value weighted portfolio. Figures 1.1-1.4 in Panel 1 show the different elements of Theorem 1. The

value of λ_1 that is used for the simulated data-sets results from the FM-estimator (6) when applied to the data from Jagannathan and Wang (1996).

Figure 1.1 contains the sampling density of $\hat{\lambda}_F$ for different values of the sample size T . It illustrates Case 1a of Theorem 1. Figure 1.1 shows that $\hat{\lambda}_F$ converges to a random variable when $\beta = 0$, the model is correctly specified and T converges to infinity since it reflects no convergence of the sampling distribution towards a point mass. The density of $\hat{\lambda}_F$ is centered around zero because the mean of $\hat{\lambda}_F$ is equal to zero as indicated by Theorem 1. Since β is equal to zero, λ_F is not identified so the density of $\hat{\lambda}_F$ is centered around a value of λ_F that contains no information about λ_F .

Figure 1.2 illustrates the divergence of $\hat{\lambda}_F$ as indicated by Case 1b of Theorem 1. In Figure 1.2, $\beta = 0$ and $E(\bar{R})$ equals a realization of a normal distributed random variable with mean zero and covariance matrix $(0.25)^2 I_n$. Figure 1.2 shows that the variance of $\hat{\lambda}_F$ rises when the sample size increases. Hence, $\hat{\lambda}_F$ converges to (plus or minus) infinity when the sample size converges to infinity. Figure 1.2 coincides with Kan and Zhang (1999).

Figure 1.3 shows the influence of adding assets whose β 's equal zero to a factor model whose assets have rather small values of β . The β 's are therefore specified such that only the first one is small and non-zero, $\beta_1 = 0.3$. All other elements of β are equal to zero, $\beta_2 = \dots = \beta_n = 0$. Figure 1.3 shows the sampling density of $\hat{\lambda}_F$ for various values of n while only the β of the first asset is unequal to zero. Figure 1.3 thus corresponds with Case 3 of Theorem 1. For small values of n , $n = 2$, the density of $\hat{\lambda}_F$ is centered around the true value of λ_F that is equal to one. When we increase the number of assets n by adding assets with a zero value of β , the density of $\hat{\lambda}_F$ becomes centered around zero which equals μ_F . Figure 1.3 illustrates the issue of estimating a risk premia for a factor that matters only for a fraction of the assets. Figure 1.3 shows that this estimate becomes more biased when we add assets for which the factor is irrelevant.

Figure 1.4 shows the consequences of misspecification for the parameter setting that is used in Figure 1.3. We therefore added some misspecification to the $E(\bar{R})$, $E(\bar{R}) = (\iota_n \lambda_1 + \beta + c)$, where c is a realization of a $N(0, 0.04 I_n)$ distributed random vector. All remaining parameter values are identical to the ones that are used for Figure 1.3. Figure 1.4 shows that the density of $\hat{\lambda}_F$ moves towards zero and that its' variance increases when we add assets with zero β 's and the model is misspecified. Figure 1.4 reveals a combined effect of the phenomena present in Figures 1.2 and 1.3.

The odd behavior of the FM risk premia estimator for small or zero values of the β 's holds for other risk premia estimators as well. The behavior of the GLS risk premia estimator,

$$\begin{pmatrix} \tilde{\lambda}_1 \\ \tilde{\lambda}_F \end{pmatrix} = \left[(\iota_n \vdots \hat{\beta})' \hat{\Omega}^{-1} (\iota_n \vdots \hat{\beta}) \right]^{-1} (\iota_n \vdots \hat{\beta})' \hat{\Omega}^{-1} \bar{R}, \quad (14)$$

is qualitatively similar to the behavior of the FM risk premia estimator for small or zero values of the β 's. Theorem 2 states the behavior of the GLS risk premia estimator (14) for such cases.

Theorem 2. *When Assumptions 1 and 2 hold, the behavior of the GLS risk premia estimator (14) in large samples in case of zero or weak β 's is characterized by:*

1. When $\beta = 0$ and

(a) $E(\bar{R}) = \iota_n \lambda_1 :$

$$\begin{pmatrix} \sqrt{T}(\hat{\lambda}_1 - \lambda_1) \\ \hat{\lambda}_F \end{pmatrix} \xrightarrow{d} \begin{pmatrix} \frac{1}{\sqrt{T}} \frac{1}{a} b' \\ (\Psi'_\beta \Omega^{-\frac{1}{2}} M_{\Omega^{-\frac{1}{2}} \iota_n} \Omega^{-\frac{1}{2}} \Psi_\beta)^{-1} \Psi'_\beta \Omega^{-\frac{1}{2}} M_{\Omega^{-\frac{1}{2}} \iota_n} \Omega^{-\frac{1}{2}} \end{pmatrix} \psi_R, \quad (15)$$

so $\hat{\lambda}_F$ converges to a multi-variate t distributed random variable with location 0, scale parameter $Q_{11.F}Q_{FF}^{-1}$ and $n-k$ degrees of freedom and the mean of the limit behavior of $\hat{\lambda}_F$ equals zero when $n > k$.

(b) $E(\bar{R}) = c$:

$$\begin{aligned} & \left(\begin{array}{c} \sqrt{T}(\hat{\lambda}_1 - \frac{1}{a}b'\Omega^{-\frac{1}{2}}c) \\ \hat{\lambda}_F - \sqrt{T}(\Psi'_\beta\Omega^{-\frac{1}{2}}M_{\Omega^{-\frac{1}{2}}\iota_n} \Omega^{-\frac{1}{2}}\Psi_\beta)^{-1}\Psi'_\beta\Omega^{-\frac{1}{2}}M_{\Omega^{-\frac{1}{2}}\iota_n} \Omega^{-\frac{1}{2}}c \end{array} \right) \xrightarrow{d} \\ & \left(\begin{array}{c} \frac{1}{\sqrt{T}}\frac{1}{a}b' \\ (\Psi'_\beta\Omega^{-\frac{1}{2}}M_{\Omega^{-\frac{1}{2}}\iota_n} \Omega^{-\frac{1}{2}}\Psi_\beta)^{-1}\Psi'_\beta\Omega^{-\frac{1}{2}}M_{\Omega^{-\frac{1}{2}}\iota_n} \end{array} \right) \Omega^{-\frac{1}{2}}\psi_R \end{aligned} \quad (16)$$

where $\Psi_\beta = \text{vecinv}_k(\psi_\beta)$, $a = \iota'_n\Omega^{-\frac{1}{2}}M_{\Omega^{-\frac{1}{2}}\Psi_\beta} \Omega^{-\frac{1}{2}}\iota_n$, $b = M_{\Omega^{-\frac{1}{2}}\Psi_\beta} \Omega^{-\frac{1}{2}}\iota_n$, $Q_{11.F} = Q_{11} - Q_{1F}Q_{FF}^{-1}Q_{F1}$.

2. When β has a weak value such that $\beta = \frac{1}{\sqrt{T}}B$ and

(a) $E(\bar{R}) = \iota_n\lambda_1 + \frac{1}{\sqrt{T}}B\lambda_F$:

$$\begin{aligned} & \left(\begin{array}{c} \sqrt{T}(\hat{\lambda}_1 - \lambda_1) \\ \hat{\lambda}_F - \lambda_F \end{array} \right) \xrightarrow{d} \left(\begin{array}{c} \frac{1}{\sqrt{T}}\frac{1}{a}b' \\ [(B + \Psi_\beta)'\Omega^{-\frac{1}{2}}M_{\Omega^{-\frac{1}{2}}\iota_n} \Omega^{-\frac{1}{2}}(B + \Psi_\beta)]^{-1}(B + \Psi_\beta)'\Omega^{-\frac{1}{2}}M_{\Omega^{-\frac{1}{2}}\iota_n} \\ \Omega^{-\frac{1}{2}}(\psi_{R,\beta} + \Psi_\beta(\rho - \lambda_F)) \end{array} \right) \end{aligned} \quad (17)$$

and the order of finite moments of the limit behavior of $\hat{\lambda}_F - \rho$ is equal to $n - k$.

(b) $E(\bar{R}) = c$:

$$\begin{aligned} & \left(\begin{array}{c} \sqrt{T}(\hat{\lambda}_1 - \frac{1}{a}b'\Omega^{-\frac{1}{2}}c) \\ \hat{\lambda}_F - \sqrt{T}[(B + \Psi_\beta)'\Omega^{-\frac{1}{2}}M_{\Omega^{-\frac{1}{2}}\iota_n} \Omega^{-\frac{1}{2}}(B + \Psi_\beta)]^{-1}(B + \Psi_\beta)'\Omega^{-\frac{1}{2}}M_{\Omega^{-\frac{1}{2}}\iota_n} \Omega^{-\frac{1}{2}}c \end{array} \right) \\ & \xrightarrow{d} \left(\begin{array}{c} \frac{1}{\sqrt{T}}\frac{1}{a}b' \\ [(B + \Psi_\beta)'\Omega^{-\frac{1}{2}}M_{\Omega^{-\frac{1}{2}}\iota_n} \Omega^{-\frac{1}{2}}(B + \Psi_\beta)]^{-1}(B + \Psi_\beta)'\Omega^{-\frac{1}{2}}M_{\Omega^{-\frac{1}{2}}\iota_n} \end{array} \right) \Omega^{-\frac{1}{2}}\psi_R, \end{aligned} \quad (18)$$

where $\Psi_\beta = \text{vecinv}_k(\psi_\beta)$, $a = \iota'_n\Omega^{-\frac{1}{2}}M_{\Omega^{-\frac{1}{2}}(B+\Psi_\beta)} \Omega^{-\frac{1}{2}}\iota_n$, $b = M_{\Omega^{-\frac{1}{2}}(B+\Psi_\beta)} \Omega^{-\frac{1}{2}}\iota_n$.

Proof. see Appendix D. ■

Theorems 1 and 2 show that the dependence of the risk premia estimators on the β 's makes statistical inference that is based on standard Wald t statistics unreliable when the β 's are small. This emphasizes the need for statistics that remain reliable when the β 's are small.

3 Tests of the risk premium

Theorems 1 and 2 show that the usual risk premia estimators converge to random variables for small and zero values of the β 's when the sample size gets large. The convergence of Wald and t -statistics to test hypotheses on the risk premia is therefore non-standard for small or zero values of the β 's. To overcome this problem, we propose some alternative statistics to test a hypothesis on the risk premia, like, for example, $H_0 : \lambda_F = \lambda_{F,0}$. Since our interest is on the risk premia λ_F , we remove λ_1 from the

model by removing the return on the n -th asset and taking all other asset returns in deviation from the return on the n -th asset.² The moment conditions then become

$$\begin{aligned} E(\mathcal{R}_t) &= \mathcal{B}\lambda_F \\ \text{cov}(\mathcal{R}_t, F_t) &= \mathcal{B}\text{var}(F_t) \\ E(F_t) &= \mu_F, \end{aligned} \quad (19)$$

with $\mathcal{R}_t = R_{1t} - \iota_{n-1}R_{nt}$ and $\mathcal{B} = \beta_1 - \iota_{n-1}\beta_n$, for $R_t = (R'_{1t} \ : \ R'_{nt})'$, $\beta = (\beta'_1 \ : \ \beta'_n)'$; $R_{1t} : (n-1) \times 1$, $R_{nt} : 1 \times 1$, $\beta_1 : (n-1) \times k$, $\beta_n : 1 \times k$. Under $H_0 : \lambda_F = \lambda_{F,0}$, a least squares estimator for \mathcal{B} is given by

$$\tilde{\mathcal{B}} = \sum_{t=1}^T \mathcal{R}_t(\bar{F}_t + \lambda_{F,0}) \left[\sum_{j=1}^T (\bar{F}_j + \lambda_{F,0})(\bar{F}_j + \lambda_{F,0})' \right]^{-1}. \quad (20)$$

We use Assumption 1 to determine the joint behavior of the average returns and the least squares estimator $\tilde{\mathcal{B}}$ under $H_0 : \lambda_F = \lambda_{F,0}$ in large samples.

Lemma 2. *Under $H_0 : \lambda_F = \lambda_{F,0}$ and when Assumption 1 holds,*

$$\sqrt{T} \begin{pmatrix} \bar{\mathcal{R}} - \tilde{\mathcal{B}}\lambda_{F,0} \\ \text{vec}(\tilde{\mathcal{B}} - \mathcal{B}) \\ \bar{F} - \mu_F \end{pmatrix} \xrightarrow{d} \begin{pmatrix} \psi_{\mathcal{R}} \\ \psi_{\mathcal{B}} \\ \psi_F \end{pmatrix},$$

where $\psi_{\mathcal{R}}$, $\psi_{\mathcal{B}}$ and ψ_F are independent $(n-1) \times 1$, $(n-1)k \times 1$ and $k \times 1$ normal distributed random vectors with mean zero and covariance matrices $(1 - \lambda'_{F,0}\mathcal{Q}(\lambda_F)_{FF}\lambda_{F,0}) \otimes \Sigma$, $\mathcal{Q}(\lambda_F)_{FF} \otimes \Sigma$ and V_{FF} , with $\mathcal{Q}(\lambda_F)_{FF} = \left[\lim_{T \rightarrow \infty} E \left(\frac{1}{T} \sum_{t=1}^T (\bar{F}_t + \lambda_F)(\bar{F}_t + \lambda_F)' \right) \right]^{-1}$ and $\Sigma = \Omega_{11} - \iota_{n-1}\omega_{1n} - \omega_{n1}\iota'_{n-1} + \iota_{n-1}\omega_{nn}\iota'_{n-1}$ for $\Omega = \begin{pmatrix} \Omega_{11} & \omega_{1n} \\ \omega_{n1} & \omega_{nn} \end{pmatrix}$, $\Omega_{11} : (n-1) \times (n-1)$, $\omega_{1n} = \omega'_{n1} : 1 \times (n-1)$ and $\omega_{nn} : 1 \times 1$.

Proof. see Appendix E. ■

The Wald statistic that is based on the FM risk premia estimator to test $H_0 : \lambda_F = \lambda_{F,0}$ reads

$$\text{FM-W}(\lambda_{F,0}) = (\hat{\lambda}_F - \lambda_{F,0})' \text{var}(\hat{\lambda}_F)^{-1} (\hat{\lambda}_F - \lambda_{F,0}). \quad (21)$$

For size-able full rank values of β and under $H_0 : \lambda_F = \lambda_{F,0}$, $\text{FM-W}(\lambda_{F,0})$ converges in distribution to a $\chi^2(k)$ distributed random variable when the sample size gets large. Theorem 1 shows that $\hat{\lambda}_F$ converges to a non-normal distributed random variable that is centered around zero for weak and zero values of the β 's. The Wald statistic $\text{W}_{\text{FM}}(\lambda_{F,0})$ does therefore not converge to a $\chi^2(m)$ distributed random variable for such values of the β 's. It implies that the Wald statistic is unreliable for conducting inference in these cases. To indicate a solution to this problem, we note that the Wald statistic that tests H_0 can be specified as

$$\text{FM-W}(\lambda_{F,0}) = (\bar{R} - \iota_N \hat{\lambda}_1 - \hat{\beta}\lambda_{F,0})' \hat{\beta} \hat{\Theta}^{-1} \hat{\beta}' (\bar{R} - \iota_N \hat{\lambda}_1 - \hat{\beta}\lambda_{F,0}), \quad (22)$$

with $\hat{\Theta} = (\hat{\beta}' \hat{\Omega} \hat{\beta} - \hat{\beta} \hat{\Omega} \iota_n (\iota'_n \hat{\Omega} \iota_n)^{-1} \iota'_n \hat{\Omega} \hat{\beta}) / (1 + \hat{\lambda}'_F (\frac{1}{T} \sum_{t=1}^T \bar{F}_t \bar{F}'_t)^{-1} \hat{\lambda}_F)$ and where $\hat{\Omega}$ is a least squares covariance matrix estimator of Ω , $\hat{\Omega} = \frac{1}{T-k-1} \sum_{t=1}^T (\bar{R}_t - \hat{\beta} \bar{F}_t)(\bar{R}_t - \hat{\beta} \bar{F}_t)'$, see Shanken (1992). Lemma 1 shows that $\bar{R} - \iota_N \hat{\lambda}_1 - \beta \lambda_{F,0}$ and $\hat{\beta}$ are independently distributed in large samples. Because of this dependence, $\bar{R} - \iota_N \hat{\lambda}_1 - \hat{\beta} \lambda_{F,0}$, is, however, not independent of $\hat{\beta}$ in large samples which explains the

²The results are invariant with respect to the asset return that is dropped and with respect to which all assets returns are taken in deviation from.

convergence issues with the Wald statistic when the sample size gets large and the β 's are relatively small.

Lemma 2 shows that $\bar{\mathcal{R}} - \tilde{\mathcal{B}}\lambda_{F,0}$ and $\tilde{\mathcal{B}}$ are independently distributed when the sample size gets large. Statistics that are based on $\bar{\mathcal{R}} - \tilde{\mathcal{B}}\lambda_{F,0}$ and/or $\tilde{\mathcal{B}}$ therefore possess better convergence properties than the Wald-statistic. We propose a few of these statistics which are based on the GMM and instrumental variable statistics that are proposed in Kleibergen (2002, 2005a) and Moreira (2003). Since the factor model is not accomodated by the GMM framework from Kleibergen (2005a), the statistics do not result directly from the GMM and instrumental variable statistics.

Fama-MacBeth LM statistic The independence of $\bar{\mathcal{R}} - \tilde{\mathcal{B}}\lambda_{F,0}$ and $\tilde{\mathcal{B}}$ in large samples implies that we can base a statistic with convenient properties on their product.

Theorem 3. *Under $H_0 : \lambda_F = \lambda_{F,0}$ and when Assumption 1 holds,*

$$\text{FM-LM}(\lambda_{F,0}) = \frac{T}{1 - \lambda'_{F,0} \hat{\mathcal{Q}}(\lambda_F)_{FF}^{-1} \lambda_{F,0}} (\bar{\mathcal{R}} - \tilde{\mathcal{B}}\lambda_{F,0})' \tilde{\mathcal{B}} (\tilde{\mathcal{B}}' \tilde{\Sigma} \tilde{\mathcal{B}})^{-1} \tilde{\mathcal{B}}' (\bar{\mathcal{R}} - \tilde{\mathcal{B}}\lambda_{F,0}), \quad (23)$$

where $\hat{\mathcal{Q}}(\lambda_F) = \frac{1}{T} \sum_{t=1}^T (\bar{F}_t + \lambda_F)(\bar{F}_t + \lambda_F)'$ and $\tilde{\Sigma} = \frac{1}{T-k} \sum_{t=1}^T (\bar{\mathcal{R}}_t - \tilde{\mathcal{B}}(\bar{F}_t + \lambda_{F,0}))(\bar{\mathcal{R}}_t - \tilde{\mathcal{B}}(\bar{F}_t + \lambda_{F,0}))'$, converges to a $\chi^2(m)$ distributed random variable when the sample size gets large for all values of the β 's.

Proof. results from the independence of $\bar{\mathcal{R}} - \tilde{\mathcal{B}}\lambda_{F,0}$ and $\tilde{\mathcal{B}}$ shown in Lemma 2 and that $\hat{\mathcal{Q}}(\lambda_F) \xrightarrow{p} \mathcal{Q}(\lambda_F)$ and $\tilde{\Sigma} \xrightarrow{p} \Sigma$. ■

The FM-Lagrange multiplier (LM) statistic is based on the FM-Wald statistic in (22) which explains why we refer to it as a Fama-MacBeth LM statistic. Compared with the FM-Wald statistic (22), the primary differences are the estimators of the β 's, $\tilde{\mathcal{B}}$ instead of $\hat{\beta}$, and that $\hat{\lambda}_1$ is no longer present since it has been removed by taking the returns in deviation from the returns on the n -th asset. Identical to the FM risk premia estimator, FM-LM($\lambda_{F,0}$) is not invariant to transformations of the asset returns so it differs in this respect from the standard LM statistics which are invariant to transformations.

GLS-LM statistic The GLS risk premia estimator (14) is invariant to transformations of the asset returns so we obtain an invariant LM statistic by incorporating the inverse of the covariance matrix in (23).

Theorem 4. *Under $H_0 : \lambda_F = \lambda_{F,0}$ and when Assumption 1 holds,*

$$\text{GLS-LM}(\lambda_{F,0}) = \frac{T}{1 - \lambda'_{F,0} \hat{\mathcal{Q}}(\lambda_F)_{FF}^{-1} \lambda_{F,0}} (\bar{\mathcal{R}} - \tilde{\mathcal{B}}\lambda_{F,0})' \tilde{\Sigma}^{-1} \tilde{\mathcal{B}} (\tilde{\mathcal{B}}' \tilde{\Sigma}^{-1} \tilde{\mathcal{B}})^{-1} \tilde{\mathcal{B}}' \tilde{\Sigma}^{-1} (\bar{\mathcal{R}} - \tilde{\mathcal{B}}\lambda_{F,0}), \quad (24)$$

converges to a $\chi^2(m)$ distributed random variable when the sample size gets large for all values of the β 's.

Proof. results from the independence of $\bar{\mathcal{R}} - \tilde{\mathcal{B}}\lambda_{F,0}$ and $\tilde{\mathcal{B}}$ shown in Lemma 2. ■

Unlike the FM-LM statistic (23), the GLS-LM statistic (24) is invariant to transformations of the asset returns. The GLS-LM statistic (24) does, however, depend on the inverse of the covariance matrix $\tilde{\Sigma}$ which can be of a large dimension and therefore difficult to obtain. For example, this is a 99×99 matrix in Jagannathan and Wang (1996).

Factor AR statistic When the disturbances in the linear factor model (2) are normally distributed with mean zero and covariance matrix Ω , we can construct the likelihood function of λ_1 , λ_F , β and Ω , see *e.g.* Gibbons (1982). In an identical manner, we can construct the likelihood function for a linear factor model in which the expected asset returns are not restricted to be equal to $\iota_n \lambda_1 + \beta \lambda_F$. After concentrating with respect to λ_1 and β , the difference between the logarithms of the likelihoods of the unrestricted linear factor model and the restricted factor model under $H_0 : \lambda_F = \lambda_{F,0}$ is proportional to³:

$$\text{FAR}(\lambda_{F,0}) = \frac{T}{1 - \lambda'_{F,0} \hat{\mathcal{Q}}(\lambda_F)_{FF}^{-1} \lambda_{F,0}} (\bar{\mathcal{R}} - \tilde{\mathcal{B}} \lambda_{F,0})' \tilde{\Sigma}^{-1} (\bar{\mathcal{R}} - \tilde{\mathcal{B}} \lambda_{F,0}). \quad (25)$$

We refer to this statistic as the factor AR statistic (FAR) since it is similar to the Anderson-Rubin statistic, see Anderson and Rubin (1949), in the instrumental variables regression model. The FAR statistic is proportional to the square of the Hansen-Jagannathan distance evaluated at $\lambda_{F,0}$ when we use $\tilde{\Sigma}$ as the covariance matrix estimator, see Hansen and Jagannathan (1997)

Theorem 5. *Under $H_0 : \lambda_F = \lambda_{F,0}$ and when Assumption 1 holds, the Factor AR statistic (25) converges to a $\chi^2(n-1)$ distributed random variable when the sample size gets large for all values of the β 's.*

Proof. results from the large sample behavior of $\bar{\mathcal{R}} - \tilde{\mathcal{B}} \lambda_{F,0}$ stated in Lemma 2. ■

Because of the relationship between the FAR statistic (25) and the likelihood function, minimizing the FAR statistic over $\lambda_{F,0}$ is identical to maximizing the likelihood. Hence, the maximum likelihood estimator for $\lambda_{F,0}$ is obtained by minimizing $\text{FAR}(\lambda_{F,0})$ over $\lambda_{F,0}$.

Misspecification factor pricing tests The GLS-LM statistic (24) is a quadratic form of the derivative of $\text{FAR}(\lambda_{F,0})$ since⁴

$$\frac{\partial}{\partial \lambda_{F,0}} \text{FAR}(\lambda_{F,0}) = c (\bar{\mathcal{R}} - \tilde{\mathcal{B}} \lambda_{F,0})' \tilde{\Sigma}^{-1} \tilde{\mathcal{B}}, \quad (26)$$

with $c = -2 \frac{T}{1 - \lambda'_{F,0} \hat{\mathcal{Q}}(\lambda_F)_{FF}^{-1} \lambda_{F,0}} \left[1 + \frac{1}{T-k-1} \text{FAR}(\lambda_{F,0}) \right]$. The GLS-LM statistic is therefore equal to zero at the maximum likelihood estimator of λ_F from Gibbons (1982) which also minimizes $\text{FAR}(\lambda_{F,0})$. Besides the maximum likelihood estimator, the GLS-LM statistic is also equal to zero at other values of $\lambda_{F,0}$ which set the derivative of the likelihood to zero, like, for example, local maxima and inflexion points. This hampers inference using the GLS-LM statistic.

To overcome the difficulty with the GLS-LM statistic, it is convenient to conduct a pre-test that verifies the validity of the moment conditions $E(\mathcal{R}_t) = \mathcal{B} \lambda_F$ at $\lambda_F = \lambda_{F,0}$ which are part of the moment equations (19). A statistic that is well suited for this purpose is the factor pricing statistic

$$\text{JGLS}(\lambda_{F,0}) = \text{FAR}(\lambda_{F,0}) - \text{GLS-LM}(\lambda_{F,0}). \quad (27)$$

Theorem 6. *Under $H_0 : \lambda_F = \lambda_{F,0}$ and when Assumption 1 holds, the factor pricing JGLS($\lambda_{F,0}$) statistic (27) converges to a $\chi^2(n-k-1)$ distributed random variable that is independently distributed of the $\chi^2(k)$ distributed random variable where GLS-LM($\lambda_{F,0}$) (24) converges to when the sample size gets large for all possible values of the β 's.*

Proof. results from the large sample behavior of $\bar{\mathcal{R}} - \tilde{\mathcal{B}} \lambda_{F,0}$ documented in Lemma 2. The independence from $\text{GLS-LM}(\lambda_{F,0})$ results since $\text{JGLS}(\lambda_{F,0})$ projects $\tilde{\Sigma}^{-\frac{1}{2}} (\bar{\mathcal{R}} - \tilde{\mathcal{B}} \lambda_{F,0})$ on the space orthogonal to $\tilde{\Sigma}^{-\frac{1}{2}} \tilde{\mathcal{B}}$ where $\text{GLS-LM}(\lambda_{F,0})$ projects it onto. ■

³We refer to Appendix G for the construction of $\text{FAR}(\lambda_{F,0})$.

⁴see Appendix F for a proof.

The factor pricing statistic tests $E(\mathcal{R}_t) = \mathcal{B}\lambda_F$ for a pre-specified value of the risk premia λ_F . The FM-LM statistic can be used as well for this purpose

$$\text{JFM}(\lambda_{F,0}) = \text{FAR}(\lambda_{F,0}) - \text{FM-LM}(\lambda_{F,0}). \quad (28)$$

Theorem 7. Under $H_0 : \lambda_F = \lambda_{F,0}$ and when Assumption 1 holds, the factor pricing JFM($\lambda_{F,0}$) statistic (28) converges to a $\chi^2(n-k-1)$ distributed random variable that is independently distributed of the $\chi^2(k)$ distributed random variable where FM-LM($\lambda_{F,0}$) (23) converges to when the sample size gets large for all possible values of the β 's.

Proof. results from the large sample behavior of $\bar{\mathcal{R}} - \tilde{\mathcal{B}}\lambda_{F,0}$ documented in Lemma 2. The independence from FM-LM($\lambda_{F,0}$) results since JFM($\lambda_{F,0}$) projects $\tilde{\Sigma}^{-\frac{1}{2}}(\bar{\mathcal{R}} - \tilde{\mathcal{B}}\lambda_{F,0})$ on the space orthogonal to $\tilde{\Sigma}^{\frac{1}{2}}\tilde{\mathcal{B}}$ where FM-LM($\lambda_{F,0}$) projects it onto. ■

The factor pricing statistics JGLS($\lambda_{F,0}$) and JFM($\lambda_{F,0}$) resolve the inferential problems with the GLS-LM and FM-LM statistics when we use them as pre-tests before we apply the GLS-LM and FM-LM statistics. A test which tests the adequacy of the moment conditions and $H_0 : \lambda_F = \lambda_{F,0}$ with $(1-\alpha) \times 100\%$ significance then results by applying the JGLS($\lambda_{F,0}$) or JFM($\lambda_{F,0}$) statistics with $(1-\alpha_J) \times 100\%$ significance and in case we do not reject, we apply the GLS-LM($\lambda_{F,0}$) or FM-LM($\lambda_{F,0}$) statistics with $(1-\alpha_{LM}) \times 100\%$ significance, where $(1-\alpha) = (1-\alpha_J)(1-\alpha_{LM})$ so $\alpha = \alpha_J + \alpha_{LM} + \alpha_J\alpha_{LM} \approx \alpha_J + \alpha_{LM}$, see Kleibergen (2005a,b).

Conditional likelihood ratio statistic The likelihood ratio statistic to test $H_0 : \lambda_F = \lambda_{F,0}$ against $H_1 : \lambda_F \neq \lambda_{F,0}$ equals twice the difference between the logarithms of the concentrated likelihoods under H_1 and H_0 :

$$\text{LR}(\lambda_{F,0}) = 2 [\max_{\lambda_F} \ln \mathcal{L}_{\text{res}}(\lambda_F) - \ln \mathcal{L}_{\text{res}}(\lambda_{F,0})]. \quad (29)$$

When $k = 1$, the likelihood ratio statistic can be specified as⁵

$$\text{CLR}(\lambda_{F,0}) = \frac{1}{2} \left[\text{FAR}(\lambda_{F,0}) - r(\lambda_{F,0}) + \sqrt{(\text{FAR}(\lambda_{F,0}) + r(\lambda_{F,0}))^2 - 4r(\lambda_{F,0})\text{JFAR}(\lambda_{F,0})} \right], \quad (30)$$

with $r(\lambda_{F,0}) = \hat{\mathcal{Q}}(\lambda_F)_{FF}^{-1} \tilde{\mathcal{B}}' \tilde{\Sigma}^{-1} \tilde{\mathcal{B}}$, which corresponds with the conditional likelihood (CLR) ratio statistic of Moreira (2003) for the linear instrumental variables regression model.

Theorem 8. Under $H_0 : \lambda_F = \lambda_{F,0}$, $k = 1$, Assumption 1 and given $r(\lambda_{F,0})$, the conditional likelihood ratio statistic CLR($\lambda_{F,0}$) (30) converges to the random variable:

$$\frac{1}{2} \left[\varphi_k + \varphi_{n-k-1} - r(\lambda_{F,0}) + \sqrt{(\varphi_k + \varphi_{n-k-1} + r(\lambda_{F,0}))^2 - 4r(\lambda_{F,0})\varphi_{n-k-1}} \right], \quad (31)$$

where φ_k and φ_{n-k-1} are independent $\chi^2(k)$ and $\chi^2(n-k-1)$ distributed random variables, when the sample size gets large.

Proof. results from the large sample behavior of $\bar{\mathcal{R}} - \tilde{\mathcal{B}}\lambda_{F,0}$ documented in Lemma 2 and that $\text{FAR}(\lambda_{F,0}) \xrightarrow{d} \varphi_k + \varphi_{n-k-1}$, $\text{JFAR}(\lambda_{F,0}) \xrightarrow{d} \varphi_{n-k-1}$. ■

The expression of the large sample distribution of the CLR statistic in Theorem 8 only applies when $k = 1$. It provides a tight upperbound on the large sample distribution of the CLR statistic when $r(\lambda_{F,0})$ is a statistic that tests for the rank of \mathcal{B} using $\tilde{\mathcal{B}}$ when k exceeds one, see Kleibergen (2005b). The conditional distribution of the CLR statistic (31) is straightforward to simulate from and results from simulating independent $\chi^2(k)$ and $\chi^2(n-k-1)$ distributed random variables which are alongside $r(\lambda_{F,0})$ used to compute the expression in (31).

⁵see Appendix G for a derivation of the expression of CLR($\lambda_{F,0}$).

Kan and Zhang Misspecification The factor pricing JGLS and JFM statistics, which test the hypothesis of factor pricing at $H_0 : \lambda_F = \lambda_{F,0}$, and the FAR statistic, which tests the joint hypothesis of factor pricing and H_0 , have power against the misspecification analyzed by Kan and Zhang (1999). The combinations of the JGLS and GLS-LM statistics and JFM and FM-LM statistics, that were proposed to overcome the spurious power declines of the GLS-LM and FM-LM statistics, therefore also provide power against misspecification of the asset returns.

4 Power and Size Comparison

We analyze the size and power of the different statistics in a simulation experiment that is based on Jagannathan and Wang (1996). We therefore generate assets returns from the one factor asset pricing model

$$R_t = \iota_n \lambda_1 + \beta(\bar{F}_t + \lambda_F) + \varepsilon_t, \quad t = 1, \dots, T, \quad (32)$$

with R_t the $n \times 1$ vector of asset returns, β the $n \times 1$ vector of β 's, λ_1 the zero- β return, λ_F the risk premium and the disturbances ε_t are independently normal distributed with mean zero and covariance matrix Ω . We use parameter settings that are obtained from the data used by Jagannathan and Wang (1996) so $n = 100$, $T = 330$, \bar{F}_t is the demeaned return on the value weighted portfolio, λ_1 equals the FM estimate $\hat{\lambda}_1$ and Ω equals the least squares covariance matrix estimator.

We generate asset returns from the linear factor model (32) for three different values of β , 1. $\beta = \hat{\beta}_{VW}$, 2. $\beta = 0.25\hat{\beta}_{VW}$ and 3. $\beta = 0.1\hat{\beta}_{VW}$, with $\hat{\beta}_{VW}$ the least squares estimator for the β 's of the asset returns with respect to the value weighted portfolio. We use a range of values of λ_F and the different statistics that we discussed previously to test $H_0 : \lambda_F = 3$ with 95% significance.

The power curves in Panel 2 reveal the issues involved with the different statistics. The lefthandside of Panel 2, which consists of Figures 2.1, 2.3 and 2.5, displays the power curves of the FM and GLS based Wald and LM statistics. The righthandside of Panel 2, which consists of Figures 2.2, 2.4 and 2.6, displays the power curves of the CLR, FAR and combinations of the JGLS and GLS-LM statistics and the JFM and FM-LM statistics. These combinations are such that a 99% critical value is applied to JGLS and JFM and a 96% critical value to GLS-LM and FM-LM so the overall size of the tests equals 5%.

The power curves of the FM and GLS Wald statistics in Figure 2.1 show that the rejection frequency of the FM Wald statistic is approximately 5% at $\lambda_F = 3$ while the size of the GLS Wald statistic differs considerably from 5%. Hence, the GLS Wald statistic is size distorted. Figures 2.3 and 2.5 show that both Wald statistics are size distorted for smaller values of β while the FM and GLS-LM statistics are not size distorted. These Figures also demonstrate a spurious decline of power of the FM and GLS-LM statistics at values of λ_F which are considerable different from the hypothesized value of three. These power declines result because the FM and GLS LM statistics are proportional to the derivatives of resp. $(\bar{R} - \tilde{\beta}\lambda_F)'(\bar{R} - \tilde{\beta}\lambda_F)$ and the FAR statistic with respect to λ_F . Hence, the FM and GLS-LM statistics are equal to zero at minima, maxima and inflexion points of these statistics which explains the spurious power declines of the FM and GLS-LM statistics.

It is remarkable that the GLS Wald statistics is already size distorted when β equals $\hat{\beta}_{VW}$ in Figure 2.1. This size distortion results to a large extent from the inversion of the 100×100 covariance matrix $\hat{\Omega}$. Appendix H explains why this size distortion does not occur for the FM-GLS and the other statistics which contain the inverse of the 99×99 matrix $\tilde{\Sigma}$.

Panel 2: Power curves of statistics that test $H_0 : \lambda_F = 3$ with 95% significance

$\beta = \hat{\beta}_{VW} :$

Figure 2.1

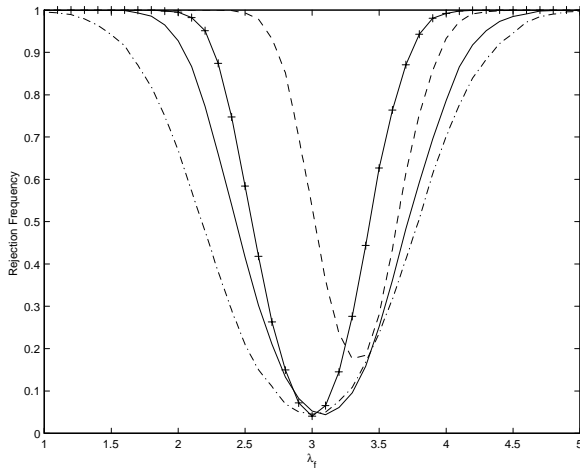
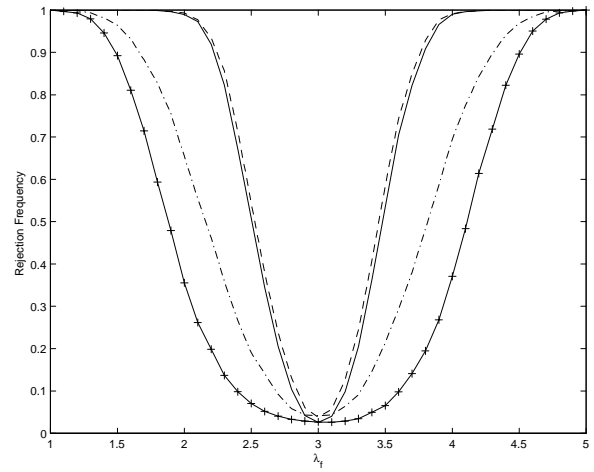


Figure 2.2



$\beta = 0.25\hat{\beta}_{VW} :$

Figure 2.3

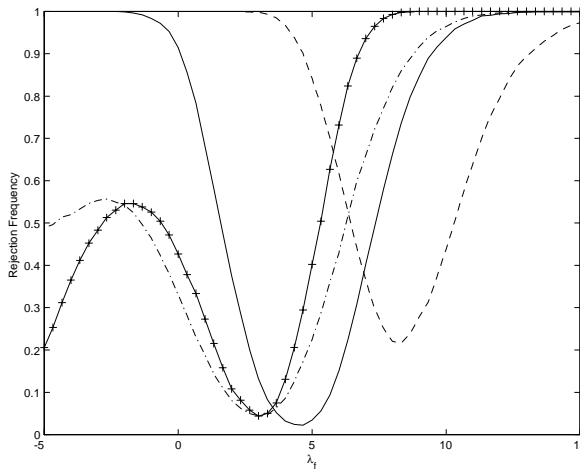
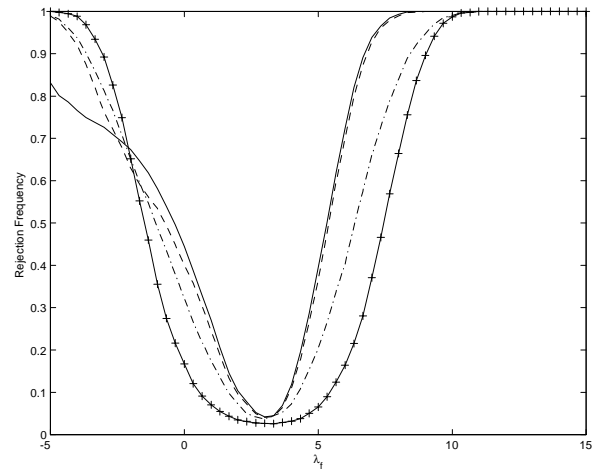


Figure 2.4



$\beta = 0.1\hat{\beta}_{VW} :$

Figure 2.5

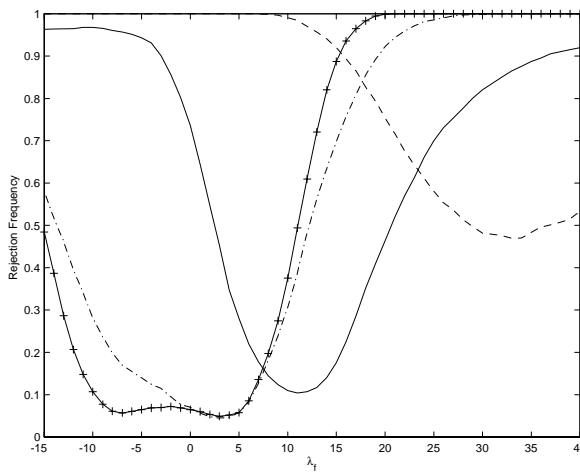
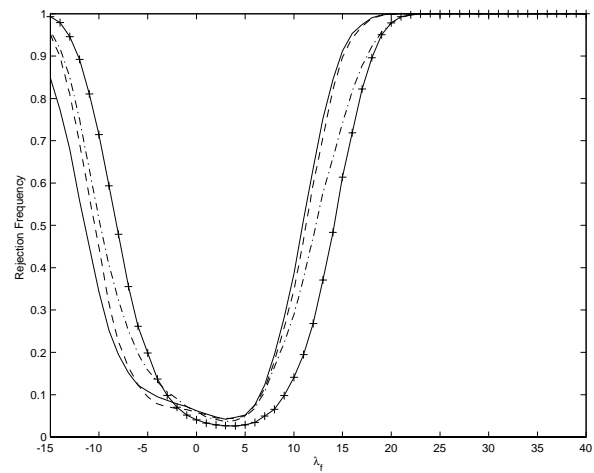


Figure 2.6



Figures 2.1, 2.3, 2.5: Power curves of FM-W (solid), GLS-W (dashed), FM-LM (dashed-dotted) and GLS-LM (solid with plusses) statistics.

Figures 2.2, 2.4, 2.6: Power curves of CLR (solid), combination of JGLS and GLS-LM (dashed), combination of JFM and FM-LM (dashed-dotted) and FAR (solid with plusses) statistics.

Figures 2.2, 2.4 and 2.6 only contain power curves of statistics whose large sample distributions are unaffected by small or zero values of the β 's. These Figures therefore reveal no size distortion of any of the depicted statistics. Figures 2.2-2.6 show that the power of the FAR statistic is in general below the power of the other (combinations of) statistics but the FAR statistic has similar power than the other statistics when the β 's are very small. The smaller power of the FAR statistic in case of reasonably large β 's results from the larger degree of freedom parameter of its large sample distribution compared to the other statistics.

Figures 2.2-2.6 show that the combinations of the JGLS and GLS-LM statistics and JFM and FM-LM statistics overcome the spurious power decline of the GLS-LM and FM-LM statistics. These combined statistics therefore perform appropriately and also provide power against misspecification, *i.e.* no factor pricing. The combination of the JGLS and GLS-LM statistics and the CLR statistics have comparable power curves and outperform the combination of the JFM and FM-LM statistics. We note that the CLR statistic has no power against misspecification.

The power curves in Panel 2 clearly show the appeal of the statistics proposed in the previous section. Compared to the FM-Wald statistic, the optimal configuration of the statistics, which are the combined JGLS and GLS-LM statistics and the CLR statistic, do not indicate any sacrifice of power to improve the size. For example, in Figure 2.1, these statistics already dominate the FM-Wald statistic in terms of power. This continues to hold in the other Figures where the FM-Wald statistics becomes eventually enormously size distorted for very small values of the β 's.

5 Tests on some risk premia

The statistics that we proposed sofar test hypothezes that are specified on all risk premia, $H_0 : \lambda_F = \lambda_{F,0}$. Instead of testing hypothezes specified on all parameters, we often want to test hypothezes that are specified on sub-sets of the parameters, for example, $H_0^* : \theta_F = \theta_{F,0}$, where $\lambda_F = (\nu_F' : \theta_F')'$, with $\nu_F : k_\nu \times 1$; $\theta_{F,0}, \theta_F : k_\theta \times 1$ and $k = k_\nu + k_\theta$. Kleibergen (2005c) shows that the limiting distributions of the sub-set statistics in the linear instrumental variables regression model are boundedly pivotal (similar) when the partialled out parameter (ν_F) are estimated by maximum likelihood. This result extends to the linear factor model. The maximum likelihood estimator for ν_F given $H_0^* : \theta_F = \theta_{F,0}$, $\tilde{\nu}_F(\theta_{F,0})$, results from the first order condition for ν_F that results from (26), so

$$(\bar{\mathcal{R}} - \tilde{\mathcal{B}})' \tilde{\Sigma}^{-1} \tilde{\mathcal{B}}_\nu = 0, \quad (33)$$

with $\tilde{\mathcal{B}} = (\tilde{\mathcal{B}}_\nu : \tilde{\mathcal{B}}_\theta) = \sum_{t=1}^T \mathcal{R}_t (\bar{F}_t + (\tilde{\nu}_F(\theta_{F,0})'_{\theta_{F,0}})')' \left[\sum_{j=1}^T (\bar{F}_j + (\tilde{\nu}_F(\theta_{F,0})'_{\theta_{F,0}})') (\bar{F}_j + (\tilde{\nu}_F(\theta_{F,0})'_{\theta_{F,0}})')' \right]^{-1}$, $\tilde{\mathcal{B}}_\nu : (n - 1) \times k_\nu$, $\tilde{\mathcal{B}}_\theta : (n - 1) \times k_\theta$. The maximum likelihood estimator $\tilde{\nu}_F(\theta_{F,0})$ seems complicated to construct but, since it corresponds with the maximum likelihood estimator of Gibbons (1982), results in a straightforward manner from an eigenvalue problem, see Appendix G.

Theorem 9. *Under $H_0^* : \theta_F = \theta_{F,0}$, when Assumption 1 holds and $\tilde{\nu}_F(\theta_{F,0})$ is a maximum likelihood estimate of ν_F , the large sample distribution(s) of*

1. *FM-LM($(\tilde{\nu}_F(\theta_{F,0})' : \theta'_{F,0})'$) and GLS-LM($(\tilde{\nu}_F(\theta_{F,0})' : \theta'_{F,0})'$) are bounded from above by $\chi^2(k_\theta)$ distributions.*
2. *FAR($(\tilde{\nu}_F(\theta_{F,0})' : \theta'_{F,0})'$) is bounded from above by a $\chi^2(n - k_\nu - 1)$ distribution.*

3. $JFM((\tilde{\nu}_F(\theta_{F,0})' : \theta'_{F,0})')$ and $JGLS((\tilde{\nu}_F(\theta_{F,0})' : \theta'_{F,0})')$ are bounded from above by $\chi^2(n - k - 1)$ distributions which are independent from the large sample distributions of $FM-LM((\tilde{\nu}_F(\theta_{F,0})' : \theta'_{F,0})')$ and $GLS-LM((\tilde{\nu}_F(\theta_{F,0})' : \theta'_{F,0})')$.
4. $CLR((\tilde{\nu}_F(\theta_{F,0})' : \theta'_{F,0})')$ given $r((\tilde{\nu}_F(\theta_{F,0})' : \theta'_{F,0})')$ is bounded from above by

$$\frac{1}{2} \sqrt{\frac{\varphi_{k_\theta} + \varphi_{n-k-1} - r((\tilde{\nu}_F(\theta_{F,0})' : \theta'_{F,0})') + \sqrt{(\varphi_{k_\theta} + \varphi_{n-k-1} + r((\tilde{\nu}_F(\theta_{F,0})' : \theta'_{F,0})'))^2 - 4r((\tilde{\nu}_F(\theta_{F,0})' : \theta'_{F,0})')\varphi_{n-k-1}}{2}}$$

where φ_{k_θ} and φ_{n-k-1} are independent $\chi^2(k_\theta)$ and $\chi^2(n - k - 1)$ distributed random variables.

for all possible values of B_θ .

Proof. see Kleibergen (2005c). ■

Theorem 9 shows that the statistics that we constructed previously can still be used to conduct tests on some of the risk premia and become conservative when we do so. This results holds for all possible values of the β 's.

6 Confidence sets of the risk premia from Jagannathan and Wang (1996)

When we specify a range of values of $\lambda_{F,0}$, or $\theta_{F,0}$, we can use the statistics to construct an asymptotic confidence set for λ_F , or θ_F . The $(1 - \alpha) \times 100\%$ asymptotic confidence set contains all values of $\lambda_{F,0}$, or $\theta_{F,0}$, for which the value of the statistic that is used to test $H_0 : \lambda_F = \lambda_{F,0}$, or $H_0^* : \theta_F = \theta_{F,0}$, is below its $(1 - \alpha) \times 100\%$ asymptotic critical value.

We construct the asymptotic confidence sets for the risk premia for the asset returns from Jagannathan and Wang (1996). Jagannathan and Wang (1996) use the factor model,

$$R_t = \iota_n \lambda_1 + \text{size} \times \lambda_s + \beta(\bar{F}_t + \lambda_F) + \varepsilon_t, \quad t = 1, \dots, T, \quad (34)$$

to describe the return on hundred size and β sorted portfolios that are collected into the vector of asset returns R_t , so $n = 100$, for three hundred and thirty monthly observations, so $T = 330$. The $n \times 1$ vector size reflects the relative size of the different portfolios and is constant over time. The $k \times 1$ vector of (demeaned) factors \bar{F}_t contains three different factors: the return on a value weighted portfolio, the yield premium between low and high grade corporate bonds and the growth of per capita labor income. Since our interest is on the risk premia λ_F , we remove λ_1 and λ_s from the factor model (34) in an analogous manner as by taking the returns in deviation from the n -th asset return as we proposed previously to remove λ_1 . The removal of the size factor implies that the value of n in the degrees of freedom parameters of the (bounding) limiting distributions in Theorem 9 reduces to $n - 1$.

Panel 3: One minus p -value plots for λ_{vw} , λ_{prem} and λ_{labor} .

λ_{vw} :

Figure 3.1

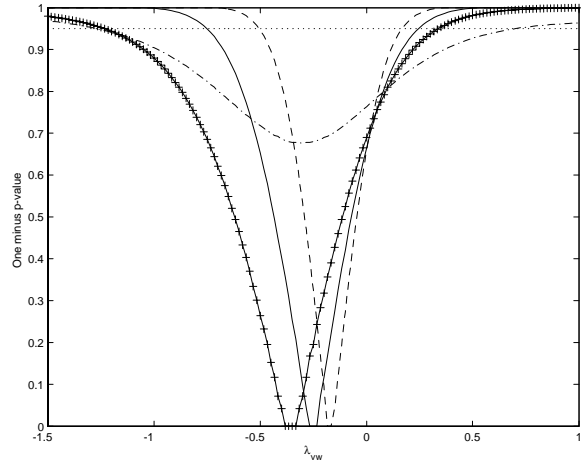
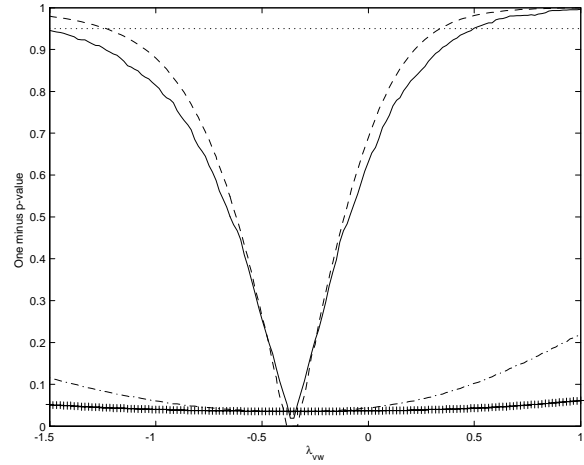


Figure 3.2



λ_{prem} :

Figure 3.3

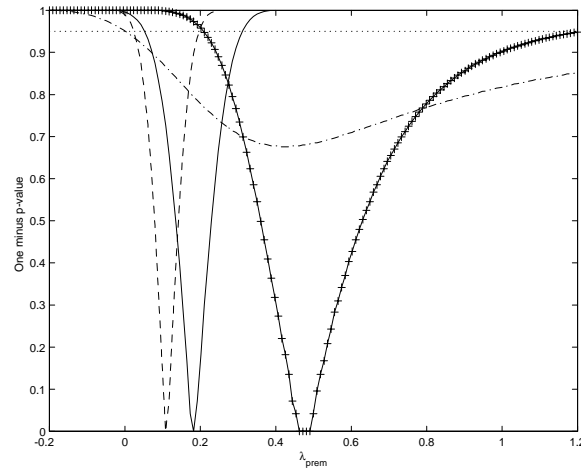
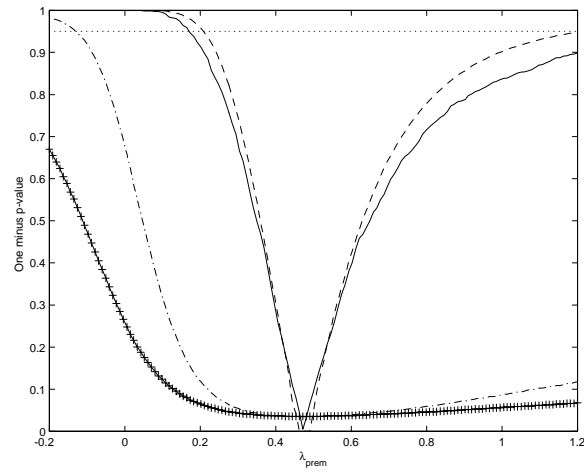


Figure 3.4



λ_{labor} :

Figure 3.5

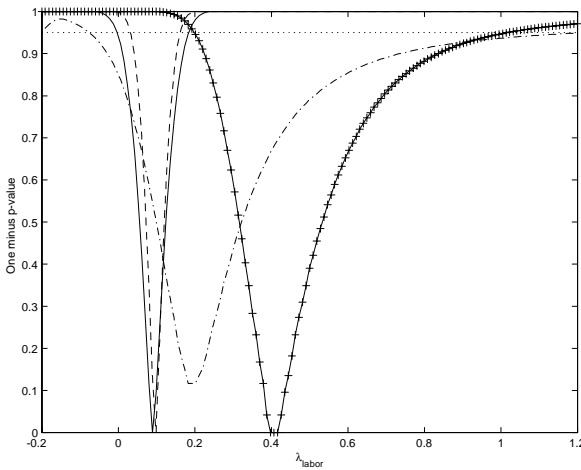
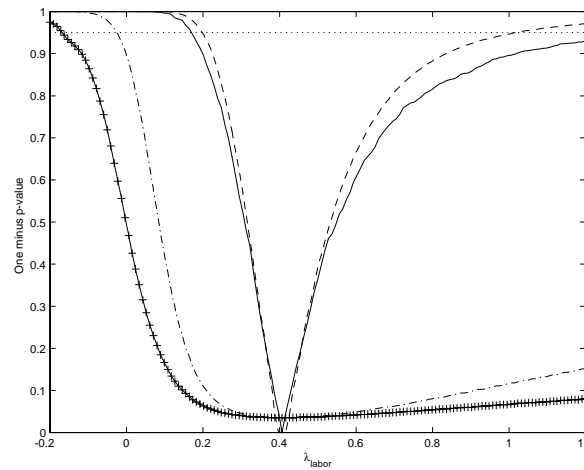


Figure 3.6



Figures 3.1-3.5: One minus p -value plot for FM-W (solid), GLS-W (dashed), FM-LM (dashed-dotted) and GLS-LM (solid with plusses) statistics.

Figures 3.2-3.6: One minus p -value plot for CLR (solid), JGLS (solid with plusses), FAR (dashed-dotted) and GLS-LM (dashed) statistics.

Panel 3 contains the p -value plots of the tests on the different risk premia. Figures 3.1 and 3.2 in Panel 3 show the p -value plots of tests of the hypothesis $H_0 : \lambda_{vw} = \lambda_{vw,0}$, where λ_{vw} is the risk premium on the value weighted portfolio, for a range of values of $\lambda_{vw,0}$. Figures 3.3 and 3.4 contain the p -value plots of tests of $H_0 : \lambda_{\text{prem}} = \lambda_{\text{prem},0}$, where λ_{prem} is the risk premium on the yield premium, for a range of values of $\lambda_{\text{prem},0}$ and Figures 3.5 and 3.6 contain the p -value plots of tests of $H_0 : \lambda_{\text{labor}} = \lambda_{\text{labor},0}$, where λ_{labor} is the risk premium on the labor income growth, for a range of values of $\lambda_{\text{labor},0}$. The tests of the hypotheses on the risk premia are all conducted in the linear factor model (34) and thus test only one element of the vector of risk premia. The maximum likelihood estimator is used for the other risk premia. The p -values therefore result from the bounding limiting distributions in Theorem 9 and are thus conservative. The p -values for the CLR statistic were obtained by using its bounding conditional limiting distribution given $r((\tilde{\nu}_F(\theta_{F,0})' : \theta'_{F,0})')$ which is computed for a large range of values of $r((\tilde{\nu}_F(\theta_{F,0})' : \theta'_{F,0})')$. All Figures in Panel 3 contain a dotted line at 95% whose intersection with the p -value plots reveals the 95% confidence set.

Figures 3.1 and 3.2 show the p -values for tests on the risk premium of the value weighted return. Figure 3.1 (and Figures 3.3 and 3.5) contain the p -value plots that result from the FM-W, GLS-W, FM-LM and GLS-LM statistics. Most of the p -value plots in Figure 3.1 are rather similar which results since the β 's of the value weighted return are sizeable. The main difference between the p -value plots of the FM and GLS Wald statistics and the GLS-LM statistic is therefore that the latter leads to a larger 95% confidence set. The p -value plot of the FM-LM statistic is rather strange and given its inferior behavior in the power study compared to the size-correct GLS-LM and CLR statistics, we do not discuss this statistic any further. All statistics do not reject the hypothesis of a zero risk premium at the 95% significance level which is as expected since the risk premium on portfolio returns, like the value weighted index, is typically assumed to be equal to zero.

Figure 3.2 (and Figures 3.4, 3.6) only contains p -value plots of statistics that remain size-correct in case of small and/or many β 's, *i.e.* the GLS-LM, CLR, JGLS and FAR statistics. The p -value plots of the JGLS and FAR statistic are all rather low which is due to the large degree of freedom parameter of their (bounding) χ^2 limiting distributions, *i.e.* 95 and 96. These p -value plots show that the factor pricing hypothesis is not rejected. The p -value plots of the GLS-LM and CLR statistic are very similar which results since the conditioning statistic which is used in the CLR statistic is rather large.

Figures 3.3 and 3.4 show the p -values for tests of the risk premium on the yield premium. The p -value plots in Figure 3.3 differ considerably and show that, because of the rather small β 's of the yield premium which are probably also only non-zero for a fraction of the large number of portfolios, there is a considerable downward bias in the FM and GLS risk premium estimators which results since λ_F exceeds μ_F in the bias expression in Theorem 1.3. The maximum likelihood estimate of the risk premium, which is the value of the risk premium where the p -value plot of GLS-LM equals zero, is therefore considerably larger than the FM estimate, which is the value of the risk premium where the p -value plot of FM-W equals zero. The GLS-LM statistic therefore leads to a much larger confidence set for the risk premium on the yield premium which also lies at considerably larger values.

Figure 3.4 shows the p -values for tests on the risk premium using the size correct GLS-LM, CLR, JGLS and FAR statistics. The p -value plots show that the results from the CLR statistic are similar to that of the GLS-LM statistic. The p -values that result from the JGLS and FAR reveal that the factors are the only priced elements and factor pricing is thus not rejected.

Figures 3.5 and 3.6 show the p -values for tests of the risk premium on labor income growth. Because the β 's on labor income growth are rather small and are probably only non-zero for a fraction of the large number of portfolios, there is again a considerable downward bias in the FM and GLS risk premium estimators. The p -value plots that result from the FM and GLS Wald statistics and the GLS-LM statistic differ therefore considerably. The 95% confidence set that results from the GLS-LM statistic

is much larger than the one that results from the FM Wald statistic and contains much larger values of the risk premium. It is striking to see that the intersection of both confidence sets is empty. The p -value plot of the CLR statistic in Figure 3.6 is similar to that of the GLS-LM statistic. The JGLS and FAR statistics in Figure 3.6 show that no mispricing is occurring and that factor pricing is not rejected.

Panel 3 shows the importance of using the size-correct statistics that we introduced previously. Because of the large number of portfolios in the Jagannathan-Wang data, the bias in the FM risk premia estimator becomes considerable when the β 's of the associated factors are small or only non-zero for a fraction of the portfolios. The Wald statistic based on the FM risk premia estimator therefore becomes size distorted and unreliable in these cases. The statistics that we introduced remain size-correct and lead to inference that is centered around the maximum likelihood estimator of Gibbons (1982). These statistics therefore do a better job in reflecting the risk premium on the factor whose β 's are relatively small and/or only matter for a fraction of the assets. This is nicely shown for the Jagannathan-Wang data where the confidence sets that result from the FM Wald statistic show that the risk premia are almost negligible while the confidence sets that result from the CLR and GLS-LM statistics indicate sizeable and thus important risk premia. This shows that the usage of improved inferential procedures can lead to more interesting conclusions.

7 Conclusions

We reveal the inadequacy of standard inferential procedures that are based on the FM and GLS risk premia estimators when the β 's are small and/or the number of assets is large. We propose some new statistics whose large sample distributions remain trustworthy in such cases. We apply these statistics to the size and β sorted portfolio return series from Jagannathan and Wang (1996). The confidence sets that result from the FM and GLS Wald statistics show that the risk premia are rather small and perhaps negligible. The confidence set that result from the new statistics indicate a more interesting conclusion as they indicate much larger values of the risk premia which thus show that the risk premia are important.

The linear factor models analyzed in this paper assume a constant covariance matrix of the asset returns over time. We will relax this assumption in future work.

Appendix

A. Assumption 1. The central limit theorem for the specification of the linear factor model (2) implied by Assumption 1 reads

$$\begin{aligned}
& \frac{1}{\sqrt{T}} \sum_{t=1}^T \begin{pmatrix} \begin{pmatrix} 1 \\ F_t \end{pmatrix} \otimes u_t \\ F_t - \mu_f \end{pmatrix} \\
&= \frac{1}{\sqrt{T}} \sum_{t=1}^T \begin{pmatrix} \begin{pmatrix} 1 \\ F_t \end{pmatrix} \otimes (R_t - \iota_n \lambda_1 - \beta(F_t - \mu_F + \lambda_F)) \\ F_t - \mu_f \end{pmatrix} \\
&= \frac{1}{\sqrt{T}} \sum_{t=1}^T \begin{pmatrix} \begin{pmatrix} 1 \\ F_t \end{pmatrix} \otimes (R_t - \iota_n \lambda_1 - \beta(F_t - \bar{F} + \bar{F} - \mu_F + \lambda_F)) \\ F_t - \mu_f \end{pmatrix} \\
&= \frac{1}{\sqrt{T}} \begin{pmatrix} \sum_{t=1}^T \begin{pmatrix} 1 \\ F_t \end{pmatrix} \otimes (R_t - \iota_n \lambda_1 - \beta(\bar{F}_t + \lambda_F)) - \sum_{t=1}^T \begin{pmatrix} 1 \\ F_t \end{pmatrix} \otimes \beta(\bar{F} - \mu_F) \\ \sum_{t=1}^T F_t - \mu_f \end{pmatrix} \\
&= \begin{pmatrix} \frac{1}{\sqrt{T}} \sum_{t=1}^T \begin{pmatrix} 1 \\ F_t \end{pmatrix} \otimes (R_t - \iota_n \lambda_1 - \beta(\bar{F}_t + \lambda_F)) - \sqrt{T} \begin{pmatrix} 1 \\ \bar{F} \end{pmatrix} \otimes \beta(\bar{F} - \mu_F) \\ \sqrt{T}(\bar{F} - \mu_f) \end{pmatrix} \\
&\xrightarrow{d} \begin{pmatrix} \varphi_R - \beta\varphi_F \\ \varphi_\beta - (\mu_F \otimes \beta)\varphi_F \\ \varphi_F \end{pmatrix}.
\end{aligned}$$

which results in convergence to a normal distributed random vector with mean zero and covariance matrix

$$\begin{pmatrix} \begin{pmatrix} 1 & \mu'_F \\ \mu_F & \mu_F \mu'_F \end{pmatrix} \otimes (\Omega + \beta V_{FF} \beta') + \begin{pmatrix} 0 & 0 \\ 0 & V_{FF} \end{pmatrix} \otimes \Omega & \begin{matrix} -\beta V_{FF} \\ -(\mu_F \otimes \beta) V_{FF} \end{matrix} \\ -V_{FF} \beta' & V_{FF} \end{pmatrix}.$$

B. Proof of Lemma 1. Assumption 1 states that

$$\frac{1}{\sqrt{T}} \sum_{t=1}^T \begin{pmatrix} \begin{pmatrix} 1 \\ F_t \end{pmatrix} \otimes (R_t - \iota_n \lambda_1 - \beta(\bar{F}_t + \lambda_F)) \\ F_t - \mu_F \end{pmatrix} \xrightarrow{d} \begin{pmatrix} \varphi_R \\ \varphi_\beta \\ \varphi_F \end{pmatrix},$$

with $(\varphi'_R \ \varphi'_\beta \ \varphi'_F)' \sim N(0, V)$. We pre-multiply it with

$$\begin{pmatrix} \begin{matrix} 1 & 0 \\ -(\frac{1}{T} \sum_{j=1}^T \bar{F}_j \bar{F}'_j)^{-1} \bar{F} & (\frac{1}{T} \sum_{j=1}^T \bar{F}_j \bar{F}'_j)^{-1} \end{matrix} \otimes I_n & 0 \\ 0 & I_k \end{pmatrix}$$

to obtain, since $\sum_{t=1}^T \bar{F}_t = 0$, $\hat{\beta} = \sum_{t=1}^T R_t \bar{F}'_t (\sum_{j=1}^T \bar{F}_j \bar{F}'_j)^{-1}$,

$$\begin{aligned}
& \sum_{t=1}^T [(\iota_n \lambda_1 + \beta(\bar{F}_t + \lambda_F)) F'_t - (\iota_n \lambda_1 + \beta(\bar{F}_t + \lambda_F)) \bar{F}'_t] (\sum_{j=1}^T \bar{F}_j \bar{F}'_j)^{-1} = \\
& \sum_{t=1}^T [\beta \bar{F}'_t (F_t - \bar{F})'] (\sum_{j=1}^T \bar{F}_j \bar{F}'_j)^{-1} = \beta,
\end{aligned}$$

and

$$\lim_{T \rightarrow \infty} \begin{pmatrix} 1 & 0 \\ -(\frac{1}{T} \sum_{j=1}^T \bar{F}_j \bar{F}'_j)^{-1} \bar{F} & (\frac{1}{T} \sum_{j=1}^T \bar{F}_j \bar{F}'_j)^{-1} \end{pmatrix} Q \begin{pmatrix} 1 & 0 \\ -(\frac{1}{T} \sum_{j=1}^T \bar{F}_j \bar{F}'_j)^{-1} \bar{F} & (\frac{1}{T} \sum_{j=1}^T \bar{F}_j \bar{F}'_j)^{-1} \end{pmatrix}' = \\
\begin{pmatrix} Q_{11} & 0 \\ 0 & Q_{FF.1}^{-1} \end{pmatrix}$$

with $Q_{FF,1} = \lim_{T \rightarrow \infty} E(\frac{1}{T} \sum_{j=1}^T \bar{F}_j \bar{F}_j')$, that

$$\sqrt{T} \begin{pmatrix} \bar{R} - \iota_n \lambda_1 - \beta \lambda_F \\ \text{vec}(\hat{\beta} - \beta) \\ \bar{F} - \mu_F \end{pmatrix} \xrightarrow{d} \begin{pmatrix} \psi_R \\ \psi_\beta \\ \psi_F \end{pmatrix},$$

where ψ_R , ψ_β and ψ_F are independent $n \times 1$, $nk \times 1$ and $k \times 1$ normal distributed random vectors with mean 0 and covariance matrices, Ω , $Q_{FF,1}^{-1} \otimes \Omega$ and V_{FF} .

C. Proof of Theorem 1. To analyze the behavior of the FM estimator, we rephrase Lemma 1 as

$$\begin{pmatrix} \bar{R} \\ \text{vec}(\hat{\beta}) \\ \bar{F} \end{pmatrix} \underset{a}{=} \begin{pmatrix} \iota_n \lambda_1 + \beta \lambda_k \\ \text{vec}(\beta) \\ \mu_F \end{pmatrix} + \frac{1}{\sqrt{T}} \begin{pmatrix} \psi_R \\ \psi_\beta \\ \psi_F \end{pmatrix} + o_p(\frac{1}{\sqrt{T}}),$$

where $\underset{a}{=}$ implies equality in large samples and $o_p(\frac{1}{\sqrt{T}})$ indicates that all other elements are of a lower order than $\frac{1}{\sqrt{T}}$.

1. $\beta = 0$. We use that since $(\iota_n : \frac{1}{\sqrt{T}} \Psi_\beta)' (\iota_n : \frac{1}{\sqrt{T}} \Psi_\beta) = \begin{pmatrix} \iota_n' \iota_n & \frac{1}{\sqrt{T}} \iota_n' \Psi_\beta \\ \frac{1}{\sqrt{T}} \Psi_\beta' \iota_n & \frac{1}{T} \Psi_\beta' \Psi_\beta \end{pmatrix}$,

$$\begin{pmatrix} (\iota_n : \frac{1}{\sqrt{T}} \Psi_\beta)' (\iota_n : \frac{1}{\sqrt{T}} \Psi_\beta) \\ a^{-1} & -a^{-1} \iota_n' \Psi_\beta (\Psi_\beta' \Psi_\beta)^{-1} \\ -\sqrt{T} (\Psi_\beta' M_{i_n} \Psi_\beta)^{-1} \frac{\Psi_\beta' \iota_n}{n} & T (\Psi_\beta' M_{i_n} \Psi_\beta)^{-1} \end{pmatrix}^{-1}.$$

with $a = \iota_n' M_{\Psi_\beta} \iota_n$.

- 1a. $E(\bar{R}) = \iota_n \lambda_1$ and

$$\begin{aligned} \begin{pmatrix} \hat{\lambda}_1 \\ \hat{\lambda}_F \end{pmatrix} &= \left[(\iota_n : \hat{\beta})' (\iota_n : \hat{\beta}) \right]^{-1} (\iota_n : \hat{\beta})' \bar{R} \\ &\underset{a}{=} \left[(\iota_n : \frac{1}{\sqrt{T}} \Psi_\beta)' (\iota_n : \frac{1}{\sqrt{T}} \Psi_\beta) \right]^{-1} (\iota_n : \frac{1}{\sqrt{T}} \Psi_\beta)' (\iota_n \lambda_1 + \frac{1}{\sqrt{T}} \psi_R) \\ &\underset{a}{=} \begin{pmatrix} a^{-1} & -a^{-1} \iota_n' \Psi_\beta (\Psi_\beta' \Psi_\beta)^{-1} \\ -\sqrt{T} (\Psi_\beta' M_{i_n} \Psi_\beta)^{-1} \frac{\Psi_\beta' \iota_n}{n} & T (\Psi_\beta' M_{i_n} \Psi_\beta)^{-1} \end{pmatrix} \\ &\begin{pmatrix} \iota_n' \iota_n \lambda_1 + \frac{1}{\sqrt{T}} \iota_n' \psi_R \\ \frac{1}{\sqrt{T}} \Psi_\beta' \iota_n \lambda_1 + \frac{1}{T} \Psi_\beta' \psi_R \end{pmatrix} \\ &\underset{a}{=} \begin{pmatrix} a^{-1} (a \lambda_1 + \frac{1}{\sqrt{T}} b' \psi_R) \\ (\Psi_\beta' M_{i_n} \Psi_\beta)^{-1} \Psi_\beta' M_{i_n} \psi_R \end{pmatrix} \\ &\underset{a}{=} \begin{pmatrix} \lambda_1 \\ 0 \end{pmatrix} + \begin{pmatrix} \frac{1}{\sqrt{T}} \frac{1}{a} b' \\ (\Psi_\beta' M_{i_n} \Psi_\beta)^{-1} \Psi_\beta' M_{i_n} \end{pmatrix} \psi_R \end{aligned}$$

where $\Psi_\beta = \text{vecin}_k(\psi_\beta)$ and $b = M_{\Psi_\beta} \iota_n$, so, since Ψ_β and ψ_R are independent,

$$\begin{pmatrix} \sqrt{T}(\hat{\lambda}_1 - \lambda_1) \\ \hat{\lambda}_F \end{pmatrix} \xrightarrow{d} \begin{pmatrix} \frac{1}{\sqrt{T}} \frac{1}{a} b' \\ (\Psi_\beta' M_{i_n} \Psi_\beta)^{-1} \Psi_\beta' M_{i_n} \end{pmatrix} \psi_R.$$

Since ψ_R is a normal distributed random vector that is independent of Ψ_β , the moments of limiting distribution of $\hat{\lambda}_F$ are determined by the moments of $(\Psi_\beta' M_{i_n} \Psi_\beta)^{-1} \Psi_\beta' M_{i_n}$. When Ω is an identity

matrix, $\Psi'_\beta M_{i_n} \Psi_\beta$ is a Wishart distributed random matrix with $n - 1$ degrees of freedom and scale matrix Q_{FF} , see *e.g.* Muirhead (1982). Hence, $(\Psi'_\beta M_{i_n} \Psi_\beta)^{-1} \Psi'_\beta M_{i_n}$ is the square root of the inverse of the Wishart distributed random matrix. To proof the order of finite moments of this matrix, we use an alternative definition of a multi-variate t distributed random variable.

A multi-variate t distributed random $k \times 1$ vector x with ν degrees, $k \times k$ dimensional scale matrix Σ^{-1} and $k \times 1$ scale vector α can be defined by

$$x = \alpha + W^{-\frac{1}{2}} z,$$

where $z \sim N(0, I)$ and independent of the $k \times k$ Wishart distributed random matrix W with $\nu + k - 1$ degrees of freedom and scale matrix Σ . Since the moments of the multi-variate t exist up to order ν , the moments of the square root of the inverted-Wishart thus exist up to order $\nu + k - 1$. This definition of a multi-variate t distributed random matrix results from the definition of a matrix-variate t distributed random matrix, see *e.g.* Zellner (1971) and Drèze and Richard (1983).

In our case $W = \Psi'_\beta M_{i_n} \Psi_\beta$ which is a $k \times k$ Wishart distributed matrix with $n - 1$ degrees of freedom. The moments of $(\Psi'_\beta M_{i_n} \Psi_\beta)^{-1} \Psi'_\beta M_{i_n}$ exist therefore up to order $n - 1 - k + 1 = n - k$ which holds as well for the moments of $\hat{\lambda}_F$. The difference of the covariance matrix of Ω from the identity matrix does not change the order of finite moments.

1b. When β is equal to zero and $E(\bar{R}) = c$:

$$\begin{aligned} \begin{pmatrix} \hat{\lambda}_1 \\ \hat{\lambda}_F \end{pmatrix} &= \left[(\iota_n : \hat{\beta})' (\iota_n : \hat{\beta}) \right]^{-1} (\iota_n : \hat{\beta})' \bar{R} \\ &= \frac{1}{a} \left[(\iota_n : \frac{1}{\sqrt{T}} \Psi_\beta)' (\iota_n : \frac{1}{\sqrt{T}} \Psi_\beta) \right]^{-1} (\iota_n : \frac{1}{\sqrt{T}} \Psi_\beta)' (c + \frac{1}{\sqrt{T}} \psi_R) \\ &= \begin{pmatrix} a^{-1} & -a^{-1} \iota_n' \Psi_\beta (\Psi'_\beta \Psi_\beta)^{-1} \\ -\sqrt{T} (\Psi'_\beta M_{i_n} \Psi_\beta)^{-1} \frac{\Psi'_\beta \iota_n}{n} & T (\Psi'_\beta M_{i_n} \Psi_\beta)^{-1} \end{pmatrix} \\ &\quad \begin{pmatrix} \iota_n' c + \frac{1}{\sqrt{T}} \iota_n' \psi_R \\ \frac{1}{\sqrt{T}} \Psi'_\beta c + \frac{1}{T} \Psi'_\beta \psi_R \end{pmatrix} \\ &= \begin{pmatrix} a^{-1} b' c + \frac{1}{\sqrt{T}} a^{-1} b' \psi_R \\ \sqrt{T} (\Psi'_\beta M_{i_n} \Psi_\beta)^{-1} \Psi_\beta^{-1} \Psi'_\beta M_{i_n} c + (\Psi'_\beta M_{i_n} \Psi_\beta)^{-1} \Psi_\beta^{-1} \Psi'_\beta M_{i_n} \psi_R \end{pmatrix}. \\ &= \begin{pmatrix} \frac{1}{a} b' \\ \sqrt{T} (\Psi'_\beta M_{i_n} \Psi_\beta)^{-1} \Psi_\beta^{-1} \Psi'_\beta M_{i_n} c \end{pmatrix} c + \begin{pmatrix} \frac{1}{\sqrt{T}} \frac{1}{a} b' \\ (\Psi'_\beta M_{i_n} \Psi_\beta)^{-1} \Psi_\beta^{-1} \Psi'_\beta M_{i_n} \end{pmatrix} \psi_R \end{aligned}$$

and

$$\begin{pmatrix} \sqrt{T} (\hat{\lambda}_1 - \frac{1}{a} b' c) \\ \hat{\lambda}_F - \sqrt{T} (\Psi'_\beta M_{i_n} \Psi_\beta)^{-1} \Psi_\beta^{-1} \Psi'_\beta M_{i_n} c \end{pmatrix} \xrightarrow{d} \begin{pmatrix} \frac{1}{\sqrt{T}} \frac{1}{a} b' \\ (\Psi'_\beta M_{i_n} \Psi_\beta)^{-1} \Psi_\beta^{-1} \Psi'_\beta M_{i_n} \end{pmatrix} \psi_R$$

so $\hat{\lambda}_F$ diverges when T gets large.

2a. $\beta = \frac{1}{\sqrt{T}}B$, with B a $n \times k$ matrix of full rank and $E(\bar{R}) = \iota_n \lambda_1 + \frac{1}{\sqrt{T}}B\lambda_F$ so

$$\begin{aligned}
& \begin{pmatrix} \hat{\lambda}_1 \\ \hat{\lambda}_F \end{pmatrix} = \frac{1}{a} \left[(\iota_n : \hat{\beta})' (\iota_n : \hat{\beta}) \right]^{-1} (\iota_n : \hat{\beta})' \bar{R} \\
& = \frac{1}{a} \left[(\iota_n : \frac{1}{\sqrt{T}}(B + \Psi_\beta))' (\iota_n : \frac{1}{\sqrt{T}}(B + \Psi_\beta)) \right]^{-1} (\iota_n : \frac{1}{\sqrt{T}}(B + \Psi_\beta))' (\iota_n \lambda_1 + \frac{1}{\sqrt{T}}(B\lambda_F + \psi_R)) \\
& = \frac{1}{a} \begin{pmatrix} a^{-1} & -a^{-1}\sqrt{T}\iota_n'(B + \Psi_\beta)((B + \Psi_\beta)'(B + \Psi_\beta))^{-1} \\ -\sqrt{T}[(B + \Psi_\beta)'M_{\iota_n}(B + \Psi_\beta)]^{-1}\frac{(B + \Psi_\beta)'\iota_n}{n} & [(B + \Psi_\beta)'M_{\iota_n}(B + \Psi_\beta)]^{-1} \\ \left(\begin{array}{c} \iota_n'\iota_n\lambda_1 + \frac{1}{\sqrt{T}}\iota_n'(B\lambda_F + \psi_R) \\ \frac{1}{\sqrt{T}}(B + \Psi_\beta)'\iota_n\lambda_1 + \frac{1}{T}(B + \Psi_\beta)'(B\lambda_F + \psi_R) \end{array} \right) \end{pmatrix} \\
& = \frac{1}{a} \begin{pmatrix} a^{-1}(a\lambda_1 + \frac{1}{\sqrt{T}}a^{-1}b'(B\lambda_F + \psi_R)) \\ ((B + \Psi_\beta)'M_{\iota_n}(B + \Psi_\beta))^{-1}(B + \Psi_\beta)'M_{\iota_n}(B\lambda_F + \psi_R) \end{pmatrix} \\
& = \frac{1}{a} \begin{pmatrix} \lambda_1 \\ 0 \end{pmatrix} + \begin{pmatrix} \frac{1}{\sqrt{T}}\frac{1}{a}b' \\ ((B + \Psi_\beta)'M_{\iota_n}(B + \Psi_\beta))^{-1}(B + \Psi_\beta)'M_{\iota_n} \end{pmatrix} (B\lambda_F + \psi_R)
\end{aligned}$$

with $a = \iota_n' M_{(B + \Psi_\beta)} \iota_n$, $b = M_{(B + \Psi_\beta)} \iota_n$, so,

$$\begin{pmatrix} \sqrt{T}(\hat{\lambda}_1 - \lambda_1) \\ \hat{\lambda}_F - \lambda_F \end{pmatrix} \xrightarrow{d} \begin{pmatrix} \frac{1}{\sqrt{T}}\frac{1}{a}b' \\ [(B + \Psi_\beta)'(B + \Psi_\beta)]^{-1}(B + \Psi_\beta)'(I_n - \frac{1}{a}\iota_n b') \end{pmatrix} (\psi_{R,\beta} - \Psi_\beta \lambda_F).$$

In this case $(B + \Psi_\beta)'(B + \Psi_\beta)$ has a non-central Wishart distribution when $\Omega = I_n$ which has no implications for the order of the finite moments compared to the case when $B = 0$. Hence, the moments of $\hat{\lambda}_F$ exist up to order $n - k$.

2b. $\beta = \frac{1}{\sqrt{T}}B$, with B a $n \times k$ matrix of full rank and $E(\bar{R}) = c$:

$$\begin{aligned}
& \begin{pmatrix} \hat{\lambda}_1 \\ \hat{\lambda}_F \end{pmatrix} = \frac{1}{a} \left[(\iota_n : \hat{\beta})' (\iota_n : \hat{\beta}) \right]^{-1} (\iota_n : \hat{\beta})' \bar{R} \\
& = \frac{1}{a} \left[(\iota_n : \frac{1}{\sqrt{T}}(B + \Psi_\beta))' (\iota_n : \frac{1}{\sqrt{T}}(B + \Psi_\beta)) \right]^{-1} (\iota_n : \frac{1}{\sqrt{T}}(B + \Psi_\beta))' (c + \frac{1}{\sqrt{T}}\psi_R) \\
& = \frac{1}{a} \begin{pmatrix} a^{-1} & -a^{-1}\sqrt{T}\iota_n'(B + \Psi_\beta)((B + \Psi_\beta)'(B + \Psi_\beta))^{-1} \\ -\sqrt{T}[(B + \Psi_\beta)'M_{\iota_n}(B + \Psi_\beta)]^{-1}\frac{(B + \Psi_\beta)'\iota_n}{n} & [(B + \Psi_\beta)'M_{\iota_n}(B + \Psi_\beta)]^{-1} \\ \left(\begin{array}{c} \iota_n'c + \frac{1}{\sqrt{T}}\iota_n'\psi_R \\ \frac{1}{\sqrt{T}}(B + \Psi_\beta)'c + \frac{1}{T}(B + \Psi_\beta)'\psi_R \end{array} \right) \end{pmatrix} \\
& = \frac{1}{a} \begin{pmatrix} a^{-1}b'c + \frac{1}{\sqrt{T}}a^{-1}b'\psi_R \\ [(B + \Psi_\beta)'M_{\iota_n}(B + \Psi_\beta)]^{-1}(B + \Psi_\beta)'M_{\iota_n}[\sqrt{T}c + \psi_R] \end{pmatrix} \\
& = \frac{1}{a} \begin{pmatrix} \frac{1}{a}b' \\ \sqrt{T}[(B + \Psi_\beta)'M_{\iota_n}(B + \Psi_\beta)]^{-1}(B + \Psi_\beta)'M_{\iota_n} \end{pmatrix} c + \\
& \quad \begin{pmatrix} \frac{1}{\sqrt{T}}\frac{1}{a}b' \\ [(B + \Psi_\beta)'M_{\iota_n}(B + \Psi_\beta)]^{-1}(B + \Psi_\beta)'M_{\iota_n} \end{pmatrix} \psi_R
\end{aligned}$$

with $a = \iota_n' M_{(B + \Psi_\beta)} \iota_n$, $b = M_{(B + \Psi_\beta)} \iota_n$, and

$$\begin{aligned}
& \begin{pmatrix} \sqrt{T}(\hat{\lambda}_1 - \lambda_1) \\ \hat{\lambda}_F - \sqrt{T}[(B + \Psi_\beta)'(B + \Psi_\beta)]^{-1}(B + \Psi_\beta)'(I_n - \frac{1}{a}\iota_n b')c \end{pmatrix} \\
& \xrightarrow{d} \begin{pmatrix} \frac{1}{\sqrt{T}}\frac{1}{a}b' \\ [(B + \Psi_\beta)'(B + \Psi_\beta)]^{-1}(B + \Psi_\beta)'(I_n - \frac{1}{a}\iota_n b') \end{pmatrix} \psi_R.
\end{aligned}$$

3. $\beta = \begin{pmatrix} \beta_1 \\ 0 \end{pmatrix}$, with $\beta_1 : n_1 \times 1$ and $E(\bar{R}) = \iota_n \lambda_1 + B \lambda_F$. The specification of the FM risk premia estimator is

$$\begin{aligned} \begin{pmatrix} \hat{\lambda}_1 \\ \hat{\lambda}_F \end{pmatrix} &= \frac{1}{a} \left[(\iota_n \vdots \hat{\beta})' (\iota_n \vdots \hat{\beta}) \right]^{-1} (\iota_n \vdots \hat{\beta})' \bar{R} \\ &= \frac{1}{a} \left[(\iota_n \vdots ((\beta_1)_0) + \frac{1}{\sqrt{T}} \Psi_\beta)' (\iota_n \vdots ((\beta_1)_0) + \frac{1}{\sqrt{T}} \Psi_\beta) \right]^{-1} (\iota_n \vdots ((\beta_1)_0) + \frac{1}{\sqrt{T}} \Psi_\beta)' (\iota_n \lambda_1 + (B \lambda_F + \psi_R)) + o_p\left(\frac{1}{\sqrt{T}}\right). \end{aligned}$$

We construct the limit of the two different elements of $\begin{pmatrix} \hat{\lambda}_1 \\ \hat{\lambda}_F \end{pmatrix}$. We assume that n_1 remains fixed and that the following limits hold when n and T become large: $\frac{\Psi'_\beta \Psi_\beta}{n} \xrightarrow{p} W$, $W = [\lim_{n \rightarrow \infty} \text{tr}(\frac{1}{n} \Omega)] V_{FF}^{-1}$, $\frac{n}{T} \rightarrow c$, $\frac{n_1}{T} \rightarrow 0$ such that $\frac{1}{\sqrt{T}} \beta'_1 \Psi_{\beta_1} \xrightarrow{p} 0$, $\frac{1}{\sqrt{n}} \beta'_1 \iota_{n_1} \rightarrow 0$ such that $\frac{n_1}{\sqrt{n}}$

$$\begin{aligned} &\begin{pmatrix} \frac{1}{\sqrt{n}} & 0 \\ 0 & 1 \end{pmatrix} (\iota_n \vdots ((\beta_1)_0) + \frac{1}{\sqrt{T}} \Psi_\beta)' (\iota_n \vdots ((\beta_1)_0) + \frac{1}{\sqrt{T}} \Psi_\beta) \begin{pmatrix} \frac{1}{\sqrt{n}} & 0 \\ 0 & 1 \end{pmatrix} = \\ &\begin{pmatrix} 1 & \frac{1}{\sqrt{n}} \iota'_{n_1} \beta_1 + \frac{1}{\sqrt{T}} (\frac{1}{\sqrt{n}} \iota'_{n_1} \Psi_\beta) \\ \frac{1}{\sqrt{n}} \beta'_1 \iota_{n_1} + \frac{1}{\sqrt{T}} (\frac{1}{\sqrt{n}} \Psi'_\beta \iota_n) & \beta'_1 \beta_1 + \frac{1}{\sqrt{T}} \beta'_1 \Psi_{\beta_1} + \frac{1}{\sqrt{T}} \Psi'_{\beta_1} \beta_1 + \frac{n}{T} \frac{\Psi'_\beta \Psi_\beta}{n} \end{pmatrix} = \\ &\begin{pmatrix} 1 & b'_1 \\ b_1 & B_1 \end{pmatrix} \end{aligned}$$

with $b_1 = \frac{1}{\sqrt{n}} \beta'_1 \iota_{n_1} + \frac{1}{\sqrt{T}} (\frac{1}{\sqrt{n}} \Psi'_\beta \iota_n)$, $B_1 = \beta'_1 \beta_1 + \frac{1}{\sqrt{T}} \beta'_1 \Psi_{\beta_1} + \frac{1}{\sqrt{T}} \Psi'_{\beta_1} \beta_1 + \frac{n}{T} \frac{\Psi'_\beta \Psi_\beta}{n}$, $\Psi_{\beta_1} = (I_{n_1} \vdots 0) \Psi_\beta$, and its inverse reads

$$\begin{pmatrix} (1 - b'_1 B_1^{-1} b_1)^{-1} & -(1 - b'_1 B_1^{-1} b_1)^{-1} b'_1 B_1^{-1} \\ -(B_1 - b_1 b'_1)^{-1} b_1 & (B_1 - b_1 b'_1)^{-1} \end{pmatrix}.$$

The large sample behavior of the second part of the FM estimator reads

$$\begin{aligned} &\begin{pmatrix} \frac{1}{\sqrt{n}} & 0 \\ 0 & 1 \end{pmatrix} (\iota_n \vdots ((\beta_1)_0) + \frac{1}{\sqrt{T}} \Psi_\beta)' (\iota_n \lambda_1 + ((\beta_1)_0) \lambda_F + \frac{1}{\sqrt{T}} \psi_R) = \\ &\begin{pmatrix} \frac{1}{\sqrt{n}} & 0 \\ 0 & 1 \end{pmatrix} \begin{pmatrix} \iota'_n \iota_n \lambda_1 + \iota'_{n_1} \beta_1 \lambda_F + \frac{1}{\sqrt{T}} \iota'_n \psi_R \\ \beta'_1 \iota_{n_1} \lambda_1 + \beta'_1 \beta_1 \lambda_F + \frac{1}{\sqrt{T}} \beta'_1 \psi_{R_1} + \frac{1}{\sqrt{T}} \Psi'_\beta \iota_n \lambda_1 + \frac{1}{\sqrt{T}} \Psi'_{\beta_1} \beta_1 \lambda_F + \frac{1}{T} \Psi'_\beta \Psi_R \end{pmatrix} = \\ &\begin{pmatrix} \sqrt{n} \lambda_1 + \frac{1}{\sqrt{n}} \beta'_1 \iota_{n_1} \lambda_F + \frac{1}{\sqrt{T}} \frac{1}{\sqrt{n}} \iota'_n \psi_R \\ \beta'_1 \iota_{n_1} \lambda_1 + n_1 B_1 \lambda_F + \frac{1}{\sqrt{T}} \beta'_1 \psi_{R_1} + \frac{1}{\sqrt{T}} \Psi'_\beta \iota_n \lambda_1 + \frac{1}{\sqrt{T}} \Psi'_{\beta_1} \beta_1 \lambda_F + \frac{1}{T} \Psi'_\beta \Psi_R \end{pmatrix} = \\ &\sqrt{n} \begin{pmatrix} 1 & \frac{1}{\sqrt{n}} \beta'_1 \iota_{n_1} + \frac{1}{\sqrt{nT}} \Psi'_\beta \iota_n \\ \frac{1}{\sqrt{n}} \beta'_1 \iota_{n_1} + \frac{1}{\sqrt{nT}} \Psi'_\beta \iota_n & \beta'_1 \beta_1 + \frac{n}{T} \frac{\Psi'_\beta \Psi_\beta}{n} + \frac{1}{\sqrt{T}} \Psi'_{\beta_1} \beta_1 + \frac{1}{\sqrt{T}} \beta'_1 \Psi_{\beta_1} \end{pmatrix} \lambda_1 + \begin{pmatrix} \frac{1}{\sqrt{n}} \beta'_1 \iota_{n_1} + \frac{1}{\sqrt{T}} \frac{1}{\sqrt{n}} \iota'_n \Psi_\beta \\ \beta'_1 \beta_1 + \frac{n}{T} \frac{\Psi'_\beta \Psi_\beta}{n} + \frac{1}{\sqrt{T}} \Psi'_{\beta_1} \beta_1 + \frac{1}{\sqrt{T}} \beta'_1 \Psi_{\beta_1} \end{pmatrix} \lambda_F + \\ &+ \begin{pmatrix} 0 \\ \frac{n}{T} \frac{\Psi'_\beta \Psi_\beta}{n} \end{pmatrix} (\mu_F - \lambda_F) - \begin{pmatrix} \frac{1}{\sqrt{T}} \frac{1}{\sqrt{n}} \iota'_n \Psi_\beta \\ \frac{n}{T} \frac{\Psi'_\beta \Psi_\beta}{n} + \frac{1}{\sqrt{T}} \beta'_1 \Psi_{\beta_1} \end{pmatrix} \lambda_F + \begin{pmatrix} \frac{1}{\sqrt{T}} \frac{1}{\sqrt{n}} \iota'_n \psi_{R_1} \\ \frac{1}{\sqrt{T}} \beta'_1 \psi_{R_1} + \frac{1}{T} \Psi'_\beta \Psi_R \end{pmatrix} = \\ &\sqrt{n} \begin{pmatrix} 1 \\ b_1 \end{pmatrix} \lambda_1 + \begin{pmatrix} b_1 \\ B_1 \end{pmatrix} \lambda_F + \begin{pmatrix} \frac{1}{\sqrt{T}} \frac{1}{\sqrt{n}} \iota'_n \Psi_\beta \\ \frac{n}{T} \frac{\Psi'_\beta \Psi_\beta}{n} + \frac{1}{\sqrt{T}} \beta'_1 \Psi_{\beta_1} \end{pmatrix} (\mu_F - \lambda_F) + \\ &\begin{pmatrix} \frac{1}{\sqrt{T}} \frac{1}{\sqrt{n}} \iota'_n \psi_{R_1} \\ \frac{1}{\sqrt{T}} \beta'_1 \psi_{R_1} + \frac{1}{T} \Psi'_\beta \Psi_R \end{pmatrix}, \end{aligned}$$

where $\Psi_{R_1} = (I_{n_1} \ : \ 0)\Psi_R$. The combined behavior for the estimator then results as

$$\begin{aligned}
\begin{pmatrix} \hat{\lambda}_1 \\ \hat{\lambda}_F \end{pmatrix} &= \begin{pmatrix} \frac{1}{\sqrt{n}} & 0 \\ 0 & 1 \end{pmatrix} \left[\begin{pmatrix} \frac{1}{\sqrt{n}} & 0 \\ 0 & 1 \end{pmatrix} [\iota_n \ : \ ((\beta_1) + \frac{1}{\sqrt{T}}\Psi_\beta)]' [\iota_n \ : \ ((\beta_1) + \frac{1}{\sqrt{T}}\Psi_\beta)] \begin{pmatrix} \frac{1}{\sqrt{n}} & 0 \\ 0 & 1 \end{pmatrix} \right]^{-1} \\
&\begin{pmatrix} \frac{1}{\sqrt{n}} & 0 \\ 0 & 1 \end{pmatrix} (\iota_n \ : \ ((\beta_1) + \frac{1}{\sqrt{T}}\Psi_\beta))' (\iota_n \lambda_1 + ((\beta_1) \lambda_F + \frac{1}{\sqrt{T}}\psi_R)) \\
&= \begin{pmatrix} \frac{1}{\sqrt{n}} & 0 \\ 0 & 1 \end{pmatrix} \begin{pmatrix} (1 - b'_1 B_1^{-1} b_1)^{-1} & -(1 - b'_1 B_1^{-1} b_1)^{-1} b'_1 B_1^{-1} \\ -(B_1 - b_1 b'_1)^{-1} b_1 & (B_1 - b_1 b'_1)^{-1} \end{pmatrix} \\
&\left[\sqrt{n} \begin{pmatrix} 1 \\ b_1 \end{pmatrix} \lambda_1 + \begin{pmatrix} b'_1 \\ B_1 \end{pmatrix} \lambda_F + \begin{pmatrix} 0 \\ \frac{n}{T} \frac{\Psi'_\beta \Psi_\beta}{n} \end{pmatrix} (\mu_F - \lambda_F) - \begin{pmatrix} \frac{1}{\sqrt{T}} \frac{1}{\sqrt{n}} \iota'_n \Psi_\beta \\ \frac{1}{\sqrt{T}} \beta'_1 \Psi_{\beta_1} \end{pmatrix} \lambda_F \right. \\
&\left. \begin{pmatrix} \frac{1}{\sqrt{T}} \frac{1}{\sqrt{n}} \iota'_n \psi_{R_1}^* \\ \frac{1}{\sqrt{T}} \beta'_1 \psi_{R_1}^* + \frac{1}{T} \Psi'_\beta \Psi_{R^*} \end{pmatrix} \right] \\
&= \begin{pmatrix} \lambda_1 \\ \lambda_F \end{pmatrix} + \begin{pmatrix} \frac{1}{\sqrt{n}} & 0 \\ 0 & 1 \end{pmatrix} \begin{pmatrix} (1 - b'_1 B_1^{-1} b_1)^{-1} & -(1 - b'_1 B_1^{-1} b_1)^{-1} b'_1 B_1^{-1} \\ -(B_1 - b_1 b'_1)^{-1} b_1 & (B_1 - b_1 b'_1)^{-1} \end{pmatrix} \left[\begin{pmatrix} 0 \\ \frac{n}{T} \frac{\Psi'_\beta \Psi_\beta}{n} \end{pmatrix} (\mu_F - \lambda_F) + \right. \\
&\left. - \begin{pmatrix} \frac{1}{\sqrt{T}} \frac{1}{\sqrt{n}} \iota'_n \Psi_\beta \\ \frac{1}{\sqrt{T}} \beta'_1 \Psi_{\beta_1} \end{pmatrix} \lambda_F \right] + \begin{pmatrix} \frac{1}{\sqrt{n}} & 0 \\ 0 & 1 \end{pmatrix} \begin{pmatrix} (1 - b'_1 B_1^{-1} b_1)^{-1} & -b'_1 (B_1 - b_1 b'_1)^{-1} \\ -(B_1 - b_1 b'_1)^{-1} b_1 & (B_1 - b_1 b'_1)^{-1} \end{pmatrix} \begin{pmatrix} \frac{1}{\sqrt{T}} \frac{1}{\sqrt{n}} \iota'_n \psi_R^* \\ \frac{1}{\sqrt{T}} \beta'_1 \psi_{R_1}^* + \frac{1}{T} \Psi'_\beta \Psi_{R^*} \end{pmatrix}.
\end{aligned}$$

When we now use that $\frac{n}{T} \frac{\Psi'_\beta \Psi_\beta}{n} \xrightarrow{p} W$, $\frac{1}{\sqrt{T}} \beta'_1 \Psi_{\beta_1} \xrightarrow{p} 0$, $b_1 \xrightarrow{p} 0$, $B_1 \xrightarrow{p} \beta'_1 \beta_1 + W$, we obtain that the bias of the FM risk premia estimator equals $(\beta'_1 \beta_1 + W)^{-1} W (\mu_F - \lambda_F)$.

D. Proof of Theorem 2.

1. $\beta = 0$. We use that since $(\iota_n \ : \ \frac{1}{\sqrt{T}}\Psi_\beta)' \Omega^{-1} (\iota_n \ : \ \frac{1}{\sqrt{T}}\Psi_\beta) = \begin{pmatrix} \iota'_n \Omega^{-1} \iota_n & \frac{1}{\sqrt{T}} \iota'_n \Omega^{-1} \Psi_\beta \\ \frac{1}{\sqrt{T}} \Psi'_\beta \Omega^{-1} \iota_n & \frac{1}{T} \Psi'_\beta \Omega^{-1} \Psi_\beta \end{pmatrix}$,

$$\begin{aligned}
&\left[(\iota_n \ : \ \frac{1}{\sqrt{T}}\Psi_\beta)' \Omega^{-1} (\iota_n \ : \ \frac{1}{\sqrt{T}}\Psi_\beta) \right]^{-1} = \\
&\begin{pmatrix} a^{-1} & -\sqrt{T} a^{-1} \iota'_n \Omega^{-1} \Psi_\beta (\Psi'_\beta \Omega^{-1} \Psi_\beta)^{-1} \\ -\sqrt{T} (\Psi'_\beta \Omega^{-\frac{1}{2}} M_{\Omega^{-\frac{1}{2}} \iota_n} \Omega^{-\frac{1}{2}} \Psi_\beta)^{-1} \frac{\Psi'_\beta \Omega^{-1} \iota_n}{\iota'_n \Omega^{-1} \iota_n} & T (\Psi'_\beta \Omega^{-\frac{1}{2}} M_{\Omega^{-\frac{1}{2}} \iota_n} \Omega^{-\frac{1}{2}} \Psi_\beta)^{-1} \end{pmatrix}.
\end{aligned}$$

with $a = \iota'_n \Omega^{-\frac{1}{2}} M_{\Omega^{-\frac{1}{2}} \Psi_\beta} \Omega^{-\frac{1}{2}} \iota_n$.

1a. $E(\bar{R}) = \iota_n \lambda_1$ and since $\hat{\Omega} \xrightarrow{p} \Omega$,

$$\begin{aligned}
\begin{pmatrix} \hat{\lambda}_1 \\ \hat{\lambda}_F \end{pmatrix} &= \left[(\iota_n \ : \ \hat{\beta})' \hat{\Omega}^{-1} (\iota_n \ : \ \hat{\beta}) \right]^{-1} (\iota_n \ : \ \hat{\beta})' \hat{\Omega} \bar{R} \\
&\stackrel{a}{=} \left[(\iota_n \ : \ \frac{1}{\sqrt{T}}\Psi_\beta)' \Omega^{-1} (\iota_n \ : \ \frac{1}{\sqrt{T}}\Psi_\beta) \right]^{-1} (\iota_n \ : \ \frac{1}{\sqrt{T}}\Psi_\beta)' \Omega^{-1} (\iota_n \lambda_1 + \frac{1}{\sqrt{T}}\psi_R) \\
&\stackrel{a}{=} \begin{pmatrix} a^{-1} & -\sqrt{T} a^{-1} \iota'_n \Omega^{-1} \Psi_\beta (\Psi'_\beta \Omega^{-1} \Psi_\beta)^{-1} \\ -\sqrt{T} (\Psi'_\beta \Omega^{-\frac{1}{2}} M_{\Omega^{-\frac{1}{2}} \iota_n} \Omega^{-\frac{1}{2}} \Psi_\beta)^{-1} \frac{\Psi'_\beta \Omega^{-1} \iota_n}{\iota'_n \Omega^{-1} \iota_n} & T (\Psi'_\beta \Omega^{-\frac{1}{2}} M_{\Omega^{-\frac{1}{2}} \iota_n} \Omega^{-\frac{1}{2}} \Psi_\beta)^{-1} \end{pmatrix} \\
&\begin{pmatrix} \iota'_n \Omega^{-1} \iota_n \lambda_1 + \frac{1}{\sqrt{T}} \iota'_n \Omega^{-1} \psi_R \\ \frac{1}{\sqrt{T}} \Psi'_\beta \Omega^{-1} \iota_n \lambda_1 + \frac{1}{T} \Psi'_\beta \Omega^{-1} \psi_R \end{pmatrix} \\
&\stackrel{a}{=} \begin{pmatrix} a^{-1} (a \lambda_1 + \frac{1}{\sqrt{T}} b' \Omega^{-\frac{1}{2}} \psi_R) \\ (\Psi'_\beta \Omega^{-\frac{1}{2}} M_{\Omega^{-\frac{1}{2}} \iota_n} \Omega^{-\frac{1}{2}} \Psi_\beta)^{-1} \Psi'_\beta \Omega^{-\frac{1}{2}} M_{\Omega^{-\frac{1}{2}} \iota_n} \Omega^{-\frac{1}{2}} \psi_R \end{pmatrix} \\
&\stackrel{a}{=} \begin{pmatrix} \lambda_1 \\ 0 \end{pmatrix} + \begin{pmatrix} \frac{1}{\sqrt{T}} \frac{1}{a} b' \\ (\Psi'_\beta \Omega^{-\frac{1}{2}} M_{\Omega^{-\frac{1}{2}} \iota_n} \Omega^{-\frac{1}{2}} \Psi_\beta)^{-1} \Psi'_\beta \Omega^{-\frac{1}{2}} M_{\Omega^{-\frac{1}{2}} \iota_n} \end{pmatrix} \Omega^{-\frac{1}{2}} \psi_R
\end{aligned}$$

where $\Psi_\beta = \text{vecinv}_k(\psi_\beta)$ and $b = M_{\Omega^{-\frac{1}{2}}\Psi_\beta} \Omega^{-\frac{1}{2}}\iota_n$, so

$$\begin{pmatrix} \sqrt{T}(\hat{\lambda}_1 - \lambda_1) \\ \hat{\lambda}_F \end{pmatrix} \xrightarrow{d} \begin{pmatrix} \frac{1}{\sqrt{T}}\frac{1}{a}b' \\ (\Psi'_\beta \Omega^{-\frac{1}{2}} M_{\Omega^{-\frac{1}{2}}\Psi_\beta} \Omega^{-\frac{1}{2}} \Psi_\beta)^{-1} \Psi'_\beta \Omega^{-\frac{1}{2}} M_{\Omega^{-\frac{1}{2}}\Psi_\beta} \end{pmatrix} \Omega^{-\frac{1}{2}} \psi_R.$$

Since ψ_R is a normal distributed random vector that is independent of Ψ_β . The random matrix $(\Psi'_\beta \Omega^{-\frac{1}{2}} M_{\Omega^{-\frac{1}{2}}\Psi_\beta} \Omega^{-\frac{1}{2}} \Psi_\beta)^{-1} \Psi'_\beta \Omega^{-\frac{1}{2}} M_{\Omega^{-\frac{1}{2}}\Psi_\beta}$ is the square root of the inverse of the Wishart distributed random matrix $(\Omega^{-\frac{1}{2}} \Psi_\beta)' M_{\Omega^{-\frac{1}{2}}\Psi_\beta} \Omega^{-\frac{1}{2}} \Psi_\beta$ which has a Wishart distributed matrix with $n - 1$ degrees of freedom and scale matrix Q_{FF} . Hence, it results from the proof of Theorem 1 that the limiting distribution of $\hat{\lambda}_F$ is a multi-variate t distribution with $n - k$ degrees of freedom, location 0 and scale matrix $Q_{11.F} Q_{FF}^{-1}$ with $Q_{11.F} = Q_{11} - Q_{1F} Q_{FF}^{-1} Q_{F1}$.

1b. When β is equal to zero and $E(\bar{R}) = c$:

$$\begin{aligned} \begin{pmatrix} \hat{\lambda}_1 \\ \hat{\lambda}_F \end{pmatrix} &= \left[(\iota_n : \hat{\beta})' \hat{\Omega}^{-1} (\iota_n : \hat{\beta}) \right]^{-1} (\iota_n : \hat{\beta})' \hat{\Omega} \bar{R} \\ &= \frac{1}{a} \left[(\iota_n : \frac{1}{\sqrt{T}} \Psi_\beta)' \Omega^{-1} (\iota_n : \frac{1}{\sqrt{T}} \Psi_\beta) \right]^{-1} (\iota_n : \frac{1}{\sqrt{T}} \Psi_\beta)' \Omega^{-1} (\iota_n \lambda_1 + \frac{1}{\sqrt{T}} \psi_R) \\ &= \frac{1}{a} \begin{pmatrix} a^{-1} & -\sqrt{T} a^{-1} \iota_n' \Omega^{-1} \Psi_\beta (\Psi'_\beta \Omega^{-1} \Psi_\beta)^{-1} \\ -\sqrt{T} (\Psi'_\beta \Omega^{-\frac{1}{2}} M_{\Omega^{-\frac{1}{2}}\Psi_\beta} \Omega^{-\frac{1}{2}} \Psi_\beta)^{-1} \frac{\Psi'_\beta \Omega^{-1} \iota_n}{\iota_n' \Omega^{-1} \iota_n} & T (\Psi'_\beta \Omega^{-\frac{1}{2}} M_{\Omega^{-\frac{1}{2}}\Psi_\beta} \Omega^{-\frac{1}{2}} \Psi_\beta)^{-1} \\ \left(\begin{array}{c} \iota_n' \Omega^{-1} c + \frac{1}{\sqrt{T}} \iota_n' \Omega^{-1} \psi_R \\ \frac{1}{\sqrt{T}} \Psi'_\beta \Omega^{-1} c + \frac{1}{T} \Psi'_\beta \Omega^{-1} \psi_R \end{array} \right) \end{pmatrix} \\ &= \frac{1}{a} \begin{pmatrix} a^{-1} \iota_n' \Omega^{-\frac{1}{2}} M_{\Omega^{-\frac{1}{2}}\Psi_\beta} \Omega^{-\frac{1}{2}} c + \frac{1}{\sqrt{T}} b' \Omega^{-\frac{1}{2}} \psi_R \\ (\Psi'_\beta \Omega^{-\frac{1}{2}} M_{\Omega^{-\frac{1}{2}}\Psi_\beta} \Omega^{-\frac{1}{2}} \Psi_\beta)^{-1} \Psi'_\beta \Omega^{-\frac{1}{2}} M_{\Omega^{-\frac{1}{2}}\Psi_\beta} \Omega^{-\frac{1}{2}} (\sqrt{T} c + \psi_R) \end{pmatrix} \\ &= \frac{1}{a} \begin{pmatrix} a^{-1} b' \Omega^{-\frac{1}{2}} c \\ \sqrt{T} (\Psi'_\beta \Omega^{-\frac{1}{2}} M_{\Omega^{-\frac{1}{2}}\Psi_\beta} \Omega^{-\frac{1}{2}} \Psi_\beta)^{-1} \Psi'_\beta \Omega^{-\frac{1}{2}} M_{\Omega^{-\frac{1}{2}}\Psi_\beta} \Omega^{-\frac{1}{2}} c \end{pmatrix} + \\ &\quad \begin{pmatrix} \frac{1}{\sqrt{T}} \frac{1}{a} b' \\ (\Psi'_\beta \Omega^{-\frac{1}{2}} M_{\Omega^{-\frac{1}{2}}\Psi_\beta} \Omega^{-\frac{1}{2}} \Psi_\beta)^{-1} \Psi'_\beta \Omega^{-\frac{1}{2}} M_{\Omega^{-\frac{1}{2}}\Psi_\beta} \end{pmatrix} \Omega^{-\frac{1}{2}} \psi_R \end{aligned}$$

and

$$\begin{pmatrix} \sqrt{T}(\hat{\lambda}_1 - a^{-1} b' \Omega^{-\frac{1}{2}} c) \\ \hat{\lambda}_F - \sqrt{T} (\Psi'_\beta \Omega^{-\frac{1}{2}} M_{\Omega^{-\frac{1}{2}}\Psi_\beta} \Omega^{-\frac{1}{2}} \Psi_\beta)^{-1} \Psi'_\beta \Omega^{-\frac{1}{2}} M_{\Omega^{-\frac{1}{2}}\Psi_\beta} \Omega^{-\frac{1}{2}} c \end{pmatrix} \xrightarrow{d} \begin{pmatrix} \frac{1}{\sqrt{T}} \frac{1}{a} b' \\ (\Psi'_\beta \Omega^{-\frac{1}{2}} M_{\Omega^{-\frac{1}{2}}\Psi_\beta} \Omega^{-\frac{1}{2}} \Psi_\beta)^{-1} \Psi'_\beta \Omega^{-\frac{1}{2}} M_{\Omega^{-\frac{1}{2}}\Psi_\beta} \end{pmatrix} \psi_R$$

and $\hat{\lambda}_F$ diverges when T gets large.

2a. $\beta = \frac{1}{\sqrt{T}}B$, with B a $n \times k$ matrix of full rank and $E(\bar{R}) = \iota_n \lambda_1 + \frac{1}{\sqrt{T}}B$ so

$$\begin{aligned}
\begin{pmatrix} \hat{\lambda}_1 \\ \hat{\lambda}_F \end{pmatrix} &= \left[(\iota_n : \hat{\beta})' \hat{\Omega}^{-1} (\iota_n : \hat{\beta}) \right]^{-1} (\iota_n : \hat{\beta})' \hat{\Omega}^{-1} \bar{R} \\
&\stackrel{a}{=} \left[(\iota_n : \frac{1}{\sqrt{T}}(B + \Psi_\beta))' \Omega^{-1} (\iota_n : \frac{1}{\sqrt{T}}(B + \Psi_\beta)) \right]^{-1} (\iota_n : \frac{1}{\sqrt{T}}(B + \Psi_\beta))' \Omega^{-1} (\iota_n \lambda_1 + \frac{1}{\sqrt{T}}(B \lambda_F + \psi_R)) \\
&\stackrel{a}{=} \begin{pmatrix} a^{-1} \\ -\sqrt{T}((B + \Psi_\beta)' \Omega^{-\frac{1}{2}} M_{\Omega^{-\frac{1}{2}} \iota_n} \Omega^{-\frac{1}{2}} (B + \Psi_\beta))^{-1} \frac{(B + \Psi_\beta)' \Omega^{-1} \iota_n}{\iota_n' \Omega^{-1} \iota_n} \\ -\sqrt{T} a^{-1} \iota_n' \Omega^{-1} (B + \Psi_\beta) ((B + \Psi_\beta)' \Omega^{-1} (B + \Psi_\beta))^{-1} \\ T((B + \Psi_\beta)' \Omega^{-\frac{1}{2}} M_{\Omega^{-\frac{1}{2}} \iota_n} \Omega^{-\frac{1}{2}} (B + \Psi_\beta))^{-1} \end{pmatrix} \\
&\quad \begin{pmatrix} \iota_n' \Omega^{-1} \iota_n \lambda_1 + \frac{1}{\sqrt{T}} \iota_n' \Omega^{-1} (B \lambda_F + \psi_R) \\ \frac{1}{\sqrt{T}} (B + \Psi_\beta)' \Omega^{-1} \iota_n \lambda_1 + \frac{1}{T} (B + \Psi_\beta)' \Omega^{-1} (B \lambda_F + \psi_R) \end{pmatrix} \\
&\stackrel{a}{=} \begin{pmatrix} a^{-1} (a \lambda_1 + \frac{1}{\sqrt{T}} a^{-1} b' \Omega^{-\frac{1}{2}} (B \lambda_F + \psi_R)) \\ ((B + \Psi_\beta)' \Omega^{-\frac{1}{2}} M_{\Omega^{-\frac{1}{2}} \iota_n} \Omega^{-\frac{1}{2}} (B + \Psi_\beta))^{-1} (B + \Psi_\beta)' \Omega^{-\frac{1}{2}} M_{\Omega^{-\frac{1}{2}} \iota_n} \Omega^{-\frac{1}{2}} (B \lambda_F + \psi_R) \end{pmatrix} \\
&\stackrel{a}{=} \begin{pmatrix} \lambda_1 \\ 0 \end{pmatrix} + \begin{pmatrix} \frac{1}{\sqrt{T}} \frac{1}{a} b' \\ ((B + \Psi_\beta)' \Omega^{-\frac{1}{2}} M_{\Omega^{-\frac{1}{2}} \iota_n} \Omega^{-\frac{1}{2}} (B + \Psi_\beta))^{-1} (B + \Psi_\beta)' \Omega^{-\frac{1}{2}} M_{\Omega^{-\frac{1}{2}} \iota_n} \end{pmatrix} \Omega^{-\frac{1}{2}} (B \lambda_F + \psi_R).
\end{aligned}$$

Hence,

$$\begin{pmatrix} \sqrt{T}(\hat{\lambda}_1 - \lambda_1) \\ \hat{\lambda}_F - \lambda_F \end{pmatrix} \xrightarrow{d} \begin{pmatrix} \frac{1}{\sqrt{T}} \frac{1}{a} b' \\ ((B + \Psi_\beta)' \Omega^{-\frac{1}{2}} M_{\Omega^{-\frac{1}{2}} \iota_n} \Omega^{-\frac{1}{2}} (B + \Psi_\beta))^{-1} (B + \Psi_\beta)' \Omega^{-\frac{1}{2}} M_{\Omega^{-\frac{1}{2}} \iota_n} \end{pmatrix} \Omega^{-\frac{1}{2}} (\psi_R - \Psi_\beta \lambda_F).$$

In this case $(B + \Psi_\beta)'(B + \Psi_\beta)$ has a non-central Wishart distribution regardless of the value of Ω . Hence, the moments of $\hat{\lambda}_F$ exist up to order $n - k$.

2b. $\beta = \frac{1}{\sqrt{T}}B$, with B a $n \times k$ matrix of full rank and $E(\bar{R}) = c$:

$$\begin{aligned}
\begin{pmatrix} \hat{\lambda}_1 \\ \hat{\lambda}_F \end{pmatrix} &= \left[(\iota_n : \hat{\beta})' \hat{\Omega}^{-1} (\iota_n : \hat{\beta}) \right]^{-1} (\iota_n : \hat{\beta})' \bar{R} \\
&\stackrel{a}{=} \left[(\iota_n : \frac{1}{\sqrt{T}}(B + \Psi_\beta))' \Omega^{-1} (\iota_n : \frac{1}{\sqrt{T}}(B + \Psi_\beta)) \right]^{-1} (\iota_n : \frac{1}{\sqrt{T}}(B + \Psi_\beta))' \Omega^{-1} (c + \frac{1}{\sqrt{T}}\psi_R) \\
&\stackrel{a}{=} \begin{pmatrix} a^{-1} \\ -\sqrt{T}((B + \Psi_\beta)' \Omega^{-\frac{1}{2}} M_{\Omega^{-\frac{1}{2}} \iota_n} \Omega^{-\frac{1}{2}} (B + \Psi_\beta))^{-1} \frac{(B + \Psi_\beta)' \Omega^{-1} \iota_n}{\iota_n' \Omega^{-1} \iota_n} \\ -\sqrt{T} a^{-1} \iota_n' \Omega^{-1} (B + \Psi_\beta) ((B + \Psi_\beta)' \Omega^{-1} (B + \Psi_\beta))^{-1} \\ T((B + \Psi_\beta)' \Omega^{-\frac{1}{2}} M_{\Omega^{-\frac{1}{2}} \iota_n} \Omega^{-\frac{1}{2}} (B + \Psi_\beta))^{-1} \end{pmatrix} \\
&\quad \begin{pmatrix} \iota_n' \Omega^{-1} c + \frac{1}{\sqrt{T}} \iota_n' \Omega^{-1} \psi_R \\ \frac{1}{\sqrt{T}} (B + \Psi_\beta)' \Omega^{-1} c + \frac{1}{T} (B + \Psi_\beta)' \Omega^{-1} \psi_R \end{pmatrix} \\
&\stackrel{a}{=} \begin{pmatrix} a^{-1} (b' \Omega^{-\frac{1}{2}} c + \frac{1}{\sqrt{T}} b' \Omega^{-\frac{1}{2}} \psi_R) \\ ((B + \Psi_\beta)' M_{\Omega^{-\frac{1}{2}} \iota_n} (B + \Psi_\beta))^{-1} (B + \Psi_\beta)' M_{\Omega^{-\frac{1}{2}} \iota_n} (\sqrt{T} c + \psi_R) \end{pmatrix} \\
&\stackrel{a}{=} \begin{pmatrix} \frac{1}{a} b' \Omega^{-\frac{1}{2}} c \\ \sqrt{T}((B + \Psi_\beta)' \Omega^{-\frac{1}{2}} M_{\Omega^{-\frac{1}{2}} \iota_n} \Omega^{-\frac{1}{2}} (B + \Psi_\beta))^{-1} (B + \Psi_\beta)' \Omega^{-\frac{1}{2}} M_{\Omega^{-\frac{1}{2}} \iota_n} \Omega^{-\frac{1}{2}} c \end{pmatrix} \\
&\quad + \begin{pmatrix} \frac{1}{\sqrt{T}} \frac{1}{a} b' \\ ((B + \Psi_\beta)' \Omega^{-\frac{1}{2}} M_{\Omega^{-\frac{1}{2}} \iota_n} \Omega^{-\frac{1}{2}} (B + \Psi_\beta))^{-1} (B + \Psi_\beta)' \Omega^{-\frac{1}{2}} M_{\Omega^{-\frac{1}{2}} \iota_n} \end{pmatrix} \Omega^{-\frac{1}{2}} \psi_R,
\end{aligned}$$

with $a = \iota'_n \Omega^{-\frac{1}{2}} M_{\Omega^{-\frac{1}{2}}(B+\Psi_\beta)} \Omega^{-\frac{1}{2}} \iota_n$, $b = M_{\Omega^{-\frac{1}{2}}(B+\Psi_\beta)} \Omega^{-\frac{1}{2}} \iota_n$, and

$$\begin{aligned} & \left(\begin{array}{c} \sqrt{T}(\hat{\lambda}_1 - \frac{1}{a} b' \Omega^{-\frac{1}{2}} c) \\ \hat{\lambda}_F - \sqrt{T}((B + \Psi_\beta)' \Omega^{-\frac{1}{2}} M_{\Omega^{-\frac{1}{2}} \iota_n} \Omega^{-\frac{1}{2}} (B + \Psi_\beta))^{-1} (B + \Psi_\beta)' \Omega^{-\frac{1}{2}} M_{\Omega^{-\frac{1}{2}} \iota_n} \Omega^{-\frac{1}{2}} c \end{array} \right) \\ & \xrightarrow{d} \left(\begin{array}{c} \frac{1}{\sqrt{T}} \frac{1}{a} b' \\ ((B + \Psi_\beta)' \Omega^{-\frac{1}{2}} M_{\Omega^{-\frac{1}{2}} \iota_n} \Omega^{-\frac{1}{2}} (B + \Psi_\beta))^{-1} (B + \Psi_\beta)' \Omega^{-\frac{1}{2}} M_{\Omega^{-\frac{1}{2}} \iota_n} \end{array} \right) \Omega^{-\frac{1}{2}} \psi_R. \end{aligned}$$

E. Proof of Lemma 2. Assumption 1 states that

$$\frac{1}{\sqrt{T}} \sum_{t=1}^T \left(\begin{array}{c} 1 \\ F_t \end{array} \right) \otimes (R_t - \iota_n \lambda_1 - \beta(\bar{F}_t + \lambda_F)) \xrightarrow{d} \begin{pmatrix} \varphi_R \\ \varphi_\beta \\ \varphi_F \end{pmatrix},$$

with $(\varphi'_R \ \varphi'_\beta \ \varphi'_F)' \sim N(0, V)$. We take all asset returns in deviation from the n -th asset return so

$$\frac{1}{\sqrt{T}} \sum_{t=1}^T \left(\begin{array}{c} 1 \\ F_t \end{array} \right) \otimes (\mathcal{R}_t - \mathcal{B}(\bar{F}_t + \lambda_F)) \xrightarrow{d} \begin{pmatrix} \varphi_{\mathcal{R}} \\ \varphi_{\mathcal{B}} \\ \psi_F \end{pmatrix},$$

with $\varphi_{\mathcal{R}} = J_n \varphi_R$, $\varphi_{\mathcal{B}} = (I_k \otimes J_n) \varphi_\beta$ with $J_n = (I_{N-1} \vdots -\iota_{n-1})$ so $(\varphi'_{\mathcal{R}} \vdots \varphi'_{\mathcal{B}} \vdots \varphi'_F)' \sim N(0, \text{diag}(Q \otimes \Sigma, V_{FF}))$ with $\Sigma = J_n \Omega J'_n$. We pre-multiply the result from Assumption 1 with

$$\left(\begin{array}{cc} 1 & 0 \\ \left[\frac{1}{T} \sum_{j=1}^T (\bar{F}_j + \lambda_{F,0}) (\bar{F}_j + \lambda_{F,0})' \right]^{-1} (\lambda_{F,0} - \bar{F}) & \left[\frac{1}{T} \sum_{j=1}^T (\bar{F}_j + \lambda_{F,0}) (\bar{F}_j + \lambda_{F,0})' \right]^{-1} \otimes I_n & 0 \\ 0 & & I_k \end{array} \right)$$

to obtain under $H_0 : \lambda_F = \lambda_{F,0}$ that

$$\sqrt{T} \begin{pmatrix} \bar{\mathcal{R}} - \mathcal{B} \lambda_F \\ \text{vec}(\tilde{\mathcal{B}} - \mathcal{B}) \\ \bar{F} - \mu_F \end{pmatrix} \xrightarrow{d} \begin{pmatrix} \xi_{\mathcal{R}} \\ \xi_{\mathcal{B}} \\ \xi_F \end{pmatrix},$$

where $\tilde{\mathcal{B}} = \sum_{t=1}^T \mathcal{R}_t (\bar{F}_t + \lambda_{F,0})' (\sum_{j=1}^T (\bar{F}_j + \lambda_{F,0}) (\bar{F}_j + \lambda_{F,0})')^{-1}$ and $(\xi'_{\mathcal{R}} \vdots \xi'_{\mathcal{B}} \vdots \xi'_F) \sim N(0, \text{diag}(Q(\lambda_F) \otimes \Sigma, V_{FF}))$ and

$$\begin{aligned} \mathcal{Q}(\lambda_F) &= \lim_{T \rightarrow \infty} E \left[\begin{pmatrix} 1 & \lambda'_F \left[\frac{1}{T} \sum_{t=1}^T (\bar{F}_t + \lambda_F) (\bar{F}_t + \lambda_F) \right]^{-1} \\ \left[\frac{1}{T} \sum_{t=1}^T (\bar{F}_t + \lambda_F) (\bar{F}_t + \lambda_F) \right]^{-1} \lambda_F & \left[\frac{1}{T} \sum_{t=1}^T (\bar{F}_t + \lambda_F) (\bar{F}_t + \lambda_F) \right]^{-1} \end{pmatrix} \right] \\ &= \begin{pmatrix} \mathcal{Q}(\lambda_F)_{11} & \mathcal{Q}(\lambda_F)_{1F} \\ \mathcal{Q}(\lambda_F)_{F1} & \mathcal{Q}(\lambda_F)_{FF} \end{pmatrix}. \end{aligned}$$

Hence, under $H_0 : \lambda_F = \lambda_{F,0}$, $\bar{\mathcal{R}} - \tilde{\mathcal{B}} \lambda_{F,0} = \bar{\mathcal{R}} - \mathcal{B} \lambda_F - (\tilde{\mathcal{B}} - \mathcal{B}) \lambda_F$ so $\sqrt{T}(\bar{\mathcal{R}} - \tilde{\mathcal{B}} \lambda_{F,0}) \xrightarrow{d} \xi_{\mathcal{R}} - (\lambda'_F \otimes I_{n-1}) \xi_{\mathcal{B}}$, which is independent of $\xi_{\mathcal{B}}$ since $\begin{pmatrix} 1 - \lambda'_{F,0} \\ 0 \end{pmatrix} \mathcal{Q}(\lambda_F) \begin{pmatrix} 1 - \lambda'_{F,0} \\ 0 \end{pmatrix}' = \text{diag}(1 - \lambda'_{F,0} \mathcal{Q}(\lambda_F)_{FF} \lambda_{F,0}, \mathcal{Q}(\lambda_F)_{FF})$, so

$$\sqrt{T} \begin{pmatrix} \bar{\mathcal{R}} - \tilde{\mathcal{B}} \lambda_F \\ \text{vec}(\tilde{\mathcal{B}} - \mathcal{B}) \\ \bar{F} - \mu_F \end{pmatrix} \xrightarrow{d} \begin{pmatrix} \psi_{\mathcal{R}} \\ \psi_{\mathcal{B}} \\ \psi_F \end{pmatrix},$$

where $\psi_{\mathcal{R}}$, $\psi_{\mathcal{B}}$ and ψ_F are independent $(n-1) \times 1$, $(n-1)k \times 1$ and $k \times 1$ normal distributed random vectors with mean zero and covariance matrices, $(1 - \lambda'_{F,0} \mathcal{Q}(\lambda_F)_{FF} \lambda_{F,0}) \otimes \Sigma$, $\mathcal{Q}(\lambda_F)_{FF} \otimes \Sigma$ and V_{FF} .

F. The derivative of the FAR statistic with respect to $\lambda_{F,0}$. To obtain the derivative of

$$\text{FAR}(\lambda_{F,0}) = \frac{T}{1 - \lambda'_{F,0} \hat{\mathcal{Q}}(\lambda_F)_{FF}^{-1} \lambda_{F,0}} (\bar{\mathcal{R}} - \tilde{\mathcal{B}} \lambda_{F,0})' \tilde{\Sigma}^{-1} (\bar{\mathcal{R}} - \tilde{\mathcal{B}} \lambda_{F,0}),$$

we first construct the derivatives of its different elements with respect to $\lambda_{F,0}$.

$$\begin{aligned} \frac{\partial \text{vec}(\tilde{\mathcal{B}})}{\partial \lambda'_{F,0}} &= \frac{\partial}{\partial \lambda'_{F,0}} \text{vec} \left[\sum_{t=1}^T \mathcal{R}_t (\bar{F}_t + \lambda_{F,0})' (\sum_{j=1}^T (\bar{F}_j + \lambda_{F,0}) (\bar{F}_j + \lambda_{F,0})')^{-1} \right] \\ &= \left[(\sum_{j=1}^T (\bar{F}_j + \lambda_{F,0}) (\bar{F}_j + \lambda_{F,0})')^{-1} \otimes I_{n-1} \right] \frac{\partial}{\partial \lambda'_{F,0}} \text{vec} \left[\sum_{t=1}^T \mathcal{R}_t (\bar{F}_t + \lambda_{F,0})' \right] \\ &\quad + \left[I_k \otimes \sum_{t=1}^T \mathcal{R}_t (\bar{F}_t + \lambda_{F,0})' \right] \frac{\partial}{\partial \lambda'_{F,0}} \text{vec} \left[(\sum_{j=1}^T (\bar{F}_j + \lambda_{F,0}) (\bar{F}_j + \lambda_{F,0})')^{-1} \right] \\ &= \left[(\sum_{j=1}^T (\bar{F}_j + \lambda_{F,0}) (\bar{F}_j + \lambda_{F,0})')^{-1} \otimes T \bar{\mathcal{R}} \right] \\ &\quad + \left[I_k \otimes \sum_{t=1}^T \mathcal{R}_t (\bar{F}_t + \lambda_{F,0})' \right] \frac{\partial}{\partial \lambda'_{F,0}} \text{vec} \left[(\sum_{j=1}^T (\bar{F}_j + \lambda_{F,0}) (\bar{F}_j + \lambda_{F,0})')^{-1} \right] \\ &= \left[(\sum_{j=1}^T (\bar{F}_j + \lambda_{F,0}) (\bar{F}_j + \lambda_{F,0})')^{-1} \otimes T \bar{\mathcal{R}} \right] \\ &\quad - 2 \left[(\sum_{j=1}^T (\bar{F}_j + \lambda_{F,0}) (\bar{F}_j + \lambda_{F,0})')^{-1} \otimes T \tilde{\mathcal{B}} \lambda_{F,0} \right] \\ &= T \left[(\sum_{j=1}^T (\bar{F}_j + \lambda_{F,0}) (\bar{F}_j + \lambda_{F,0})')^{-1} \otimes (\bar{\mathcal{R}} - \tilde{\mathcal{B}} \lambda_{F,0}) \right] \\ \frac{\partial}{\partial \lambda'_{F,0}} \left[1 - \lambda'_{F,0} \hat{\mathcal{Q}}(\lambda_F)_{FF}^{-1} \lambda_{F,0} \right] &= -2 \lambda'_{F,0} \left(\frac{1}{T} \sum_{j=1}^T (\bar{F}_j + \lambda_{F,0}) (\bar{F}_j + \lambda_{F,0})' \right)^{-1} - \\ &\quad (\lambda'_{F,0} \otimes \lambda'_{F,0}) \frac{\partial}{\partial \lambda'_{F,0}} \text{vec} \left[\left(\frac{1}{T} \sum_{j=1}^T (\bar{F}_j + \lambda_{F,0}) (\bar{F}_j + \lambda_{F,0})' \right)^{-1} \right] \\ &= -2 \lambda'_{F,0} \left(\frac{1}{T} \sum_{j=1}^T (\bar{F}_j + \lambda_{F,0}) (\bar{F}_j + \lambda_{F,0})' \right)^{-1}_{FF} \\ &\quad (1 - \lambda'_{F,0} \left(\frac{1}{T} \sum_{j=1}^T (\bar{F}_j + \lambda_{F,0}) (\bar{F}_j + \lambda_{F,0})' \right)^{-1} \lambda_{F,0}) \\ \frac{\partial \text{vec}(\tilde{\Sigma})}{\partial \lambda'_{F,0}} &= -2 \left(\frac{1}{T-k-1} \sum_{t=1}^T (\bar{\mathcal{R}}_t - \tilde{\mathcal{B}} (\bar{F}_t + \lambda_{F,0})) \otimes \tilde{\mathcal{B}} \right) \\ &\quad - 2 \left(\frac{1}{T-k-1} \sum_{t=1}^T (\bar{\mathcal{R}}_t - \tilde{\mathcal{B}} (\bar{F}_t + \lambda_{F,0})) (\bar{F}_t + \lambda_{F,0})' \otimes I_{n-1} \right) \\ &= -2 \left(\frac{1}{T-k-1} \sum_{t=1}^T (\bar{\mathcal{R}}_t - \tilde{\mathcal{B}} (\bar{F}_t + \lambda_{F,0})) \otimes \tilde{\mathcal{B}} \right) \\ &= -2 \left(\frac{T}{T-k-1} ((\bar{\mathcal{R}} - \tilde{\mathcal{B}} \lambda_{F,0}) \otimes \tilde{\mathcal{B}}) \right) \end{aligned}$$

since $\sum_{t=1}^T (\bar{\mathcal{R}}_t - \tilde{\mathcal{B}} (\bar{F}_t + \lambda_{F,0})) (\bar{F}_t + \lambda_{F,0})' = 0$ and $\bar{F} = 0$.

Hence,

$$\begin{aligned} \frac{\partial}{\partial \lambda'_{F,0}} \text{FAR}(\lambda_{F,0}) &= -2 \left\{ \frac{T}{1 - \lambda'_{F,0} \hat{\mathcal{Q}}(\lambda_F)_{FF}^{-1} \lambda_{F,0}} (\bar{\mathcal{R}} - \tilde{\mathcal{B}} \lambda_{F,0})' \tilde{\Sigma}^{-1} \tilde{\mathcal{B}} + \right. \\ &\quad \frac{T}{1 - \lambda'_{F,0} \hat{\mathcal{Q}}(\lambda_F)_{FF}^{-1} \lambda_{F,0}} (\lambda'_{F,0} \otimes (\bar{\mathcal{R}} - \tilde{\mathcal{B}} \lambda_{F,0})' \tilde{\Sigma}^{-1}) \frac{\partial \text{vec}(\tilde{\mathcal{B}})}{\partial \lambda'_{F,0}} \\ &\quad \frac{T}{(1 - \lambda'_{F,0} \hat{\mathcal{Q}}(\lambda_F)_{FF}^{-1} \lambda_{F,0})^2} (\bar{\mathcal{R}} - \tilde{\mathcal{B}} \lambda_{F,0})' \tilde{\Sigma}^{-1} (\bar{\mathcal{R}} - \tilde{\mathcal{B}} \lambda_{F,0}) \\ &\quad \frac{\partial}{\partial \lambda'_{F,0}} \left[1 - \lambda'_{F,0} \hat{\mathcal{Q}}(\lambda_F)_{FF}^{-1} \lambda_{F,0} \right] + \frac{T}{1 - \lambda'_{F,0} \hat{\mathcal{Q}}(\lambda_F)_{FF}^{-1} \lambda_{F,0}} \\ &\quad \left. ((\bar{\mathcal{R}} - \tilde{\mathcal{B}} \lambda_{F,0})' \tilde{\Sigma}^{-1} \otimes (\bar{\mathcal{R}} - \tilde{\mathcal{B}} \lambda_{F,0})' \tilde{\Sigma}^{-1}) \frac{\partial \text{vec}(\tilde{\Sigma})}{\partial \lambda'_{F,0}} \right\} \\ &= -2 \frac{T}{1 - \lambda'_{F,0} \hat{\mathcal{Q}}(\lambda_F)_{FF}^{-1} \lambda_{F,0}} \left\{ (\bar{\mathcal{R}} - \tilde{\mathcal{B}} \lambda_{F,0})' \tilde{\Sigma}^{-1} \tilde{\mathcal{B}} + \right. \\ &\quad \left[\lambda'_{F,0} \left(\frac{1}{T} \sum_{j=1}^T (\bar{F}_j + \lambda_{F,0}) (\bar{F}_j + \lambda_{F,0})' \right)^{-1} \otimes (\bar{\mathcal{R}} - \tilde{\mathcal{B}} \lambda_{F,0})' \tilde{\Sigma}^{-1} (\bar{\mathcal{R}} - \tilde{\mathcal{B}} \lambda_{F,0}) \right] \\ &\quad - \left[\lambda'_{F,0} \left(\frac{1}{T} \sum_{j=1}^T (\bar{F}_j + \lambda_{F,0}) (\bar{F}_j + \lambda_{F,0})' \right)^{-1}_{FF} (\bar{\mathcal{R}} - \tilde{\mathcal{B}} \lambda_{F,0})' \tilde{\Sigma}^{-1} (\bar{\mathcal{R}} - \tilde{\mathcal{B}} \lambda_{F,0}) \right] \\ &\quad \left. \left(\frac{T}{T-k-1} (\bar{\mathcal{R}} - \tilde{\mathcal{B}} \lambda_{F,0})' \tilde{\Sigma}^{-1} (\bar{\mathcal{R}} - \tilde{\mathcal{B}} \lambda_{F,0}) \otimes (\bar{\mathcal{R}} - \tilde{\mathcal{B}} \lambda_{F,0})' \tilde{\Sigma}^{-1} \tilde{\mathcal{B}} \right) \right\} \\ &= -2 \frac{T}{1 - \lambda'_{F,0} \hat{\mathcal{Q}}(\lambda_F)_{FF}^{-1} \lambda_{F,0}} \left[1 + \frac{1}{T-k-1} \text{FAR}(\lambda_{F,0}) \right] (\bar{\mathcal{R}} - \tilde{\mathcal{B}} \lambda_{F,0})' \tilde{\Sigma}^{-1} \tilde{\mathcal{B}} \end{aligned}$$

G. The specification of FAR($\lambda_{F,0}$) and CLR($\lambda_{F,0}$). After concentrating with respect to $B = (c : \mathcal{B})$, the log-likelihood of the unrestricted model

$$\mathcal{R}_t = B\left(\frac{1}{\bar{F}_t}\right) + \varepsilon_t,$$

under independent normal distributed disturbances with mean zero and covariance matrix Σ , is proportional to

$$\ln \mathcal{L}_{\text{unres}}(\Sigma) = c_\pi - \frac{T}{2} \ln |\Sigma| - \frac{1}{2} \text{tr} \left[\Sigma^{-1} \sum_{t=1}^T \left(\mathcal{R}_t - \hat{B}\left(\frac{1}{\bar{F}_t}\right) \right) \left(\mathcal{R}_t - \hat{B}\left(\frac{1}{\bar{F}_t}\right) \right)' \right],$$

with $\hat{B} = \sum_{t=1}^T \mathcal{R}_t \left(\frac{1}{\bar{F}_t}\right)' \left[\sum_{j=1}^T \left(\frac{1}{\bar{F}_j}\right) \left(\frac{1}{\bar{F}_j}\right)' \right]^{-1}$ and $c_\pi = -\frac{T}{2} \ln(2\pi)$. To determine the difference with the log-likelihood of the restricted model, we specify $\hat{B}\left(\frac{1}{\bar{F}_t}\right)$ as

$$\begin{aligned} \hat{B}\left(\frac{1}{\bar{F}_t}\right) &= \sum_{i=1}^T \mathcal{R}_i \left(\frac{1}{\bar{F}_i}\right)' \left[\sum_{j=1}^T \left(\frac{1}{\bar{F}_j}\right) \left(\frac{1}{\bar{F}_j}\right)' \right]^{-1} \left(\frac{1}{\bar{F}_t}\right) \\ &= \sum_{i=1}^T \mathcal{R}_i \left(\frac{1}{\bar{F}_i}\right)' \begin{pmatrix} 1 & 0 \\ \lambda_{F,0} & I_k \end{pmatrix}' \left[\begin{pmatrix} 1 & 0 \\ \lambda_{F,0} & I_k \end{pmatrix} \sum_{j=1}^T \left(\frac{1}{\bar{F}_j}\right) \left(\frac{1}{\bar{F}_j}\right)' \begin{pmatrix} 1 & 0 \\ \lambda_{F,0} & I_k \end{pmatrix} \right]^{-1} \begin{pmatrix} 1 & 0 \\ \lambda_{F,0} & I_k \end{pmatrix} \left(\frac{1}{\bar{F}_t}\right) \\ &= \sum_{i=1}^T \mathcal{R}_i \begin{pmatrix} 1 \\ \bar{F}_i + \lambda_{F,0} \end{pmatrix}' \left[\sum_{j=1}^T \begin{pmatrix} 1 \\ \bar{F}_j + \lambda_{F,0} \end{pmatrix} \begin{pmatrix} 1 \\ \bar{F}_j + \lambda_{F,0} \end{pmatrix}' \right]^{-1} \begin{pmatrix} 1 \\ \bar{F}_t + \lambda_{F,0} \end{pmatrix} \\ &= \sum_{i=1}^T \mathcal{R}_i \begin{pmatrix} 1 \\ \bar{F}_i + \lambda_{F,0} \end{pmatrix}' \left[\sum_{j=1}^T \begin{pmatrix} 1 \\ \bar{F}_j + \lambda_{F,0} \end{pmatrix} \begin{pmatrix} 1 \\ \bar{F}_j + \lambda_{F,0} \end{pmatrix}' \right]^{-1} \begin{pmatrix} 1 \\ \bar{F}_t + \lambda_{F,0} \end{pmatrix} \\ &= \sum_{i=1}^T \mathcal{R}_i \begin{pmatrix} 1 \\ \bar{F}_i + \lambda_{F,0} \end{pmatrix}' \left[\begin{pmatrix} 0 & 0 \\ 0 & \frac{1}{T} \hat{Q}_{FF}(\lambda_{F,0}) \end{pmatrix} + \begin{pmatrix} 1 \\ -\hat{Q}_{FF}(\lambda_{F,0}) \lambda_{F,0} \end{pmatrix} \right. \\ &\quad \left. \frac{1}{T} (1 - \lambda'_{F,0} \hat{Q}_{FF}(\lambda_{F,0}) \lambda_{F,0})^{-1} \begin{pmatrix} 1 \\ -\hat{Q}_{FF}(\lambda_{F,0}) \lambda_{F,0} \end{pmatrix}' \right] \begin{pmatrix} 1 \\ \bar{F}_t + \lambda_{F,0} \end{pmatrix} \\ &= \tilde{\mathcal{B}}(\bar{F}_t + \lambda_{F,0}) + (\bar{\mathcal{R}} - \tilde{\mathcal{B}} \lambda_{F,0}) (1 - \lambda'_{F,0} \hat{Q}_{FF}(\lambda_{F,0}) \lambda_{F,0})^{-1} \\ &\quad (1 - \lambda'_{F,0} \hat{Q}_{FF}(\lambda_{F,0}) (\bar{F}_t + \lambda_{F,0})), \end{aligned}$$

with $\hat{Q}_{FF}(\lambda_{F,0}) = \left[\frac{1}{T} \sum_{t=1}^T (\bar{F}_t + \lambda_{F,0}) (\bar{F}_t + \lambda_{F,0})' \right]^{-1}$, $\tilde{\mathcal{B}} = \frac{1}{T} \sum_{i=1}^T \mathcal{R}_i \begin{pmatrix} 1 \\ \bar{F}_i + \lambda_{F,0} \end{pmatrix}' \hat{Q}_{FF}(\lambda_{F,0})$. Hence,

$$\begin{aligned} \ln \mathcal{L}_{\text{unres}}(\Sigma) + \frac{T}{2} \ln |\Sigma| - c_\pi &= -\frac{1}{2} \text{tr} \left[\Sigma^{-1} \sum_{t=1}^T \left(\mathcal{R}_t - \hat{B}\left(\frac{1}{\bar{F}_t}\right) \right) \left(\mathcal{R}_t - \hat{B}\left(\frac{1}{\bar{F}_t}\right) \right)' \right] \\ &= -\frac{1}{2} \text{tr} \left[\Sigma^{-1} \sum_{t=1}^T \left(\mathcal{R}_t - \tilde{\mathcal{B}}(\bar{F}_t + \lambda_{F,0}) \right) \left(\mathcal{R}_t - \tilde{\mathcal{B}}(\bar{F}_t + \lambda_{F,0}) \right)' \right] - \frac{1}{2} \frac{1}{1 - \lambda'_{F,0} \hat{Q}_{FF}(\lambda_{F,0}) \lambda_{F,0}} \times \\ &\quad \left\{ -\text{tr} \left[\Sigma^{-1} \sum_{t=1}^T (\bar{\mathcal{R}} - \tilde{\mathcal{B}} \lambda_{F,0}) (1 - \lambda'_{F,0} \hat{Q}_{FF}(\lambda_{F,0}) (\bar{F}_t + \lambda_{F,0})) \left(\mathcal{R}_t - \tilde{\mathcal{B}}(\bar{F}_t + \lambda_{F,0}) \right)' \right] \right. \\ &\quad \left. - \text{tr} \left[\Sigma^{-1} \sum_{t=1}^T \left(\mathcal{R}_t - \tilde{\mathcal{B}}(\bar{F}_t + \lambda_{F,0}) \right) (1 - (\bar{F}_t + \lambda_{F,0})' \hat{Q}_{FF}(\lambda_{F,0}) \lambda_{F,0}) (\bar{\mathcal{R}} - \tilde{\mathcal{B}} \lambda_{F,0}) \right] \right. \\ &\quad \left. + \frac{1}{1 - \lambda'_{F,0} \hat{Q}_{FF}(\lambda_{F,0}) \lambda_{F,0}} \times \text{tr} \left[\Sigma^{-1} \sum_{t=1}^T (\bar{\mathcal{R}} - \tilde{\mathcal{B}} \lambda_{F,0}) (1 - \lambda'_{F,0} \hat{Q}_{FF}(\lambda_{F,0}) (\bar{F}_t + \lambda_{F,0})) \right. \right. \\ &\quad \left. \left. (1 - (\bar{F}_t + \lambda_{F,0})' \hat{Q}_{FF}(\lambda_{F,0}) \lambda_{F,0}) (\bar{\mathcal{R}} - \tilde{\mathcal{B}} \lambda_{F,0}) \right] \right\} \\ &= -\frac{1}{2} \text{tr} \left[\Sigma^{-1} \sum_{t=1}^T \left(\mathcal{R}_t - \tilde{\mathcal{B}}(\bar{F}_t + \lambda_{F,0}) \right) \left(\mathcal{R}_t - \tilde{\mathcal{B}}(\bar{F}_t + \lambda_{F,0}) \right)' \right] - \frac{1}{2} \frac{1}{1 - \lambda'_{F,0} \hat{Q}_{FF}(\lambda_{F,0}) \lambda_{F,0}} \times \\ &\quad \left\{ -\text{tr} \left[\Sigma^{-1} (\bar{\mathcal{R}} - \tilde{\mathcal{B}} \lambda_{F,0}) T (\bar{\mathcal{R}} - \tilde{\mathcal{B}} \lambda_{F,0} - \lambda'_{F,0} \tilde{\mathcal{B}}' + \lambda'_{F,0} \tilde{\mathcal{B}}') \right] \right. \\ &\quad \left. - \text{tr} \left[\Sigma^{-1} (\bar{\mathcal{R}} - \tilde{\mathcal{B}} \lambda_{F,0} - \tilde{\mathcal{B}} \lambda_{F,0} + \tilde{\mathcal{B}} \lambda_{F,0}) T (\bar{\mathcal{R}} - \tilde{\mathcal{B}} \lambda_{F,0})' \right] + \frac{1}{1 - \lambda'_{F,0} \hat{Q}_{FF}(\lambda_{F,0}) \lambda_{F,0}} \times \right. \\ &\quad \left. + \text{tr} \left[\Sigma^{-1} (\bar{\mathcal{R}} - \tilde{\mathcal{B}} \lambda_{F,0}) T (1 - \lambda'_{F,0} \hat{Q}_{FF}(\lambda_{F,0}) \lambda_{F,0}) (\bar{\mathcal{R}} - \tilde{\mathcal{B}} \lambda_{F,0})' \right] \right\} \\ &= -\frac{1}{2} \text{tr} \left[\Sigma^{-1} \sum_{t=1}^T \left(\mathcal{R}_t - \tilde{\mathcal{B}}(\bar{F}_t + \lambda_{F,0}) \right) \left(\mathcal{R}_t - \tilde{\mathcal{B}}(\bar{F}_t + \lambda_{F,0}) \right)' \right] + \frac{1}{2} \frac{T}{1 - \lambda'_{F,0} \hat{Q}_{FF}(\lambda_{F,0}) \lambda_{F,0}} \times \\ &\quad \text{tr} \left[\Sigma^{-1} (\bar{\mathcal{R}} - \tilde{\mathcal{B}} \lambda_{F,0}) (\bar{\mathcal{R}} - \tilde{\mathcal{B}} \lambda_{F,0})' \right] \end{aligned}$$

where we used that $\sum_{t=1}^T \bar{F}_t = 0$. Since the log-likelihood of the restricted model reads

$$\ln \mathcal{L}_{\text{res}}(\Sigma, \lambda_F) = c_\pi - \frac{T}{2} \ln |\Sigma| - \frac{1}{2} \text{tr} \left[\Sigma^{-1} \sum_{t=1}^T \left(\mathcal{R}_t - \tilde{\mathcal{B}}(\bar{F}_t + \lambda_{F,0}) \right) \left(\mathcal{R}_t - \tilde{\mathcal{B}}(\bar{F}_t + \lambda_{F,0}) \right)' \right],$$

the FAR statistic equals:

$$\text{FAR}(\lambda_{F,0}) = 2 [\ln \mathcal{L}_{\text{unres}}(\Sigma) - \ln \mathcal{L}_{\text{res}}(\Sigma, \lambda_F)].$$

CLR($\lambda_{F,0}$). We construct the maximal value of the likelihood of the restricted model. We therefore concentrate the likelihood with respect to the different parameters. The log-likelihood reads

$$\ln \mathcal{L}_{\text{res}}(\Sigma, \lambda_F, \mathcal{B}) = c_\pi - \frac{T}{2} \ln |\Sigma| - \frac{1}{2} \text{tr} \left[\Sigma^{-1} \sum_{t=1}^T \left(\mathcal{R}_t - \mathcal{B}(\bar{F}_t + \lambda_F) \right) \left(\mathcal{R}_t - \mathcal{B}(\bar{F}_t + \lambda_F) \right)' \right].$$

When we substitute the maximum likelihood estimator for \mathcal{B} into the expression of the likelihood, the concentrated likelihood of (Σ, λ_F) becomes

$$\ln \mathcal{L}_{\text{res}}(\Sigma, \lambda_F) = c_\pi - \frac{T}{2} \ln |\Sigma| - \frac{1}{2} \text{tr} \left[\Sigma^{-1} \sum_{t=1}^T \left(\mathcal{R}_t - \tilde{\mathcal{B}}(\bar{F}_t + \lambda_F) \right) \left(\mathcal{R}_t - \tilde{\mathcal{B}}(\bar{F}_t + \lambda_F) \right)' \right].$$

The maximum likelihood estimator of Σ is

$$\tilde{\Sigma} = \frac{1}{T} \sum_{t=1}^T \left(\mathcal{R}_t - \tilde{\mathcal{B}}(\bar{F}_t + \lambda_F) \right) \left(\mathcal{R}_t - \tilde{\mathcal{B}}(\bar{F}_t + \lambda_F) \right)'$$

and the concentrated likelihood of λ_F reads

$$\begin{aligned} \ln \mathcal{L}_{\text{res}}(\lambda_F) &= c_\pi - \frac{T}{2} \ln |\tilde{\Sigma}| - \frac{1}{2} \text{tr} [I_n T^{-1}] \\ &= c_T - \frac{T}{2} \ln |\tilde{\Sigma}|, \end{aligned}$$

with $c_T = c_\pi - \frac{n}{2T}$. The determinant of $\tilde{\Sigma}$ can be decomposed as

$$\begin{aligned} |\tilde{\Sigma}| &= \left| \frac{1}{T} \sum_{t=1}^T \left(\mathcal{R}_t - \tilde{\mathcal{B}}(\bar{F}_t + \lambda_F) \right) \left(\mathcal{R}_t - \tilde{\mathcal{B}}(\bar{F}_t + \lambda_F) \right)' \right| \\ &= \left| \frac{1}{T} \sum_{t=1}^T \mathcal{R}_t \mathcal{R}_t' - \left[\frac{1}{T} \sum_{t=1}^T \mathcal{R}_t (\bar{F}_t + \lambda_F)' \right] \left[\frac{1}{T} \sum_{t=1}^T (\bar{F}_t + \lambda_F) (\bar{F}_t + \lambda_F)' \right]^{-1} \right. \\ &\quad \left. \left[\frac{1}{T} \sum_{t=1}^T \mathcal{R}_t (\bar{F}_t + \lambda_F)' \right]' \right| \\ &= \left| \begin{pmatrix} \frac{1}{T} \sum_{t=1}^T \mathcal{R}_t \mathcal{R}_t' & \frac{1}{T} \sum_{t=1}^T \mathcal{R}_t (\bar{F}_t + \lambda_F)' \\ \frac{1}{T} \sum_{t=1}^T \mathcal{R}_t (\bar{F}_t + \lambda_F)' & \frac{1}{T} \sum_{t=1}^T (\bar{F}_t + \lambda_F) (\bar{F}_t + \lambda_F)' \end{pmatrix} \right| \\ &= \left| \begin{pmatrix} I_n & 0 \\ -\left[\frac{1}{T} \sum_{t=1}^T (\bar{F}_t + \lambda_F) (\bar{F}_t + \lambda_F)' \right]^{-1} \frac{1}{T} \sum_{t=1}^T \mathcal{R}_t (\bar{F}_t + \lambda_F)' & \left[\frac{1}{T} \sum_{t=1}^T (\bar{F}_t + \lambda_F) (\bar{F}_t + \lambda_F)' \right]^{-1} \end{pmatrix} \right| \\ &= \left| \begin{pmatrix} I_n & -\left[\frac{1}{T} \sum_{t=1}^T \mathcal{R}_t \mathcal{R}_t' \right]^{-1} \frac{1}{T} \sum_{t=1}^T \mathcal{R}_t (\bar{F}_t + \lambda_F)' \\ 0 & I_{kf} \end{pmatrix} \begin{pmatrix} \frac{1}{T} \sum_{t=1}^T \mathcal{R}_t \mathcal{R}_t' & \frac{1}{T} \sum_{t=1}^T \mathcal{R}_t (\bar{F}_t + \lambda_F)' \\ \frac{1}{T} \sum_{t=1}^T \mathcal{R}_t (\bar{F}_t + \lambda_F)' & \frac{1}{T} \sum_{t=1}^T (\bar{F}_t + \lambda_F) (\bar{F}_t + \lambda_F)' \end{pmatrix} \right| \\ &= \left| \begin{pmatrix} I_n & 0 \\ -\left[\frac{1}{T} \sum_{t=1}^T (\bar{F}_t + \lambda_F) (\bar{F}_t + \lambda_F)' \right]^{-1} \frac{1}{T} \sum_{t=1}^T \mathcal{R}_t (\bar{F}_t + \lambda_F)' & \left[\frac{1}{T} \sum_{t=1}^T (\bar{F}_t + \lambda_F) (\bar{F}_t + \lambda_F)' \right]^{-1} \end{pmatrix} \right| \\ &= \left| \frac{1}{T} \sum_{t=1}^T \mathcal{R}_t \mathcal{R}_t' \right| \left| \frac{1}{T} \sum_{t=1}^T (\bar{F}_t + \lambda_F) (\bar{F}_t + \lambda_F)' - \sum_{t=1}^T \frac{1}{T} (\bar{F}_t + \lambda_F) \mathcal{R}_t' \left[\frac{1}{T} \sum_{t=1}^T \mathcal{R}_t \mathcal{R}_t' \right]^{-1} \right. \\ &\quad \left. \frac{1}{T} \sum_{t=1}^T \mathcal{R}_t (\bar{F}_t + \lambda_F)' \right| \left| \frac{1}{T} \sum_{t=1}^T (\bar{F}_t + \lambda_F) (\bar{F}_t + \lambda_F)' \right|^{-1} \end{aligned}$$

such that the minimal value of $|\tilde{\Sigma}|$ as a function of λ_F results from the characteristic polynomial

$$\left| \theta \left[\frac{1}{T} \sum_{t=1}^T \begin{pmatrix} 1 \\ \bar{F}_t \end{pmatrix} \begin{pmatrix} 1 \\ \bar{F}_t \end{pmatrix}' \right] - \left[\frac{1}{T} \sum_{t=1}^T \begin{pmatrix} 1 \\ \bar{F}_t \end{pmatrix} \mathcal{R}_t' \right] \Sigma^{-1} \left[\frac{1}{T} \sum_{t=1}^T \mathcal{R}_t \begin{pmatrix} 1 \\ \bar{F}_t \end{pmatrix}' \right] \right| = 0,$$

where $\Sigma = \frac{1}{T} \sum_{t=1}^T \mathcal{R}_t \mathcal{R}'_t$. We express the characteristic polynomial as a function of $\bar{\mathcal{R}} - \tilde{\mathcal{B}}\lambda_{F,0}$ and $\tilde{\mathcal{B}}$ which does not affect the roots θ_i , $i = 1, \dots, k$.

$$\begin{aligned}
& \left| \theta \left[\frac{1}{T} \sum_{t=1}^T \begin{pmatrix} 1 \\ \bar{F}_t \end{pmatrix} \begin{pmatrix} 1 \\ \bar{F}_t \end{pmatrix}' \right] - \left[\frac{1}{T} \sum_{t=1}^T \begin{pmatrix} 1 \\ \bar{F}_t \end{pmatrix} \mathcal{R}'_t \right] \Sigma^{-1} \left[\frac{1}{T} \sum_{t=1}^T \mathcal{R}_t \begin{pmatrix} 1 \\ \bar{F}_t \end{pmatrix}' \right] \right| &= 0 \Leftrightarrow \\
& \left| \begin{pmatrix} 1 & 0 \\ \lambda_F & I_k \end{pmatrix} \left[\theta \frac{1}{T} \sum_{t=1}^T \begin{pmatrix} 1 \\ \bar{F}_t \end{pmatrix} \begin{pmatrix} 1 \\ \bar{F}_t \end{pmatrix}' - \frac{1}{T} \sum_{t=1}^T \begin{pmatrix} 1 \\ \bar{F}_t \end{pmatrix} \mathcal{R}'_t \Sigma^{-1} \frac{1}{T} \sum_{t=1}^T \mathcal{R}_t \begin{pmatrix} 1 \\ \bar{F}_t \end{pmatrix}' \right] \begin{pmatrix} 1 & 0 \\ \lambda_F & I_k \end{pmatrix}' \right| &= 0 \Leftrightarrow \\
& \left| \theta \begin{pmatrix} 1 & \lambda'_F \\ \lambda_F & \hat{Q}_{FF}(\lambda_F)^{-1} \end{pmatrix} - \left(\frac{1}{T} \sum_{t=1}^T \begin{pmatrix} \bar{\mathcal{R}}' \\ \bar{F}_t + \lambda_F \end{pmatrix} \mathcal{R}'_t \right) \Sigma^{-1} \left(\frac{1}{T} \sum_{t=1}^T \begin{pmatrix} \bar{\mathcal{R}}' \\ \bar{F}_t + \lambda_F \end{pmatrix} \mathcal{R}'_t \right)' \right| &= 0 \Leftrightarrow \\
& \left| \begin{pmatrix} 1 & -\lambda_F \hat{Q}_{FF}(\lambda_F) \\ 0 & I_k \end{pmatrix} \left[\theta \begin{pmatrix} 1 & \lambda'_F \\ \lambda_F & \hat{Q}_{FF}(\lambda_F) \end{pmatrix} - \left(\frac{1}{T} \sum_{t=1}^T \begin{pmatrix} \bar{\mathcal{R}} \\ \bar{F}_t + \lambda_F \end{pmatrix} \mathcal{R}'_t \right) \Sigma^{-1} \right. \right. \\
& \left. \left. \begin{pmatrix} \frac{1}{T} \sum_{t=1}^T \begin{pmatrix} \bar{\mathcal{R}} \\ \bar{F}_t + \lambda_F \end{pmatrix} \mathcal{R}'_t \right)' \begin{pmatrix} 1 & -\lambda_F \hat{Q}_{FF}(\lambda_F) \\ 0 & I_k \end{pmatrix}' \right] \right| &= 0 \Leftrightarrow \\
& \left| \theta \begin{pmatrix} (1-\lambda'_F \hat{Q}_{FF}(\lambda_F) \lambda_F) & 0 \\ 0 & \hat{Q}_{FF}(\lambda_F)^{-1} \end{pmatrix} - \left(\frac{1}{T} \sum_{t=1}^T \begin{pmatrix} \bar{\mathcal{R}} - \tilde{\mathcal{B}}\lambda_F' \\ \bar{F}_t + \lambda_F \end{pmatrix} \mathcal{R}'_t \right) \Sigma^{-1} \left(\frac{1}{T} \sum_{t=1}^T \begin{pmatrix} \bar{\mathcal{R}} - \tilde{\mathcal{B}}\lambda_F' \\ \bar{F}_t + \lambda_F \end{pmatrix} \mathcal{R}'_t \right)' \right| &= 0 \Leftrightarrow \\
& \left| \theta I_{k+1} - \left(\Sigma^{-\frac{1}{2}} (\bar{\mathcal{R}} - \tilde{\mathcal{B}}\lambda_F) \frac{1}{\sqrt{1-\lambda'_F \hat{Q}_{FF}(\lambda_F) \lambda_F}} \Sigma^{-\frac{1}{2}} \tilde{\mathcal{B}} \hat{Q}_{FF}(\lambda_F)^{-\frac{1}{2}} \right)' \right. \\
& \left. \left(\Sigma^{-\frac{1}{2}} (\bar{\mathcal{R}} - \tilde{\mathcal{B}}\lambda_F) \frac{1}{\sqrt{1-\lambda'_F \hat{Q}_{FF}(\lambda_F) \lambda_F}} \Sigma^{-\frac{1}{2}} \tilde{\mathcal{B}} \hat{Q}_{FF}(\lambda_F)^{-\frac{1}{2}} \right) \right| &= 0 \Leftrightarrow \\
& \left| \theta I_{k+1} - \frac{1}{T} \begin{pmatrix} \text{FAR}(\lambda_F) & s(\lambda_F)' \\ s(\lambda_F) & r(\lambda_F) \end{pmatrix} \right| &= 0,
\end{aligned}$$

with

$$\begin{aligned}
r(\lambda_F) &= T \hat{Q}_{FF}(\lambda_F)^{-\frac{1}{2}} \tilde{\mathcal{B}}' \left[\frac{1}{T} \sum_{t=1}^T \mathcal{R}_t \mathcal{R}'_t \right]^{-1} \tilde{\mathcal{B}} \hat{Q}_{FF}(\lambda_F)^{-\frac{1}{2}}, \\
s(\lambda_F) &= T \hat{Q}_{FF}(\lambda_F)^{-\frac{1}{2}} \tilde{\mathcal{B}}' \left[\frac{1}{T} \sum_{t=1}^T \mathcal{R}_t \mathcal{R}'_t \right]^{-1} (\bar{\mathcal{R}} - \tilde{\mathcal{B}}\lambda_F) \frac{1}{\sqrt{1-\lambda'_F \hat{Q}_{FF}(\lambda_F) \lambda_F}}.
\end{aligned}$$

When $k = 1$, the characteristic polynomial reads

$$\begin{aligned}
& \left| \theta I_{k+1} - \frac{1}{T} \begin{pmatrix} \text{FAR}(\lambda_F) & s(\lambda_F)' \\ s(\lambda_F) & r(\lambda_F) \end{pmatrix} \right| &= 0 \Leftrightarrow \\
& (\theta - \frac{1}{T} \text{FAR}(\lambda_F)) (\theta - \frac{1}{T} r(\lambda_F)) - s(\lambda_F) s(\lambda_F)' &= 0 \Leftrightarrow \\
& \theta^2 - \theta \frac{1}{T} (\text{FAR}(\lambda_F) + r(\lambda_F)) + \frac{1}{T^2} [\text{FAR}(\lambda_F) r(\lambda_F) - s(\lambda_F) s(\lambda_F)'] &= 0,
\end{aligned}$$

and the roots of the polynomial are characterized by

$$\theta = \frac{1}{2T} [\text{FAR}(\lambda_F) + r(\lambda_F) \pm \sqrt{(\text{FAR}(\lambda_F) + r(\lambda_F))^2 - 4r(\lambda_F) (\text{FAR}(\lambda_F) - \text{GLS-LM}(\lambda_F))}].$$

The largest root, which leads to the minimal value of $|\tilde{\Sigma}|$, is therefore equal to

$$\theta_{\max} = \frac{1}{2T} [\text{FAR}(\lambda_F) + r(\lambda_F) + \sqrt{(\text{FAR}(\lambda_F) + r(\lambda_F))^2 - 4r(\lambda_F) (\text{FAR}(\lambda_F) - \text{GLS-LM}(\lambda_F))}].$$

and the maximal value of the likelihood of the restricted model is

$$\begin{aligned}
\max_{\lambda_F} \ln \mathcal{L}_{\text{res}}(\Sigma, \lambda_F) &= c_T - \frac{T}{2} \ln [\min_{\lambda_F} |\Sigma|] \\
&= c_T - \frac{T}{2} \ln \left| \frac{1}{T} \sum_{t=1}^T \mathcal{R}_t \mathcal{R}'_t \right| - \frac{T}{2} \ln(1 - \theta_{\max}) \\
&\approx c_T - \frac{T}{2} \ln \left| \frac{1}{T} \sum_{t=1}^T \mathcal{R}_t \mathcal{R}'_t \right| + \frac{T}{2} \theta_{\max}.
\end{aligned}$$

To simplify the value of the likelihood under $H_0 : \lambda_F = \lambda_{F,0}$, we decompose the determinant of the estimator of Σ under H_0 :

$$\begin{aligned}
& \left| \frac{1}{T} \sum_{t=1}^T \left(\mathcal{R}_t - \tilde{\mathcal{B}}(\bar{F}_t + \lambda_{F,0}) \right) \left(\mathcal{R}_t - \tilde{\mathcal{B}}(\bar{F}_t + \lambda_{F,0}) \right)' \right| \\
&= \left| \frac{1}{T} \sum_{t=1}^T \mathcal{R}_t \mathcal{R}_t' - \left[\frac{1}{T} \sum_{t=1}^T \mathcal{R}_t (\bar{F}_t + \lambda_{F,0}) \right] \left[\frac{1}{T} \sum_{t=1}^T (\bar{F}_t + \lambda_{F,0}) (\bar{F}_t + \lambda_{F,0})' \right]^{-1} \right. \\
&\quad \left. \left[\frac{1}{T} \sum_{t=1}^T \mathcal{R}_t (\bar{F}_t + \lambda_{F,0}) \right]' \right| \\
&= \left| \frac{1}{T} \sum_{t=1}^T \mathcal{R}_t \mathcal{R}_t' \right| \left| \hat{Q}_{FF}(\lambda_{F,0}) \right| \\
&\quad \left| \hat{Q}_{FF}(\lambda_{F,0})^{-1} - \left[\frac{1}{T} \sum_{t=1}^T \mathcal{R}_t (\bar{F}_t + \lambda_{F,0})' \right]' \left[\frac{1}{T} \sum_{t=1}^T \mathcal{R}_t \mathcal{R}_t' \right]^{-1} \left[\frac{1}{T} \sum_{t=1}^T \mathcal{R}_t (\bar{F}_t + \lambda_{F,0})' \right] \right| \\
&= \left| \frac{1}{T} \sum_{t=1}^T \mathcal{R}_t \mathcal{R}_t' \right| \left| I_{k_f} - \hat{Q}_{FF}(\lambda_F)^{-\frac{1}{2}} \tilde{\mathcal{B}}' \left[\frac{1}{T} \sum_{t=1}^T \mathcal{R}_t \mathcal{R}_t' \right]^{-1} \tilde{\mathcal{B}} \hat{Q}_{FF}(\lambda_F)^{-\frac{1}{2}} \right| \\
&= \left| \frac{1}{T} \sum_{t=1}^T \mathcal{R}_t \mathcal{R}_t' \right| \left| I_{k_f} - \frac{1}{T} \mathbf{r}(\lambda_F) \right|.
\end{aligned}$$

When $k = 1$, we can therefore characterize the likelihood ratio statistic for testing $H_0 : \lambda_F = \lambda_{F,0}$ against $H_1 : \lambda_F \neq \lambda_{F,0}$ by

$$\begin{aligned}
& \text{LR}(\lambda_{F,0}) \\
&= 2 \left[\max_{\lambda_F} \ln \mathcal{L}_{\text{res}}(\lambda_F) - \ln \mathcal{L}_{\text{res}}(\lambda_{F,0}) \right] \\
&\approx 2 \left[c_T - \frac{T}{2} \ln \left| \frac{1}{T} \sum_{t=1}^T \mathcal{R}_t \mathcal{R}_t' \right| + \frac{T}{2} \theta_{\max} - c_T + \frac{T}{2} \ln \left(\left| \frac{1}{T} \sum_{t=1}^T \mathcal{R}_t \mathcal{R}_t' \right| \left| I_{k_f} - \frac{1}{T} \mathbf{r}(\lambda_F) \right| \right) \right] \\
&\approx 2 \left[c_T - \frac{T}{2} \ln \left| \frac{1}{T} \sum_{t=1}^T \mathcal{R}_t \mathcal{R}_t' \right| + \frac{T}{2} \theta_{\max} - c_T + \frac{T}{2} \ln \left| \frac{1}{T} \sum_{t=1}^T \mathcal{R}_t \mathcal{R}_t' \right| + \frac{T}{2} \ln \left| I_{k_f} - \frac{1}{T} \mathbf{r}(\lambda_F) \right| \right] \\
&\approx 2 \left[\frac{T}{2} \theta_{\max} - \frac{1}{2} \mathbf{r}(\lambda_F) \right] \\
&\approx \left[\text{FAR}(\lambda_{F,0}) - \mathbf{r}(\lambda_{F,0}) + \sqrt{(\text{FAR}(\lambda_{F,0}) + \mathbf{r}(\lambda_{F,0}))^2 - 4\mathbf{r}(\lambda_{F,0}) (\text{FAR}(\lambda_{F,0}) - \text{GLS-LM}(\lambda_{F,0}))} \right].
\end{aligned}$$

H. Size distortions of statistics that result from the inversion of a high dimensional covariance matrix. The expression of the GLS Wald statistic is

$$\text{GLS-W}(\lambda_{F,0}) = (\bar{R} - \iota_n \hat{\lambda}_1 - \hat{\beta} \lambda_{F,0})' \hat{\Omega}^{-1} \hat{\beta} \hat{\Theta}^{-1} \hat{\beta}' \hat{\Omega}^{-1} (\bar{R} - \iota_n \hat{\lambda}_1 - \hat{\beta} \lambda_{F,0}),$$

with $\hat{\Theta} = \hat{\beta}' \hat{\Omega}^{-1} \hat{\beta} - \hat{\beta} \hat{\Omega}^{-1} \iota_n (\iota_n' \hat{\Omega}^{-1} \iota_n)^{-1} \iota_n' \hat{\Omega}^{-1} \hat{\beta}$. The expression for $\hat{\Omega}$ is:

$$\hat{\Omega} = \frac{1}{T-k-1} \sum_{t=1}^T (\bar{R}_t - \hat{\beta} \bar{F}_t) (\bar{R}_t - \hat{\beta} \bar{F}_t)'.$$

To understand the size distortion of GLS-W, we specify it using matrix notation:

$$\text{GLS-W}(\lambda_{F,0}) = \left[\frac{1}{T} \iota_T' (R - \iota_T (\hat{\lambda}_1 \iota_n' + \lambda_{F,0}' \hat{\beta}') - \hat{F} \hat{\beta}') \right] \hat{\Omega}^{-1} \hat{\beta} \hat{\Theta}^{-1} \hat{\beta}' \hat{\Omega}^{-1} \left[\frac{1}{T} (R - \iota_T (\hat{\lambda}_1 \iota_n' + \lambda_{F,0}' \hat{\beta}') - \hat{F} \hat{\beta}')' \iota_T \right],$$

with $R = (R_1' \dots R_T')'$, $\hat{F} = (\bar{F}_1' \dots \bar{F}_T')'$, so $\iota_T' \hat{F} = 0$, and $\hat{\Omega} = \frac{1}{T-k-1} R' M_{(\iota_T \hat{F})} R$. When T equals n , $\hat{\Omega}$ is a $T \times T$ matrix of rank $T - 2$ since $M_{(\iota_T \hat{F})}$ is a $T \times T$ matrix of rank $T - 2$. The eigenvectors associated with the zero eigenvalues of $M_{(\iota_T \hat{F})}$ are spanned by ι_T and \hat{F} . The eigenvectors associated with the zero eigenvalues of $\hat{\Omega}$ are obtained by specifying R as

$$R = P_{(\iota_T \hat{F})} R + M_{(\iota_T \hat{F})} R.$$

The first part of this specification of R , $P_{(\iota_T \hat{F})} R$, is mapped onto zero when R is multiplied by $M_{(\iota_T \hat{F})}$. When $T = n$, the eigenvectors that are associated with a zero value of $\hat{\Omega}$ are therefore spanned by

$R'(\iota_T \hat{F})$ and the eigenvectors that belong to the non-zero eigenvalues are spanned by the orthogonal complement of $R'(\iota_T \hat{F})$. For large values of n smaller than T , $\hat{\Omega}$ is non-singular but will have two relatively small eigenvalues whose eigenvectors are spanned by $R'(\iota_T \hat{F})$. In the GLS-Wald statistic, $\hat{\Omega}^{-1}$ is post-multiplied by $(R - \iota_T(\hat{\lambda}_1 + \hat{\beta}'\lambda_{F,0}) - \hat{F}\hat{\beta}')'\iota_T$. This vector lies in the span of $R'(\iota_T \hat{F})$ which is the eigenvector associated with the smallest eigenvalues of $\hat{\Omega}$ or put differently the largest eigenvalues of $\hat{\Omega}^{-1}$. This shows that the large size distortion of the GLS-Wald statistics results from the association between $(R - \iota_T(\hat{\lambda}_1 + \hat{\beta}'\lambda_{F,0}) - \hat{F}\hat{\beta}')'\iota_T$ and the eigenvectors that belong to the largest eigenvalues of $\hat{\Omega}^{-1}$.

When we use the covariance matrix estimator,

$$\tilde{\Sigma} = \frac{1}{T-k-1} \sum_{t=1}^T (\mathcal{R}_t - \tilde{\mathcal{B}}(\bar{F}_t + \lambda_{F,0}))(\mathcal{R}_t - \tilde{\mathcal{B}}(\bar{F}_t + \lambda_{F,0}))',$$

the matrix expression of this covariance matrix estimator reads

$$\hat{\Omega} = \frac{1}{T-k-1} \mathcal{R}' M_{\hat{F} + \iota_T \lambda'_{F,0}} \mathcal{R}.$$

with $\mathcal{R} = (\mathcal{R}'_1 \dots \mathcal{R}'_T)'$. Since $\iota'_T(\mathcal{R} - (\hat{F} + \iota_T \lambda_{F,0})\tilde{\mathcal{B}}')$ is not spanned by $\mathcal{R}'(\hat{F} + \iota_T \lambda'_{F,0})$, $\iota'_T(\mathcal{R} - (\hat{F} + \iota_T \lambda_{F,0})\tilde{\mathcal{B}}')$ is not associated with the eigenvector of the largest eigenvalue of $\tilde{\Sigma}^{-1}$ and therefore all statistics that use $\tilde{\Sigma}^{-1}$ perform appropriately despite the inversion of a large dimensional matrix.

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