Estimation in High-Dimensional Factor Models with Structural Instabilities

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Estimation in High-Dimensional Factor Models with Structural Instabilities

by

Wen (Wendy) Gao

A Major Research Paper
Submitted to the Faculty of Graduate Studies through the Department of Mathematics and Statistics in Partial Fulfillment of the Requirements for the Degree of Master of Science at the University of Windsor

Windsor, Ontario, Canada

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ESTIMATION IN HIGH-DIMENSIONAL FACTOR MODELS WITH STRUCTURAL INSTABILITIES

by

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September 18, 2018
Author’s Declaration of Originality

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Abstract

In this major paper, we use high-dimensional models to analyze macroeconomic data which is influenced by the break point. In particular, we consider to detect the break point and study the changes of the number of factors and the factor loadings with the structural instability.

Concretely, we propose two factor models which explain the processes of pre- and post- break periods. Then, we consider the break point as known or unknown. In both situations, we derive the shrinkage estimators by minimizing the penalized least square function and calculate the estimators of the numbers of pre- and post- break factors and the existence of the break point. After that, we present some results about the asymptotic performance of the penalty least square estimators with both the cross-section dimension and the time dimension tend to infinity.

In addition, we establish the Monte Carlo simulation to evaluate the performance of the procedure and we analyze the real dataset from 2007-2009 Great Recession. The theoretical results are confirmed by simulation that the break point can be properly detected. More than half proposed post model selection estimators domain the full sample estimators while the procedure performs relatively poor in estimating the number of pre- and post- factors.
To my loving parents

Yun Gao and Bingmin Lian
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Chapter 1

Introduction

1.1 Background

This major paper is about estimation and modeling of macroeconomic data in a high-dimensional setting. In macroeconomics, it is common to study the economic system and its influence on the financial market in a long-term. In particular, the aim of studying macroeconomics is to understand the reason for the economic fluctuations, to achieve persistent economic growth and to maximize the level of employment and Gross Domestic Product (GDP) in a country (Giannone et al. 2008)[20].

The macroeconomics phenomena that we studied in this major paper is under the high-dimensional setting. Briefly, high-dimensional data refers to the situation where the number of unknown parameters is much larger than the sample size. Such a model is common in different areas of macroeconomics (Fan et al. 2011)[17]. Vector autoregressive (VAR) model (Sims et al. 1990)[29] is one example of the high
CHAPTER 1. INTRODUCTION

dimensional models, which is considered to be an effective method to evaluate the joint evolution of macroeconomic time series. However, macroeconomic problems are usually associated with hundreds of data series, which may result in quadratic or cubic growth in the number of parameters. To reduce the difficulty of estimating large number of parameters, Bernanke (2005)[7] proposed an augment standard VARs with estimated factors (FAVAR) to evaluate the consequences of monetary policy (Fan et al. 2011)[17]. Besides that, another difficulty of analyzing high dimensional data is that we have to cope with measures of uncertainty or stability (Bühlmann and Van de Geer, 2011)[11]. The additional difficulty we focus on in this major paper is the potential break of the economic or financial period. Indeed, the analysis of macroeconomics can be interrupted by structural breaks such as the 2007-2009 Great Recession. However, the structural breaks is generally unknown, which makes the statistical analysis more complicated.

In terms of a real-world application, many intriguing questions emerge: whether the Great Recession or other post-war U.S. recessions cause a change in the business cycle or not, are these breaks always accompanied by the emergence of new factors such as financial or credit factors, or are there any changes to the existing factors? (Cheng et al. 2016)[14]. This major paper is generated to answer these questions. It locates the break point if it is unknown, captures any new factors and estimates the change of the factors after the break point.
1.2 Existing Studies

Throughout the research, we reviewed the existing papers which use different techniques to analyze the problems mentioned above. The methods include locating the break point, evaluating changes of factor loadings and estimating the number of factors before and after the break point. For example, Bai (1997)[4] uses conventional residual-based procedures to estimate the break point, but his method is based on the known number of factors. Other studies given in Bai and Ng (2002)[5], Onatski (2010)[26], or Ahn and Horenstein (2013)[1] require the break point to be known and they are not able to capture the changes in factor loadings. Other works are given in Breitung and Eickmeier (2011)[10], Chen et al. (2014)[13], or Han and Inoue (2014)[22]. They all neglect the estimation of the number of factors as well as the emergence of the new factors in the process of structural break tests. Moreover, the methods in Stock and Watson (2012)[31] cannot work without the knowledge of the break point.

1.3 Improvements

Cheng et al.(2016)[14] improved over the references mentioned above by dealing with the situation that the break point is unknown. The proposed method simultaneously estimates the break point if it is unknown and determines the numbers of pre- and post-break factors. By the improved techniques, the overestimation of the number of factors is avoided. In this case, a factor model instability is considered. The definition of the instability refers to the large changes in the factor loadings when the number
of factors is constant.

The method used in this major paper applies two identification results for the factor model. First, this major paper shows that if the space spanned by the factor loadings or the scaling of the factor loadings changes, then a structural change is identifiable. Second, if the break point is unknown, then it could be determined by the factor model. Then, the number of pre- and post-break factors can be estimated more precisely under the knowledge of the break point.

The penalized least squares (PLS) method is used to obtain the estimator. The group-Least Absolute Shrinkage and Selection Operator (LASSO) penalties are applied to pre-break factor loadings and the change in loadings. A PLS criterion function is developed, and the estimator is generated from the minimization of this function.

1.4 Organization of the Major Paper

The remainder of this major paper is organized as follows. Chapter 2 introduces two versions of the factor models with the structural instability characterized by the knowledge of the break point. Chapter 3 demonstrates the whole process of deriving and calculating the shrinkage estimators, the mathematical algorithm applied to improve our work and the asymptotic result for the large sample behavior of the factors under the known break point case. The situation when the break point is unknown is explained in Chapter 4. Chapter 5 is separated into two parts. Part one displays
the design and the result of the Monte Carlo Simulations about the performance of our estimators. Part two of Chapter 5 implements our procedure to real world macroeconomic problem, the Great Recession. Finally, the result of our findings are concluded in Chapter 6. The proof of one major theorem is attached in the Appendix.
Chapter 2

The Statistical Model

In this Chapter, Section 2.1 gives an overview about the two versions of the factor models which are designed for pre- and post-break periods respectively. Section 2.2 briefly explains the identification of a structural change under known or unknown cases.

Suppose that \( t \in \{1, 2, \cdots, T_0, \cdots, T\} \), which denotes a set of \( T \) time points. \( T_0 \) is the only break point which is usually unknown. Suppose that there are \( r_a \) unobserved pre-break factors at \( T_0 \), and \( r_b \) unobserved post-break factors after \( T_0 \). We assume we have the data \( \{X_{it} \in \mathbb{R} : i = 1, \cdots, N, t = 1, \cdots, T\} \) for the \( i \)-th cross-section unit at time \( t \). Note that \( N \) and \( T \) are the cross-section dimension and the time dimension of the data set respectively (Bai and Ng, 2002)[5]. \( X_t = (X_{1t}, \cdots, X_{Nt})' \in \mathbb{R}^{N \times 1} \) are the observations at time \( t \). Two different versions of factor model are generated for the periods before and after the break point which will be discussed in the following Section.
CHAPTER 2. THE STATISTICAL MODEL

2.1 The factor Model

This section introduces the formation of the factor models with the observations $X$, the unknown factor loadings and the unknown parameters. The models have two different versions for the data observed before or after the break point.

The whole process is split into two parts by the potential structural break point. Let $F_a = (F^0_1, \cdots, F^0_T_0) \in \mathbb{R}^{T_0 \times r_a}$ denote the factors in the pre-break period, and let $\Lambda^0$ stand for an $N \times r_a$ matrix which denotes the factor loadings associated with $F_a$. For the time points at or before $T_0$, the model of pre-break period is

$$X_a = F_a \Lambda^0 + e_a,$$  \hspace{1cm} (2.1)

where $X_a = (X_1, \cdots, X_{T_0})' \in \mathbb{R}^{T_0 \times N}$, and $e_a = (e_1, \cdots, e_{T_0})' \in \mathbb{R}^{T_0 \times N}$, which is the error term. The factors, their loadings and the errors are not observable.

For the post-break model, the matrix $F_b$ are separated into $F_{b,1}$ and $F_{b,2}$, i.e., $F_b = (F_{b,1}, F_{b,2})$. $F_{b,1} \in \mathbb{R}^{T_1 \times r_a}$ stands for the pre-break factors in the post-break period, and $F_{b,2} \in \mathbb{R}^{T_1 \times (r_b-r_a)}$ stands for the new factors that emerge after the break point. Let $\Gamma^0 = (\Gamma^0_1, \Gamma^0_2)$ denote the change in the factor loadings, let $\Gamma^0_1$ denote the change of the pre-break factors $F^0_t$ and let $\Gamma^0_2$ denote the change of the new factors $F^*_t$. The matrix form of post-break model is given as follows

$$X_b = F_{b,1}(\Lambda^0 + \Gamma^0_1)' + F_{b,2}\Gamma^0_2 + e_b,$$  \hspace{1cm} (2.2)
where $X_b = (X_{T_0+1}, \cdots, X_T)'$, $F_{b,1} = (F^0_{T_0+1}, \cdots, F^0_{T_0})'$, $F_{b,2} = (F^*_{T_0+1}, \cdots, F^*_T)'$ and $e_b = (e_{T_0+1}, \cdots, e_T)'$, which is the error term.

### 2.2 Instability

In this section, we introduce the structural instability. In this major paper, we assume that the number of post-break factors is always greater than or equal to pre-break factors, which means $r_b \geq r_a$.

The structural instability is defined as following:

$$r_a = r_b \text{ and } \Gamma^0_1 \neq 0. \quad (2.3)$$

Under the instability, the number of factors remains the same after the break point, but the factor loadings changes.
Chapter 3

Preliminary Results: Known Break Point Case

In the model we discussed in the previous chapter, there are four critical parts, the structural instability, the break point, the change of factor loadings, and the number of factors. Above all, we analyze the instability, the change of factor loadings, and the number of factors given that the break point is known. First, we show the identification of the instability and then introduce the shrinkage estimator. Next, we show how we obtain the estimators of the occurrence of break point and the number of factors, which are determined by the shrinkage estimator. Then, we apply a two-step estimation algorithm to analyze the shrinkage estimator with adjusted penalty weights done by cross validation. Finally, we present the asymptotic theory for the large sample behavior of the factors.

In the first Section, we briefly explain the way to identify the instability. In Section
3.1 Identification of the Instability

Since the break point is known, the number of pre- and post-break factors $r_a$ and $r_b$ can be analyzed using the model selection method constructed by Bai and Ng (2002)[5].

To identify the instability, first, we need to have the knowledge of the normalization of the factor covariance matrix. Second, we show that either the space spanned by the factor loadings or the scaling of the factor loadings changes (Cheng et al. 2016)[14].

We define an augmented covariance matrix with dimension $(r_a + r_b) \times (r_a + r_b)$:

$$
\Sigma_{\Lambda \Psi}^+ = \begin{bmatrix}
\Sigma_{\Lambda} & \Sigma_{\Lambda \Psi} \\
\Sigma'_{\Lambda \Psi} & \Sigma_{\Psi}
\end{bmatrix}.
$$

We can identify the structural instability by using the following Assumption 1 from Cheng et al. (2016)[14].
Assumption 1. *One of the following two conditions holds:*

(i) \( \text{rank}(\Sigma_{A\Psi}^+) > r_a \);

(ii) \( \rho_\ell(\Sigma_F \Sigma_A) \neq \rho_\ell(\Sigma_F \Sigma_\Psi) \) for some \( \ell \leq r_a \).

Here, \( \rho_\ell(A) \) denotes the \( \ell \)-th largest eigenvalue of a square matrix \( A \), and \( \text{rank}(A) \) denotes the rank of an \( m \times n \) matrix \( A \), which refers to the number of pivot columns in an echelon of \( A \). Here, pivot columns refer to the columns which have entries distinct from 0.

### 3.2 Shrinkage Estimation

Shrinkage estimation is a statistical algorithm which aims to stabilize the estimates, reduce errors, and smooth fluctuations by narrowing the range of extreme values towards the sample mean (Xie et al. 2016)[33]. We also refer to Nkurunziza et al. (2016)[25] for more details about shrinkage and LASSO estimators. In this section, we introduce the purpose of using shrinkage estimator and how it works for our research. We also explain how to obtain the shrinkage estimator by achieving the minimization of the penalized least square (PLS) functions in details.

First we begin with introducing some special cases of the factor model. After the break point, if the loadings of the pre-break factors remain the same, then \( \Gamma_1^0 = 0 \); while if no new factors appear, then \( \Gamma_2^0 = 0 \). Suppose that the post-break factor loadings is \( \Psi^0 = (\Lambda^0 + \Gamma_1^0, \Gamma_2^0) \), which represents the loadings before and after break
CHAPTER 3. PRELIMINARY RESULTS: KNOWN BREAK POINT CASE

point, then the model in (2.2) can be rewritten as

$$X_b = F_b \Psi^0 + e_b. \quad (3.2)$$

In (2.1) and (3.2), the factors and their loadings have to be estimated together. So the normalization restrictions are applied to estimate the factor model (Cheng et al. 2016)[14]. The factor model in (2.1) and (3.2) can be rewritten as

$$X_a = (F_a R_a)(R_a^{-1} \Lambda^0) + e_a = F_a^R \Lambda^R' + e_a,$$
$$X_b = (F_b R_b)(R_b^{-1} \Psi^0) + e_b = F_b^R \Psi^R' + e_b,$$

where $F_a^R = F_a R_a$ and $F_b^R = F_b R_b$ and $R_a$ and $R_b$ are transformation matrices.

Suppose that $r_a + r_b \leq k$ where $k$ is an upper bound of the number of factors. The factor model in (3.3) is rewritten as the following augmented system (Cheng et al. 2016)[14].

$$X_a = \begin{bmatrix} F_a^R & F_{a,1}^R & F_{a,2}^R \end{bmatrix} \begin{bmatrix} \Lambda^R \\ 0_{(r_b-r_a)\times N} \\ 0_{(k-r_b)\times N} \end{bmatrix} + e_a = F_a^{R+}(\Lambda^{R+})' + e_a,$$
$$X_b = \begin{bmatrix} F_{b,1}^R & F_{b,2}^R & F_b^R \end{bmatrix} \begin{bmatrix} \Lambda^R + \Gamma^R \\ \Gamma^R_2 \\ 0_{(k-r_b)\times N} \end{bmatrix} + e_b = F_b^{R+}(\Lambda^{R+} + \Gamma^{R+})' + e_b.$$

The matrix $F_a^{R\perp}$ is an orthogonal complement of $F_a^R$, and it can be partitioned into
two sub-matrices $F_{a,1}^{R \perp} \in R^{T_0 \times (r_b-r_a)}$ and $F_{a,2}^{R \perp} \in R^{T_0 \times (k-r_b)}$. Similarly, $F_{b}^{R \perp}$ denotes an orthogonal complement of $F_{b}^{R} \in R^{T_b \times r_b}$ and $F_{b}^{R}$ can be partitioned into two submatrices $F_{b,1}^{R}$ and $F_{b,2}^{R}$. Here, we treat $F_{a}^{R}$ and $F_{b}^{R}$ as the true factors and $F_{a}^{R \perp}$ and $F_{b}^{R \perp}$ as the irrelevant factors. In the augmented system (3.4), we obtain $\Lambda^{R+}$ and $(\Lambda^{R+} + \Gamma^{R+})$, which are the factor loadings for pre- and post-break periods respectively. They will be further analyzed by the shrinkage estimator.

The numbers of factors $r_a$ and $r_b$ and instability in factor loadings can be estimated concurrently in model (3.4) by selecting the zero and nonzero components in $\Lambda^{R+}$ and $\Gamma^{R+}$ respectively. Therefore, it is essential to find such estimators, which turn out to be the shrinkage estimators, that can differentiate zero and nonzero components of matrices $\Lambda^{R+}$ and $\Gamma^{R+}$. In the rest of this section, the whole procedure will be explained in details.

Since the break date is known, we obtain the shrinkage estimator by minimizing the penalized least square (PLS) function which consists of two group-LASSO penalty functions (Yuan and Lin, 2006)[34].

First, we estimate $k$ potential factors. It is determined by the principle component estimators in each subsample. For instance, suppose $j \in \{a,b\}$, let $\tilde{F}_j$ be a $T_j \times k$ matrix which consists of the orthonormalized eigenvectors of $(NT_j)^{-1}X_jX_j'$ with the first $k$ largest eigenvalues. Since $\tilde{F}_a$ and $\tilde{F}_b$ are known, the shrinkage estimators $(\widehat{\Lambda}, \widehat{\Gamma})$ of the factor loadings $(\Lambda^{R+}, \Gamma^{R+})$ can be achieved by minimizing the following PLS
CHAPTER 3. PRELIMINARY RESULTS: KNOWN BREAK POINT CASE

criterion function:

\[
(\hat{\Lambda}, \hat{\Gamma}) = \arg\min_{\Lambda \in R^{N \times k}, \Gamma \in R^{N \times k}} [M(\Lambda, \Gamma) + P_1(\Lambda) + P_2(\Gamma)],
\]

where

\[
M(\Lambda, \Gamma) = (NT)^{-1} \left( \|X_a - \tilde{F}_a \Lambda\|^2 + \|X_b - \tilde{F}_b (\Lambda + \Gamma)\|^2 \right),
\]

\[
P_1(\Lambda) = \alpha_{NT} \sum_{\ell=1}^k \omega^\lambda_\ell \|\Lambda_\ell\| \quad \text{and} \quad P_2(\Gamma) = \beta_{NT} \sum_{\ell=1}^k \omega^\gamma_\ell \|\Gamma_\ell\|,
\]

where the \( \ell \) subscript of \( \Lambda_\ell \) and \( \Gamma_\ell \) denotes the \( \ell \)-th column of matrices \( \Lambda \) and \( \Gamma \), \( \alpha_{NT} \) and \( \beta_{NT} \) stand for positive constant parameters varying with \( N \) and \( T \), \( \omega^\lambda_\ell \) and \( \omega^\gamma_\ell \) are data-adapting weights defined as follow

\[
\omega^\lambda_\ell = \left( N^{-1} \|\tilde{\Lambda}_\ell\|^2 I\{\tilde{\Lambda}_\ell \neq 0_{N \times 1}\} + N^{-1} \|\tilde{\Lambda}_\ell,LS\|^2 I\{\tilde{\Lambda}_\ell = 0_{N \times 1}\} \right)^2,
\]

\[
\omega^\gamma_\ell = \left( N^{-1} \|\tilde{\Gamma}_\ell\|^2 I\{\tilde{\Gamma}_\ell \neq 0_{N \times 1}\} + N^{-1} \|\tilde{\Gamma}_\ell,LS\|^2 I\{\tilde{\Gamma}_\ell = 0_{N \times 1}\} \right)^2.
\]

Those two weights in (3.7) are constructed to distinguish between the zero and nonzero columns of \( \Lambda^{R+} \) and \( \Gamma^{R+} \). Here, \( \tilde{\Lambda} \in R^{N \times K} \) and \( \tilde{\Gamma} \in R^{N \times K} \) denote some preliminary estimators of factor loadings \( \Lambda^{R+} \) and \( \Gamma^{R+} \); \( \tilde{\Lambda}_{LS} = T_a^{-1} X_a' \tilde{F}_a \) and \( \tilde{\Psi}_{LS} = T_b^{-1} X_b' \tilde{F}_b \) are two unrestricted least square estimators of the pre- and post-break factor loading matrices. The change of factor loadings \( \tilde{\Gamma}_{LS} = \tilde{\Psi}_{LS} - \tilde{\Lambda}_{LS} \).

Note that, in this case, the simplest preliminary estimators are the unrestricted least squares estimators \( \tilde{\Lambda}_{LS} \) and \( \tilde{\Lambda}_{LS} \). We apply \( \tilde{\Lambda}_{LS} \) and \( \tilde{\Lambda}_{LS} \) if other preliminary estimators \( \tilde{\Lambda} \) and \( \tilde{\Gamma} \) have zero columns, which happens sometimes. Therefore we include
\( \mathcal{I}_A(x) \) in (3.7) which denotes the indicator function of the subset \( A \), i.e. it takes the value 1 if \( x \in A \), otherwise it takes the value 0.

The penalty functions \( P_1(\Lambda) \) and \( P_2(\Gamma) \), defined in terms of the column norms \( \|\Lambda_\ell\| \) and \( \|\Gamma_\ell\| \), are group-LASSO penalties (Yuan and Lin, 2006)[34]. We use this group-LASSO penalties to handle large-scale factor models and eliminate irrelevant factors which have zero factor loadings for all series (Cheng et al. 2016)[14]. It either sets all the elements in a group equal to zero or estimates them as non-zeros altogether.

### 3.3 Estimators of the Number of Factors and the Break Point

In this section, we show that by using the shrinkage estimators introduced in Section 3.2, the numbers of pre- and post-break factors \( r_a \) and \( r_b \) and the existence of the break point \( B_0 \) can be specified simultaneously. The structural instability can be identified as well. We use a binary variable \( B_0 \in \{0,1\} \) to indicate the occurrence of the break point. \( B_0=0 \) if there is no break, otherwise \( B_0=1 \). Thus, the structural instability in (2.3) can be rewritten as \( B_0=1 \) and \( r_a = r_b \).

The estimator of \( B_0 \) is defined as

\[
\hat{B} = \mathcal{I}_{\{\|\hat{\Gamma}\| > 0\}}.
\]  

(3.8)

To show that the estimator \( \hat{B} = 0 \), it is equivalent to show that \( \Gamma^0 = (\Gamma^0_1, \Gamma^0_2) = 0 \) in
(2.2). This holds if and only if $\Gamma^R = 0$, where $\Gamma^R = (\Gamma^R_1, \Gamma^R_2)$ is obtained by rewriting the normalized version of the post-break factor model in (3.3) as

$$X_b = F^R_b \Psi^R + e_b = F^R_{b,1} (\Lambda^R + \Gamma^R_1)' + F^R_{b,2} \Gamma^R_2' + e_b,$$

(3.9)

where $F^R_b = (F^R_{b,1}, F^R_{b,2})$, $\Psi^R = (\Lambda^R + \Gamma^R_1, \Gamma^R_2)$ and $\Gamma^R = (\Gamma^R_1, \Gamma^R_2)$.

The estimators of $r_a$ and $r_b$ are given by the numbers of the last non-zero columns of $\hat{\Lambda}$ and $\hat{\Gamma}$ respectively, which are

$$\hat{r}_a = \min \{ j \geq 1 : \|\hat{\Lambda}_\ell\| = 0 \text{ for all } \ell > j \},$$

$$\hat{r}_b = \max \left( \hat{r}_a, \min \{ j \geq 1 : \|\hat{\Gamma}_\ell\| = 0 \text{ for all } \ell > j \} \right).$$

(3.10)

### 3.4 Post Model Selection Estimation

In the previous sections, we obtain the estimators $\hat{r}_a$, $\hat{r}_b$ and $\hat{B}$ from the shrinkage estimator. In this section, we demonstrate that the shrinkage estimator also contributes to the estimation of the loading matrices $\Lambda$ and $\Gamma$. When we estimate the non-zero coefficients of the loading matrices, we have to re-estimate the loadings using least squares conditional on the selected $\hat{r}_a$, $\hat{r}_b$ and $\hat{B}$. In this case, we can optimize the penalty terms of the shrinkage estimator. The obtained estimator is called the post model selection (PMS) estimator.

First, consider there is no structural break ($\hat{B} = 0$), then we can re-estimate the fac-
tor model on the full sample. Suppose that $\tilde{F} \in R^{T \times k}$ is an orthonormalized matrix which consists of the first $k$ principal components of the full sample. Let $\tilde{\Lambda}$ denote the first $\hat{r}_a$ columns of the full sample least squares estimators $\tilde{\Lambda}_{LS} = T^{-1}X'\tilde{F} = \tilde{\Psi}_{LS}$ since there is no break. Thus, we set $\tilde{\Lambda} = \tilde{\Psi}$.

Second, if there exists a break ($\hat{B} = 1$), then we need to re-estimate the two subsamples of factors and loadings separately. Let $\tilde{F}_a$ and $\tilde{F}_b$ denote those two subsamples. Let $\tilde{\Lambda}$ denote the first $\hat{r}_a$ columns of $\tilde{\Lambda}_{LS} = T^{-1}X'_a\tilde{F}_a$. Similarly, let $\tilde{\Psi}$ denote the first $\hat{r}_b$ columns of $\tilde{\Psi}_{LS} = T^{-1}X'_b\tilde{F}_b$. The PMS estimators are defined as

$$
\hat{\Lambda}_{PMS} = (\tilde{\Lambda}, 0) \text{ and } \hat{\Psi}_{PMS} = (\tilde{\Psi}, 0), \tag{3.11}
$$

where $0$ is the zero matrix.

### 3.5 Estimation of the Tuning Parameters

In this section, we focus on the calculation of the shrinkage estimator. We propose a method to obtain the penalty functions by choosing the most appropriate tuning parameters $\alpha_{NT}$ and $\beta_{NT}$.

The group LASSO penalty functions $P_1(\Lambda)$ and $P_2(\Gamma)$ are determined by the tuning parameters $\alpha_{NT}$ and $\beta_{NT}$. They are the weights with respect to $X_a$ and $X_b$. The
tuning parameters are defined as

\[ \alpha_{NT} = \kappa_1 N^{-1/2} C_{NTa}^{-3} \quad \text{and} \quad \beta_{NT} = \kappa_2 N^{-1/2} C_{NTb}^{-3}, \] (3.12)

where \( C_{NTa} = \min \left( N^{1/2}, T_a^{1/2} \right) \), and \( C_{NTb} = \min \left( N^{1/2}, T_b^{1/2} \right) \). Let \( \kappa_1 \) and \( \kappa_2 \) be two constant numbers equal to

\[ \kappa_1 = c_1 \left\{ (NT_a)^{-1/2} \left\| e_a(\tilde{\Lambda}) \right\| + (NT_b)^{-1/2} \left\| e_b(\tilde{\Lambda} + \tilde{\Gamma}) \right\| \right\}, \]
\[ \kappa_2 = c_2 (NT_b)^{-1/2} \left\| e_b(\tilde{\Lambda} + \tilde{\Gamma}) \right\|, \] (3.13)

where \( \tilde{\Lambda} \) and \( \tilde{\Gamma} \) are preliminary estimators, \( e_a(\Lambda) \) and \( e_b(\Lambda + \Gamma) \) are residual matrices defined as

\[ e_a(\Lambda) = X_a - \tilde{F}_a \Lambda' \quad \text{and} \quad e_b(\Lambda + \Gamma) = X_b - \tilde{F}_b (\Lambda + \Gamma)' \] (3.14)

In (3.12), \( C_{NTa} \) and \( C_{NTb} \) are rates related to \( N \) and \( T \). We can balance these two rates by varying \( \alpha_{NT} \) and \( \beta_{NT} \) and replace the whole sample size \( T \) by the subsamples \( T_a \) and \( T_b \). We set up a default value for the constants \( c_1 \) and \( c_2 \) as \( c_1 = c_2 = 1 \), but we also apply a cross validation procedure to calibrate \( c_1 \) and \( c_2 \) over a fixed interval in finite samples.

### 3.6 Two-Step Estimation Algorithm

In this section, we introduce an effective two-step estimation algorithm from Cheng et al. (2016)[14]. First, we obtain a preliminary estimator, then we use it to adjust the penalty terms of the second-step shrinkage estimator.

This procedure develops the finite sample performance with respect to two aspects.
First, the residual matrices in (3.13) are optimized when \( \tilde{\Lambda} \) and \( \tilde{\Gamma} \) are based on a first-step estimator. Thus, the tuning parameters \( \alpha_{NT} \) and \( \beta_{NT} \) are improved in the second step. Second, the preliminary estimator \( \tilde{\Gamma} \) which is obtained through a rotation of the factor loadings \( \Lambda^R \) and \( \Psi^R \) also has better performance.

Let \( \tilde{\Lambda}^{(i)} \), \( \tilde{\Psi}^{(i)} \) and \( \tilde{\Gamma}^{(i)} \) denote the preliminary estimators with Step \( i \), \( i = 1 \) and 2. Let \( \hat{\Lambda}^{(i)}, \hat{\Psi}^{(i)} \) and \( \hat{\Gamma}^{(i)} \) denote the penalty least square (PLS) estimators in Step \( i \). Let \( \hat{\Lambda}_{PMS}^{(i)}, \hat{\Gamma}_{PMS}^{(i)} \) and \( \hat{\Psi}_{PMS}^{(i)} \) denote the post model selection (PMS) estimators in step \( i \).

The two-step estimation procedure is as follows:

**Algorithm 1 (Two-Step Estimation Procedure)**

1. First-Stage Shrinkage Estimation:
   1.1. Compute the unrestricted least square estimators \( \tilde{\Lambda}_{LS} \) and \( \tilde{\Gamma}_{LS} \).
   1.2. Let \( \tilde{\Lambda}^{(1)} = \tilde{\Lambda}_{LS} \) and \( \tilde{\Gamma}^{(1)} = \tilde{\Gamma}_{LS} \) be the first preliminary estimators. Calculate \( \omega_\lambda^i \), \( \omega_\gamma^i \), \( \alpha_{NT} \) and \( \beta_{NT} \) from (3.7), (3.12) and (3.13) with \( \tilde{\Lambda} = \tilde{\Lambda}^{(1)} \) and \( \tilde{\Gamma} = \tilde{\Gamma}^{(1)} \).
   1.3. Compute the shrinkage estimators \( \tilde{\Lambda}^{(1)} \) and \( \tilde{\Gamma}^{(1)} \) by minimizing the criterion function in (3.5).
   1.4. Estimate \( r_a \) and \( r_b \) from (3.10) with \( \hat{\Lambda} = \hat{\Lambda}^{(1)} \) and \( \hat{\Gamma} = \hat{\Gamma}^{(1)} \). Name the estimators as \( \hat{r}_a^{(1)} \) and \( \hat{r}_b^{(1)} \).
   1.5. Construct the PMS estimators \( \hat{\Lambda}_{PMS}^{(1)} \) and \( \hat{\Psi}_{PMS}^{(1)} \) by (3.11). If \( \hat{r}_a^{(1)} = \hat{r}_b^{(1)} \), then we define the rotation of the columns of \( \hat{\Psi}^{(1)} \) as

\[
\hat{\Psi}_{R}^{(1)} = \hat{\Psi}^{(1)}Q,
\]

(3.15)
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where $Q = VU'$ and let $UDV'$ denote the singular value decomposition of $\Lambda^{(1)'} \Psi^{(1)}$ such that $\Lambda^{(1)'} \Psi^{(1)} = UDV'$. The modified PMS estimator of $\Psi$ is defined as

$$
\hat{\Psi}_{PMS-R}^{(1)} = \left( \Psi_R^{(1)}, 0_{\Psi^{(1)}} \right) \in R^{N \times k}. \quad (3.16)
$$

2. Second-Stage Shrinkage Estimation

2.1. Let

$$
\tilde{\Lambda}^{(2)} = \hat{\Lambda}_{PMS}^{(1)},
\tilde{\Psi}^{(2)} = \begin{cases} 
\hat{\Psi}_{PMS-R}^{(1)} & \text{if } \hat{r}_b^{(1)} = \hat{r}_a^{(1)} \\
\hat{\Psi}_{PMS}^{(1)} & \text{if } \hat{r}_b^{(1)} > \hat{r}_a^{(1)} 
\end{cases},
\tilde{\Gamma}^{(2)} = \tilde{\Psi}^{(2)} - \tilde{\Lambda}^{(2)},
$$

and $\omega_{\ell}^\lambda, \omega_{\ell}^\gamma, \alpha_{NT}$, and $\beta_{NT}$ are calculated from (3.7), (3.12) and (3.13) with $\tilde{\Lambda} = \tilde{\Lambda}^{(2)}$ and $\tilde{\Gamma} = \tilde{\Gamma}^{(2)}$.

2.2. Compute the shrinkage estimators $\tilde{\Lambda}^{(2)}$ and $\tilde{\Gamma}^{(2)}$ by minimizing the criterion function (3.5).

2.3. Compute $B_0^{(2)}, \hat{r}_a^{(2)}$, and $\hat{r}_b^{(2)}$ from (3.10) with $\tilde{\Lambda} = \tilde{\Lambda}^{(2)}$ and $\tilde{\Gamma} = \tilde{\Gamma}^{(2)}$.

2.4. Construct the PMS estimator $\hat{\Lambda}_{PMS}^{(2)}$ and $\hat{\Psi}_{PMS}^{(2)}$ by the definition in (4.9) given the selection of the estimators $\tilde{B}^{(2)}$, $\tilde{r}_a^{(2)}$ and $\tilde{r}_b^{(2)}$.

The PMS estimator of the first stage is used to improve the preliminary estimators of the second stage. In Step 1.5, we obtain a transformation of $\Psi_R^{(1)}$ as $\Psi_R^{(1)}$ since the rotated matrix performs better in locating the structural break when there is no instability in the model. The transformation does not affect the procedure of the
asymptotic theory.

To give more explanation of Step 1.5, when we do the transformation, first we need to generate an orthogonal matrix $Q$ which can minimize $\|\Lambda^{(1)} - \Psi^{(1)}Q\|$. It suffices to maximize the correlation between the columns of $\Lambda^{(1)}$ and $\Psi^{(1)}Q$. The solution is $Q = VU'$ from the previous literature of Schönenmann (1966)[28]. In case of Step 1.5, $\Lambda^{(1)'}\Psi^{(1)} = UDV'$ by the singular value decomposition.

### 3.7 Cross Validation

We adopt the cross validation method (Cheng et al. 2016)[14] to calibrate the constant $c = (c_1, c_2) \in \mathcal{C}$ which appears in the penalty weights in (3.13). It has default value $c_1 = c_2 = 1$. We have already partitioned the sample according to $T$ time points. Since we are not able to observe the factors in current environment, we also need to subdivide the sample cross-sectionally.

First, the data in the cross-sectional dimension creates disjoint subsamples $X_{(-jN)}$ (N-regression) and $X_{jN}$ (N-prediction). Then, the procedure in Section 3.3 is applied to this subsample $X_{(-jN)}$. This produces the estimator of the unobserved factors and the estimators $\hat{r}_a(-jN, c)$, $\hat{r}_b(-jN, c)$ and $\hat{B}(-jN, c)$ for a given value $c$. We then partition the left subsample $X_{jN}$ along the $T$ dimension into regression and prediction samples. If the structural break is taken place in the partitioned subsample, then the regression and prediction samples need to be treated separately for the pre- and post-break periods. Given the estimators of the factors and the loadings, we can then
generate sample forecasts for the prediction sample.

The cross-validation criterion used in this major paper is built on mean-squared forecast errors (MSFE). The tuning constants are chosen to minimize the MSFE for given $c$. The minimization is accomplished over a bounded set $C$. The adjusted constant improves the performance of the finite sample.

### 3.8 Asymptotic Theory

In this section, we study the asymptotic performance of the PLS estimator. To estimate the number of factors from the observed data, the penalty for overfitting must be a function of both the cross-section dimension $N$ and the time dimension $T$ (Bai and Ng, 2002)[5]. In the research of Lewbel (1991)[21] and Donald (1997)[15], they assume either $N$ or $T$ is fixed. The reason is that the functions of $N$ and $T$, the AIC and BIC, usually do not work when both dimensions are large. However, in this major paper, we assume that both $N$ and $T$ converge to infinity. This adjustment is necessary because of our empirical knowledge. The time dimension of the dataset is not able to be assumed as a fixed $T$ since it is too large (Bai and Ng, 2002)[5]. Moreover, the cross-section dimension is larger relative to the time dimension. We present some assumptions and state some theorems, which are mainly from Cheng et al. (2016)[14], on the large sample properties of the preliminary estimators $\tilde{\Lambda}$ and $\tilde{\Gamma}$ and on the convergence rates of the sequences $\alpha_{NT}$ and $\beta_{NT}$ below.
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First, Theorem 3.1 and 3.2 give the stochastic order of the preliminary estimators \( \tilde{\Lambda} \) and \( \tilde{\Gamma} \) which may result in the change of the data-dependent weights \( \omega_\lambda^\ell \) and \( \omega_\gamma^\ell \) defined in (3.7). Their columns are separated into two parts, the first part is from 1 to the number of pre-break factors \( r_a \) and the second part is from the number of post-break factors \( r_b \) to \( k \). Let \( O_p \) (Big O) denote the stochastic boundedness. For example, \( X_n = O_p(a_n) \) means that \( X_n/a_n \) is stochastically bounded. Specifically, for any \( \epsilon > 0 \), there exists a finite \( M > 0 \) and a finite \( N > 0 \) such that \( \forall n > N \), \( P(|X_n/a_n| > M) < \epsilon \).

Theorem 3.1. If \( N, T \to \infty \) with \( \sqrt{N}/T \to \infty \), the preliminary estimators \( \tilde{\Lambda} \) and \( \tilde{\Gamma} \) satisfy

(i) \( \lim_{N,T \to \infty} \Pr(N^{-1}\|\tilde{\Lambda}_\ell\|^2 \geq C) = 1 \) for \( \ell = 1, \ldots, r_a \), \( N^{-1}\|\tilde{\Lambda}_\ell\|^2 = O_p(C_{NT}^{-2}) \) for \( \ell = r_a + 1, \ldots, k \);

(ii) If \( \Gamma^0 \neq 0 \), \( \lim_{N,T \to \infty} \Pr(N^{-1}\|\tilde{\Gamma}_\ell\|^2 \geq C) = 1 \) for \( \ell = 1, \ldots, r_b \), \( N^{-1}\|\tilde{\Gamma}_\ell\|^2 = O_p(C_{NT}^{-2}) \) for \( \ell = r_b + 1, \ldots, k \);

(iii) If \( \Gamma^0 = 0 \), \( N^{-1}\|\tilde{\Gamma}_\ell\|^2 = O_p(C_{NT}^{-2}) \) for \( \ell = 1, \ldots, k \).

For the proof of this Theorem, we refer to Cheng et al. (2016)[14]. Here, the large penalties generate shrinkage estimators that equal to 0 with probability approaching to 1. Also, \( \tilde{\Lambda}_\ell = 0 \) is a special case of \( N^{-1}\|\tilde{\Lambda}_\ell\|^2 = O_p(C_{NT}^{-2}) \), and so as \( \tilde{\Gamma}_\ell \).

Theorem 3.2. If \( N, T \to \infty \) with \( \sqrt{N}/T \to \infty \), the preliminary estimators \( \tilde{\Lambda}_{LS} \) and \( \tilde{\Gamma}_{LS} \) satisfy

(i) \( \lim_{N,T \to \infty} \Pr(N^{-1}\|\tilde{\Lambda}_{LS,\ell}\|^2 \geq C) = 1 \) for \( \ell = 1, \ldots, r_a \), \( N^{-1}\|\tilde{\Lambda}_{LS,\ell}\|^2 = O_p(C_{NT}^{-2}) \) for \( \ell = r_a + 1, \ldots, k \);

(ii) If \( \Gamma^0 \neq 0 \), \( \lim_{N,T \to \infty} \Pr(N^{-1}\|\tilde{\Gamma}_{LS,\ell}\|^2 \geq C) = 1 \) for \( \ell = 1, \ldots, r_b \),
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\[ N^{-1}\|\tilde{\Gamma}^{LS,\ell}\|^2 = O_p(C_{NT}^{-2}) \text{ for } \ell = r_b + 1, \cdots, k; \]

(iii) If \( \Gamma^0 = 0 \), \( N^{-1}\|\tilde{\Gamma}^{LS,\ell}\|^2 = O_p(C_{NT}^{-2}) \) for \( \ell = 1, \cdots, k \). \( \square \)

The proof of this Theorem follows from the results in Cheng et al. (2016)[14]. Theorem 3.2 takes care of the situation that \( \tilde{\Lambda} \) or \( \tilde{\Gamma} \) has zero columns. In this case, \( \omega^\Lambda_\ell \) and \( \omega^\gamma_\ell \) depend on \( \tilde{\Lambda}^{LS} \) and \( \tilde{\Lambda}^{LS} \).

Then, the following Assumptions 2 and 3 are designed for the large sample behaviour of the factors and they correspond to Assumptions A and B from Cheng et al. (2016)[14]. For the post-break period, let \( \overline{F}_t^0 = (\overline{F}_t^{0w}, \overline{F}_t^{0s})' \in \mathbb{R}^{r_b} \) denote the \( r_b \) factors after the break and \( C \in \mathbb{R} \) denotes a generic positive constant.

**Assumption 2. — Factors:**

\[ \mathbb{E}[\|F_t^0\|^4] \leq C, \mathbb{E}[\|F_t^0\|^4] \leq C \text{ and there exist positive definite nonrandom matrices } \Sigma_F \text{ and } \Sigma_T \text{ not depending on } N \text{ such that } T_0^{-1} \sum_{t=1}^{T_0} F_t^0 F_t^0' = \Sigma_F + O_p(T_0^{-1/2}) \text{ and } T_1^{-1} \sum_{t=T_0+1}^{T} \overline{F}_t^0 \overline{F}_t^{0w} = \Sigma_T + O_p(T_1^{-1/2}). \]

(2) The matrices \( \Sigma_F \Sigma_F \) and \( \Sigma_T \Sigma_T \) both have distinct eigenvalues. \( \square \)

**Assumption 3. — Factor Loadings:**

1. \( \|\lambda^0_i\| \leq C, \|\psi^0_i\| \leq C \text{ and there exist nonrandom matrices } \Sigma_\Lambda, \Sigma_\Psi \text{ and } \Sigma_{\Lambda \Psi} \text{ not depending on } N \text{ such that } \|\Lambda^0 \Lambda^0 / N - \Sigma_\Lambda\| \to 0, \|\Psi^0 \Psi^0 / N - \Sigma_\Psi\| \to 0 \text{ and } \|\Lambda^0 \Psi^0 / N - \Sigma_{\Lambda \Psi}\| \to 0 \text{ as } N \to \infty, \text{ where } \Sigma_\Lambda \text{ and } \Sigma_\Psi \text{ are positive definite.} \]

2. The matrices \( \Sigma_\Lambda \Sigma_\Psi \) and \( \Sigma_\Psi \Sigma_T \) both have distinct eigenvalues. \( \square \)
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Assumption 3 guarantees that each factor has a nontrivial contribution to the variance of $X_{it}$. We only consider nonrandom factor loadings.

Next, suppose that $T_0/T \to \tau_0$ for some constant $\tau_0 \in (0,1)$ as $T \to \infty$. Let $e = [e_1, \cdots, e_T] \in \mathbb{R}^{N \times T}$ be the matrix of errors and $e_{it}$ denote the $(i,t)$ element of series $i$ in $t$. Below, we present Assumption 4 which restricts the high dependence between the time dimensional partitions and the cross-sectional partitions in the error terms. Meanwhile, we also need Assumption 5 to restrict high dependence between the factors and the error terms.

Assumption 4. —Time and Cross-Section Dependence:

(i) $\mathbb{E}[e_{it}] = 0$, $\mathbb{E}[|e_{it}|] \leq C$;

(ii) $\mathbb{E}[N^{-1} \sum_{i=1}^{N} e_{is}e_{it}] = \sigma_N(s,t)$, $|\sigma_N(s,s)| \leq C$ for all $s$,

$T^{-1} \sum_{s=1}^{T} \sum_{t=1}^{T} |\sigma_N(s,t)| \leq C$;

(iii) $\mathbb{E}[e_{it}e_{jt}] = \tau_{ij,t}$ with $|\tau_{ij,t}| \leq |\tau_{ij}|$ for some $\tau_{ij}$ and for all $t$, and

$N^{-1} \sum_{i=1}^{N} \sum_{j=1}^{N} |\sigma_N(s,t)| \leq C$;

(iv) $\mathbb{E}[e_{it}e_{js}] = \tau_{ij,ts}$ and $(NT)^{-1} \sum_{i=1}^{N} \sum_{j=1}^{N} \sum_{t=1}^{T} \sum_{s=1}^{T} |\tau_{ij,ts}| \leq C$;

(v) For every $(t, s)$, $\mathbb{E}\left[N^{-1/2} \sum_{i=1}^{N} \left|e_{is}e_{it} - \mathbb{E}[e_{is}e_{it}]\right|\right] \leq C$;

(vi) $\rho_1((NT)^{-1}e_{at}e_{a}'_t) = O_p\left(\max[N^{-1}, T^{-1}]\right)$ and

$\rho_1((NT)^{-1}e_{bt}e_{b}'_t) = O_p\left(\max[N^{-1}, T^{-1}]\right)$.

Assumption 5. —Weak Dependence between Factors and Errors:

$\mathbb{E}\left[N^{-1} \sum_{i=1}^{N} \left\|T^{-1/2}\left(\sum_{t=1}^{T_0} F_t^0 e_{it} + \sum_{t=T_0+1}^{T} F_t^0 e_{it}\right)\right\|^2\right] \leq C$.

These two assumptions are analogous to the Assumptions C and D in Bai and Ng (2002)[5]. Note that Assumption 4(vi) is proposed as a condition to select the number
of factors (Amengual and Watson, 2007)\[3\].

Moreover, we need another assumption for the tuning parameter $\alpha_{NT}$ and $\beta_{NT}$, which determine the overall penalization. The following Assumption 6 states the convergence rate $C_{NT}$ of the unrestricted least square estimator at which the tuning parameters $\alpha_{NT}$ and $\beta_{NT}$ vanish asymptotically. As in Bai and Ng (2002)[5], we set $C_{NT} = \min(T^{1/2}, N^{1/2})$, where $C_{NT}$ is the convergence rate of the unrestricted least square estimator.

**Assumption 6.** The tuning parameters $\alpha_{NT}$ and $\beta_{NT}$ satisfy

(i) $\alpha_{NT} = O(N^{-1/2}C_{NT}^{-1})$ and $\beta_{NT} = O(N^{-1/2}C_{NT}^{-1})$;

(ii) $N^{-1/2}C_{NT}^{-5} = o(\alpha_{NT})$ and $N^{-1/2}C_{NT}^{-5} = o(\beta_{NT})$. $\square$

In Assumption 6, the boundaries of $\alpha_{NT}$ and $\beta_{NT}$ limit the extent of the overall penalization of shrinkage estimators. Assumption 6(i) ensures that the penalties on the nonzero columns are small when the weights $\omega_{\ell}^\lambda$ and $\omega_{\ell}^\gamma$ are stochastically bounded. Also, we propose to shrink the estimators of zero columns to zero. In Assumption 6(ii), we introduce the lower bound which requires the tuning parameter $\alpha_{NT}$ and $\beta_{NT}$ converge to zero slowly.

Finally, we present Theorem 3.3 which gives the asymptotic limits of the PLS estimators $\mathbf{\hat{\Lambda}}$ and $\mathbf{\hat{\Gamma}}$ as the following Theorem 3.3.

**Theorem 3.3.** Suppose that Assumptions 2-6 holds, then

(i) Pre-break loadings of relevant factors: $N^{-1}\|\mathbf{\hat{\Lambda}}_\ell - \Lambda^R_\ell\|^2 = O_p(C_{NT}^{-2})$ for $\ell = 1, \ldots, r_a$;

(ii) Pre-break loadings of irrelevant factors: $\lim_{N,T \to \infty} \Pr(\|\mathbf{\hat{\Lambda}}_\ell\|^2 = 0) = 1$
for $\ell = r_a + 1, \cdots, k$

(iii) Post-break changes in loadings of relevant factors: If $\Gamma^0 \neq 0$,

$$N^{-1}||\hat{\Gamma}_\ell - \Gamma^R_\ell||^2 = O_p(C_{NT}^{-2}) \text{ for } \ell = 1, \cdots, r_b;$$

(iv) No-break: If $\Gamma^0 = 0$, \[ \lim_{N,T \to \infty} \Pr(||\hat{\Gamma}_\ell||^2 = 0) = 1 \text{ for } \ell = 1, \cdots, r_b; \]

(v) Post-break changes in loadings of irrelevant factors:

$$\lim_{N,T \to \infty} \Pr(||\hat{\Gamma}_\ell||^2 = 0) = 1 \text{ for } \ell = r_b + 1, \cdots, k. \quad \square$$

The proof of Theorem 3.3 follows from the results in Cheng et al. (2016)[14]. Here, Theorem 3.3(i) and (ii) show that the factor loadings of the irrelevant factors are turned out to be 0 with probability approaching 1 because of the penalization. For $\ell = 1, \cdots, r_a$, the PLS estimators $\hat{\Lambda}_\ell$ and $\hat{\Gamma}_\ell$ converge in probability to the factor loadings $\Lambda^R_\ell$ and $\Gamma^R_\ell$ in the transformed factor model (3.3) respectively. Theorem 3.3(iii) to (v) are used to detect the structural instability by the asymptotic properties of the PLS estimators of the changes in the factor loadings. If there is no structural instability, like in part (iv), the PLS estimators of the changes $\hat{\Gamma}_\ell$ are turned out to have columns from 1 to $r_b$ equal to 0 with probability approaching to 1. If there does exists a structural instability, only the estimators $\hat{\Gamma}_\ell$ such that with column $\ell = r_b + 1, \cdots, k$ satisfies part (v).

To sum up, Theorem 3.3 states that the factor loadings of irrelevant factors can be estimated as 0 w.p.a 1. Meanwhile, the changes in the loadings of the relevant factors can be estimated as 0 with probability approaching to 1 if their factor loadings are not in the control of the structural instability. An equivalent way to get the PLS estimator is to show that the asymptotic limits of $N^{-1}||\Lambda^R_\ell||^2$ and $N^{-1}||\Gamma^R_\ell||^2$ in Part (i) and (iii) are bounded away from 0 which can be obtained from Assumption 1. The following Theorem 3.4 displays the asymptotic result of the estimators.
Theorem 3.4. Suppose that Assumptions 1-6 hold, then the following hold

\[\lim_{N,T \to \infty} \Pr(\hat{B} = B_0) = 1, \quad \lim_{N,T \to \infty} \Pr(\hat{r}_a = r_a) = 1, \quad \lim_{N,T \to \infty} \Pr(\hat{r}_b = r_b) = 1.\]

Theorem 3.4 shows that the model can be estimated simultaneously for any set of preliminary estimators \(\tilde{\Lambda}\) and \(\tilde{\Gamma}\) that satisfy Theorems 3.1 and 3.2. Theorem 3.4 is given in Cheng et al. (2016)[14]. Furthermore, for the convenience of the reader, we also present a detailed proof in the appendix.

Given the existence of the structural instability, we obtain a set of columns as following,

\[
Z = \left\{ \ell : N^{-1} \| \Gamma^R \ell \|^2 = N^{-1} \| \Psi^R W - \Lambda^R \ell \|^2 \geq C \right\}. \tag{3.18}
\]

This columns of set \(Z\) are essential for the estimation of the structural instability. Since the Step 1.5 of Algorithm 1 makes a transformation of the estimators, we require the following additional assumption:

**Assumption 7.** If \(r_a = r_b\), then \(\inf_{\| W \|=1} N^{-1} \| \Psi^R W - \Lambda^R \ell \|^2 \geq C\) for \(\ell \in Z\). \(\square\)

Assumption 7 is not restricted by the condition of \(\Lambda^R\). It is true whenever \(\Lambda^R\) is in the column space generated by \(\Psi^R\) or not. It is enforced on the factor loadings \(\Lambda^R\) of the normalized version of the factor model in (3.3) rather than the original factor model in (2.1) and (3.2). Additionally, if no structural instability exists, then \(Z\) is empty and Assumption 7 is not required.
Chapter 4

Proposed Methods in Unknown Break Point Case

In the previous chapter, we consider the case where the break point is known. In this chapter, we generalize the result to deal with the case where the break point is unknown. Similarly, we start with the identification of the instability, the calculation of the shrinkage estimator, the estimators of the number of factors and the break point and PMS estimator. We highlight the different analysis method from the known break point situation. We also point out the distinct implementation of estimating the tuning parameters and two-step estimation algorithm for unknown break point case.

Section 4.1 shows how to identify the instability. In Section 4.2, we present the process of calculating the shrinkage estimator. In Section 4.3, we describe the estimators of the number of factors and the break point. Section 4.4 presents the estimator of
\section{Identification of the Instability}

Given that \(T\) is the number of time points in the sample, we introduce a new parameter \(\pi = T_0/T \in \Pi\), which stands for the break point fraction, and \(\Pi\) is some closed subset inside \([0,1]\). Regarding to any \(\pi \in \Pi\), we split the full sample by the break point into pre- and post- subsets \(X_a(\pi) = (X_1, \cdots, X_{T_a})' \in \mathbb{R}^{T_a \times N}\) and \(X_b(\pi) = (X_{T_a+1}, \cdots, X_T)' \in \mathbb{R}^{T_b \times N}\), where \(T_a\) is the integer part of \(T\pi\) and \(T_b = T - T_a\). To obtain the unknown break point fraction \(\pi_0\), we need to study the number of factors in the observed \(X_a(\pi)\) and \(X_b(\pi)\), which are \(r_a(\pi)\) and \(r_b(\pi)\), with other value of \(\pi\). \(r_a(\pi)\) and \(r_b(\pi)\) are defined as the number of non-vanishing eigenvalues of \((NT)^{-1}X_a(\pi)'X_a(\pi)\) and \((NT)^{-1}X_b(\pi)'X_b(\pi)\) as \(N, T \to \infty\).

The existing study of Breitung and Eickmeier (2011)\cite{10} claims that the subsample of pre- and post-break observations will have one or more additional factors with a misspecified break point. More specifically, to correctly estimate the break point, we need to minimize the sum of the numbers of pre- and post-break factors by adjusting the potential break point fraction \(\pi\).
4.2 Shrinkage Estimator

When we estimate the shrinkage estimator, the different situation for unknown break point is that we need to consider an unknown and varying break point fraction \( \pi \in \Pi = [\pi_1, \pi_2] \), where \( \pi_1 > 0 \) and \( \pi_2 < 1 \).

Suppose that \( \tilde{F}_a(\pi) \in \mathbb{R}^{T_a \times k} \) consists of the orthonormalized eigenvectors of \( (NT_a)^{-1} X_a(\pi) X_a(\pi)' \) with its first \( k \) largest eigenvalues. Similarly, \( \tilde{F}_b(\pi) \in \mathbb{R}^{T_b \times k} \) consists of the orthonormalized left eigenvectors of \( (NT_b)^{-1} X_b(\pi) X_b(\pi)' \) with its first \( k \) largest eigenvalues. The unrestricted estimators of the factor loadings are \( \tilde{\Lambda}_{LS}(\pi) = T_a^{-1} X_a(\pi)' \tilde{F}_a(\pi) \), \( \tilde{\Psi}_{LS}(\pi) = T_b^{-1} X_b(\pi)' \tilde{F}_b(\pi) \), and the difference is \( \tilde{\Gamma}_{LS}(\pi) = \tilde{\Psi}_{LS}(\pi) - \tilde{\Lambda}_{LS}(\pi) \).

We can apply the procedure in Sections 3.2 with \( \pi_0 \) instead of \( \pi \). The shrinkage estimator with respect to different \( \pi \in \Pi \) estimates the corresponding \( r_a(\pi) \) and \( r_b(\pi) \). However, this procedure is not ideal because the estimators of \( r_a(\pi) \) and \( r_b(\pi) \) are highly dependent on \( \pi \). To overcome this difficulty, we introduce the averaging penalty functions to stabilize the shrinkage estimators. The shrinkage estimators can be obtained by the following function:

\[
(\hat{\Lambda}(\pi), \hat{\Gamma}(\pi)) = \arg\min_{\Lambda \in \mathbb{R}^{N \times k}, \Gamma \in \mathbb{R}^{N \times k}} \left[ M(\Lambda, \Gamma; \pi) + P_1^*(\Lambda) + P_2^*(\Gamma) \right],
\]

(4.1)

where

\[
M(\Lambda, \Gamma; \pi) = (NT)^{-1} \left[ \| X_a(\pi) - \tilde{F}_a(\pi) \Lambda' \|^2 + \| X_b(\pi) - \tilde{F}_b(\pi)(\Lambda + \Gamma)' \| \right],
\]

(4.2)
CHAPTER 4. PROPOSED METHODS IN UNKNOWN BREAK POINT CASE

\[ P_1^*(\Lambda) = \sum_{\ell=1}^{k} \mathbb{E}_\xi[\alpha_{NT}(\xi)\omega_\ell^{\Lambda}(\xi)] \|\Lambda_\ell\| \quad \text{and} \quad P_2^*(\Gamma) = \sum_{\ell=1}^{k} \mathbb{E}_\xi[\beta_{NT}(\xi)\omega_\ell^{\Gamma}(\xi)] \|\Gamma_\ell\|, \quad (4.3) \]

where \( \xi \) has uniform distribution on \( \Pi \) and \( \mathbb{E}_\xi[\cdot] \) stands for the expected value of the functions of \( \xi \). The terms \( P_1^*(\Lambda) \) and \( P_2^*(\Gamma) \) represent the averaging penalty functions.

The constant parameters \( \alpha_{NT}(\pi) \) and \( \beta_{NT}(\pi) \) depend on \( N \) and \( T \) for different value of \( \pi \). For \( \pi \in \Pi \), let \( \tilde{\Lambda}(\pi) \), \( \tilde{\Psi}(\pi) \) and \( \tilde{\Gamma}(\pi) \) be some preliminary estimators. We compute the adaptive weights \( \omega_\ell^{\Lambda}(\pi) \) and \( \omega_\ell^{\Gamma}(\pi) \) as:

\[
\omega_\ell^{\Lambda}(\pi) = \left( N^{-1} \|\tilde{\Lambda}_\ell(\pi)\|^2 I_{\{\tilde{\Lambda}_\ell(\pi) \neq 0_{N \times 1}\}} + N^{-1} \|\tilde{\Lambda}_\ell,LS(\pi)\|^2 I_{\{\tilde{\Lambda}_\ell(\pi) = 0_{N \times 1}\}} \right)^2,
\]

\[
\omega_\ell^{\Gamma}(\pi) = \left( N^{-1} \min \left\{ \|\tilde{\Gamma}_\ell(\pi)\|^2, \|\tilde{\Psi}_\ell(\pi)\|^2 \right\} I_{\{\tilde{\Gamma}_\ell(\pi) \neq 0_{N \times 1}\}} \right)^{-2},
\]

\[
+ \left( N^{-1} \min \left\{ \|\tilde{\Gamma}_\ell,LS(\pi)\|^2, \|\tilde{\Psi}_\ell,LS(\pi)\|^2 \right\} I_{\{\tilde{\Gamma}_\ell(\pi) = 0_{N \times 1}\}} \right)^2.
\]

Comparing two versions of the weights in (3.7) and (4.4), it should be noticed that \( \omega_\ell^{\Lambda}(\pi_0) = \omega_\ell^{\Lambda} \) but \( \omega_\ell^{\Gamma}(\pi_0) \neq \omega_\ell^{\Gamma} \). When the break point is unknown, we need to use the modified adaptive weights \( \omega_\ell^{\Gamma}(\pi) \) for estimation of \( r_b \) to generate larger penalties.

4.3 Estimators of the Number of factors and the Break Point

When we analyze the estimators, we consider the influence of varying break point fraction \( \pi \) and adopt *-superscripts to distinguish the one from the known break point case.

The estimators \( \hat{B}^*, \hat{r}_a^* \) and \( \hat{r}_b^* \) can be obtained as follows. The parameter \( B \) is estimated
CHAPTER 4. PROPOSED METHODS IN UNKNOWN BREAK POINT CASE

by,

\[ \hat{B}^* = \mathcal{I}\left\{ \sup_{\pi \in \Pi} \| \hat{\Gamma}(\pi) \| > 0 \right\}. \] (4.5)

The number of pre- and post-break factors can be estimated by

\[ \hat{r}_a^* = \min_{\pi \in \Pi} \hat{r}_a(\pi) \quad \text{and} \quad \hat{r}_b^* = \max_{\pi \in \Pi} \hat{r}_b(\pi), \] (4.6)

where \( \hat{r}_a(\pi) \) and \( \hat{r}_b(\pi) \) are defined as in (3.10), replacing \( \hat{\Lambda} \) and \( \hat{\Gamma} \) by \( \hat{\Lambda}(\pi) \) and \( \hat{\Gamma}(\pi) \) respectively.

4.4 Post Model Selection Estimation

In Section 3.4, we derive the PMS estimator in the known break point situation. If the break point is unknown, the PMS estimator is similar. The only difference is to add the notation \( * \)-superscripts and the \( (\pi) \)-arguments.

The PMS estimators are defined as

\[ \hat{\Lambda}_{PMS}(\pi) = (\hat{\Lambda}(\pi), 0) \quad \text{and} \quad \hat{\Psi}_{PMS}(\pi) = (\hat{\Psi}(\pi), 0), \] (4.7)

where 0 is the zero matrix.

Based on the method of Bai (1997)[4], one can estimate \( \pi_0 \) by using the least-squares objective function. Namely, we have

\[ \hat{\pi} = \arg\min_{\pi \in \Pi} Q_{NT}(\pi; \hat{r}_a^*, \hat{r}_b^*), \] (4.8)
where
\[
Q_{NT}(\pi; \widehat{r}_a^*, \widehat{r}_b^*) = (NT)^{-1} \left[ \left\| X_a(\pi) - \tilde{F}_a(\pi)\Lambda'_{PMS}(\pi) \right\| + \left\| X_b(\pi) - \tilde{F}_b(\pi)\Psi'_{PMS}(\pi) \right\| \right].
\] (4.9)

When we are able to detect a break, this procedure can be applied to identify the location of the break point.

### 4.5 Implementation of the Method and Algorithm

In this section, we generalize the method of estimating the tuning parameters first. Then, we apply the two-step estimation algorithm and calibrate the penalty weights by cross validation. The tuning parameters $\alpha_{NT}$ and $\beta_{NT}$ dealing with the unknown break point are similar to the known break point case described in Section 3.5. In this case, the tuning parameters are as follows:

\[
\alpha_{NT}(\pi) = \kappa_1(\pi) N^{-1/2} C^{-3}_{NTa} \quad \text{and} \quad \beta_{NT}(\pi) = \kappa_2(\pi) N^{-1/2} C^{-3}_{NTb},
\] (4.10)

where $\kappa_1(\pi) \in [\kappa_1, \bar{\kappa}_1]$ and $\kappa_2(\pi) \in [\kappa_2, \bar{\kappa}_2]$ for some $\kappa_1, \bar{\kappa}_2 < \infty$. The value of $\kappa_1(\pi)$ and $\kappa_2(\pi)$ are given as defined in (3.13) but with $\widetilde{\Lambda}$ and $\widetilde{\Gamma}$ replaced by $\widetilde{\Lambda}(\pi)$ and $\widetilde{\Gamma}(\pi)$.

A two-step procedure in Section 3.6 can be implemented to the model of unknown break point with $\pi_0$ replaced by the possible break point fraction $\pi$. Next, set the first step preliminary estimators $\tilde{\Lambda}^{(1)}(\pi)$, $\tilde{\Psi}^{(1)}(\pi)$ and $\tilde{\Gamma}^{(1)}(\pi)$ as unrestricted least squares estimators $\tilde{\Lambda}_{LS}(\pi)$, $\tilde{\Psi}_{LS}(\pi)$ and $\tilde{\Gamma}_{LS}(\pi)$; replace $\omega^\lambda$, $\omega^\gamma$, $\alpha_{NT}$ and $\beta_{NT}$ with $\omega^\lambda_{NT}(\pi)$,
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\[ \omega^*_\ell(\pi), \alpha_{NT}(\pi) \text{ and } \beta_{NT}(\pi); \]
replace the PLS criterion function (3.5) with (4.2); and use the numbers of factors \( \hat{r}_a \) and \( \hat{r}_b \) defined in (4.7). Here, the first step numbers of factors \( \hat{r}_a^{(1)} \) and \( \hat{r}_b^{(1)} \) still remain the same with varying value of \( \pi \) according to their definition in (4.6). Hence, we need to attain the first step shrinkage estimators \( \hat{\Lambda}^{(1)}(\pi), \hat{\Gamma}^{(1)}(\pi), \hat{r}_a^{(1)} \) and \( \hat{r}_b^{(1)} \). Next, we can access the second-step estimators \( \hat{\Lambda}^{(2)}(\pi) \) and \( \hat{\Gamma}^{(2)}(\pi) \). Finally, we can obtain the selected model with specific estimators \( \hat{B}^*, \hat{r}_a^*, \) and \( \hat{r}_b^* \) by those two-step PLS estimator \( \hat{\Lambda}^{(2)}(\pi) \) and \( \hat{\Gamma}^{(2)}(\pi) \) following the calculation algorithm in Section 3.5.

Meanwhile, the cross validation procedure is applied. We need to obtain a common value \( c \) which works for all possible break dates. For each potential \( \pi \), the subsample \( X_{(-jN)} \) are constructed in a similar way as in Section 4.4, only replacing \( \pi_0 \) by \( \pi \). For each \( \pi \) and the corresponding value of \( c \), we obtain a selected model. For the cross validation subsample \( X_{(jN)} \), the observations that are not inside the conjectured break interval \( \Pi \) are ignored.

We need to estimate the following four parameters before using the shrinkage estimator: the maximum number of potential factors \( k \), the break point fraction interval \( \Pi \), the domain \( \mathcal{C} \) for the tuning constants, and the numbers of sample partitions \( n_N \) and \( n_T \). First, \( k \) is inclined to be obtained from some preliminary examination of the data. The literature about specifying the value of \( k \) is associated with the estimation of the numbers of factors (Stock and Watson, 2012)[31]. Overestimating the value of \( k \) results in generating a large number of potential regressors, i.e. the factors, and may weaken the behavior of the shrinkage estimator. If \( \hat{r}_b = k \), then the value of \( k \)
may be set too small. Second, the break point fraction interval \( \Pi \), or in other words the break interval, is determined by the research object. The centre of the interval could be set around year 1984 if we are interested in studying the breaks related to the Great Moderation; or around 2007 if the aim is to analyze the Great Recession. The performance of the estimators can be optimized if we correctly determine a suitable length of the interval. Finally, we choose a set of \( C, n_N, \) and \( n_T \) that give good result for the Monte Carlo study.

### 4.6 Asymptotic Theory

In this section, we show that the shrinkage estimator with the averaging penalty in (4.3) can be extended to achieve the estimation of the unknown-break-point model. We use the two-step estimation algorithm detailed in Section 4.5. In Theorem 3.3 for known break point case, we first establish the asymptotic behaviour of the shrinkage estimators \( \hat{\Lambda}(\pi) \) and \( \hat{\Gamma}(\pi) \) for all \( \pi \). However, for an unknown break date, this step can be skipped. The principal reason is that the estimation of \( r_a(\pi) \) and \( r_b(\pi) \) can not be accomplished for all \( \pi \) by shrinkage estimators with the averaging penalty terms. The averaging penalties are inclined to be over-penalized if \( \pi \neq \pi_0 \). However, we can still estimate \( r_a \) and \( r_b \) because \( r_a \leq r_a(\pi) \) and \( r_b \leq r_b(\pi) \).

To indicate that the two-step PLS estimator in Section 4.5 generates the estimation of the model, we rewrite Assumption 7 to allow for the impact of the unknown break point and the averaging penalty. That gives us Assumption 8. For any \( \pi \in \Pi \), the
normalized system in (2.3) can be rewritten as

\[ X_a(\pi) = F^R_a(\pi) \Lambda^R(\pi)' + e_a(\pi) \quad \text{and} \quad X_b(\pi) = F^R_b(\pi) \Psi^R(\pi)' + e_b(\pi), \]

where \( F^R_a(\pi) \) and \( \Lambda^R(\pi) \) are \( T_a \times (r_a + r_b) \) and \( N \times (r_a + r_b) \) matrices, respectively, and \( F^R_a(\pi) \) and \( \Psi^R(\pi) \) are \( T_a \times (r_a + r_b) \) and \( N \times (r_a + r_b) \) matrices, respectively.

**Assumption 8.** (i) If \( r_a = r_b \), then

\[ \inf_{\pi \in \Pi, \|W\| = 1} N^{-1}\|\Psi^R(\pi)W - \Lambda^R_\ell(\pi)\|^2 \geq C \quad \text{for} \quad \ell \in \mathcal{Z}. \]

Assumption 8(i) is a generalization of Assumption 7 by replacing the break point \( \pi = \pi_0 \) with unknown \( \pi \in \Pi \).

We establish a theorem which shows that even for the unknown break point case, we can still get the asymptotic result of the estimators which include the existence of the break point and the numbers of factors. Before stating the theorem, we present Assumption 9 and 10.

**Assumption 9.** \( \mathbb{E}[\|F_0^\prime\|^4] \leq C, \quad \mathbb{E}[\|F_0^\prime\|_F^4] \leq C \) and there exist nonrandom positive definite matrices \( \Sigma_F \) and \( \Sigma_F^* \) such that

\[ T^{-1} \sum_{t=1}^{\lfloor T \pi \rfloor} F_0^\prime F_0 = \pi \Sigma_F + O_p(T^{-1/2}) \quad \text{for} \quad \pi \leq \pi_0 \quad \text{and} \quad T^{-1} \sum_{t=\lfloor T \pi \rfloor+1}^T F_0^\prime F_0 = (1-\pi) \Sigma_F + O_p(T^{-1/2}) \quad \text{for} \quad \pi \geq \pi_0, \]

where both \( O_p(T^{-1/2}) \) terms are uniform over \( \pi \in \Pi \).

**Assumption 10.** Assumption 4 holds with \( e_a \) and \( e_b \) replaced by \( e_a(\pi) \) and \( e_b(\pi) \) and Assumption 4(vi) holds uniformly over \( \pi \in \Pi \).

**Theorem 4.1.** Suppose that Assumptions 1, 3, 5, 8, 9 and 10 hold. Then the estimators within the model selected by the two-step estimator in Algorithm 1 satisfy
CHAPTER 4. PROPOSED METHODS IN UNKNOWN BREAK POINT CASE

\[ \lim_{N,T \to \infty} \Pr(\hat{B}^* = B_0) = 1, \quad \lim_{N,T \to \infty} \Pr(\hat{r}_a^* = r_a) = 1, \quad \lim_{N,T \to \infty} \Pr(\hat{r}_b^* = r_b) = 1. \]

The proof of this theorem is similar to that given for Theorem 3.4. We also refer to Cheng et al. (2016)[14].

If the difference between \( \pi \) and \( \pi_0 \) is considerable, the averaging penalty terms working in the shrinkage estimators tend to over-penalize the loadings that would be set to zero for \( \pi = \pi_0 \). It leads to a tendency of underestimating either \( r_a(\pi) \) or \( r_b(\pi) \) if the conjectured break point is improperly specified. The estimation of the break point can be identified by applying the estimates \( \hat{r}_a^* \) and \( \hat{r}_b^* \) to the least squares objective function in (4.8). In the next chapter, we implement the Monte Carlo simulation to evaluate the finite-sample performance of the shrinkage estimators and analyze the empirical dataset.
Chapter 5

Numerical Results

In the previous chapters, we develop a shrinkage estimator and design estimators of the occurrence of break point and the number of factors. In the first section of Chapter 5, we carry out Monte Carlo Simulations to evaluate the performance of the estimators. Thus, we compute the probability of correctly estimating the numbers of pre- and post- break factors and the existence of the break point. Moreover, we also compute the mean-squared errors for out-of-sample forecasts (MSFE). In the second section, we deal with the real world financial problem, the Great Recession.

5.1 Monte Carlo Simulations

Monte Carlo Simulation is a widely used technique to evaluate the performance of the proposed statistical procedure. The behaviour of the procedure in random samples can be estimated by the empirical process of actually drawing numerous random samples as the experiments (Mahadevan, 1997)[23]. The strategy of doing this is to create
an artificial "world," or pseudo-population, which simulates the real world, and to conduct numerous trials to investigate how that procedure behaves across samples (Mooney, 1997)[24].

Section 5.1 is divided into two subsections. Section 5.1.1 introduces the design of the factor models and simulations used in the experiments. The simulation outcomes and interpretation are displayed in Section 5.1.2.

5.1.1 Design of the Factor Models

In this section, we describe the design procedure of the factor model and simulations. The factor models of this major paper refer to the paper of Bates et al. (2013)[6], with the modification of stabilizing the structural instability and emphasizing the influence of large breaks instead of small breaks. The factor models have the forms as follows

\[
\begin{align*}
\text{Pre-break:} & \quad X_{it} = \lambda_i' F_t + e_{it}, \quad F_{t,\ell} = \rho_a F_{t-1,\ell} + u_{t,\ell}, \\
& \quad t = 1, \cdots, [T\pi_0], \quad \ell = 1, \cdots, r_a, \\
\text{Post-break:} & \quad X_{it} = \psi_i' F_t + e_{it}, \quad F_{t,\ell} = \rho_b F_{t-1,\ell} + u_{t,\ell}, \\
& \quad t = [T\pi_0] + 1, \cdots, T, \quad \ell = 1, \cdots, r_b,
\end{align*}
\]

(5.1)

where \(i = 1, \cdots, N, \ F_t = (F_{t,1}, \cdots, F_{t,r_a})', \ F_t = (F_{t,1}, \cdots, F_{t,r_b})'. \) For \( t = 1, \cdots, T \) and \( \ell = 1, \cdots, r_b \), \( u_{t,\ell} \) are i.i.d as \( N(0, 1) \). To take into account for the time dimen-
sional and cross-sectional dependence of the errors, we suppose that

\[ e_{it} = \alpha e_{i,t-1} + v_{it}, \quad v_{it} = (v_{1t}, \ldots, v_{Nt})' \sim N(0, \Omega), \tag{5.2} \]

where the \((i, j)\)-th element of \(\Omega\) is \(\beta|i-j|\). We suppose that the processes \(\{u_{t,\ell} : \ell = 1, \ldots, r_b\}\) and \(\{v_{it} : i = 1, \ldots, N\}\) in (5.2) are mutually independent and are i.i.d across \(t\). \(F_0\) and \(e_0 = (e_{10}, \ldots, e_{N0})'\) denote the initial values of the factors and the errors. We suppose that they are drawn from the stationary distribution. If there is no break, which means \(r_b = r_a\), then \(F_{T_0} = F_{T_b}\). If there are some new factors after the break point, which means \(r_b > r_a\), then \(F_{T_0} = (F_{T_0}', F_{T_b}')'\), where each element of \(F_{T_0}'\) is drawn independently from the distribution of \(F_{t,\ell}\).

First, we explain how to construct the pre-break factor loadings \(\{\lambda_i : i = 1, \ldots, N\}\). Let \(\lambda_i \sim N(0, \Sigma_i)\) independently, where \(\Sigma_i\) is a diagonal matrix with distinctive elements \(\sigma_i^2(1), \ldots, \sigma_i^2(r_a)\). The sum of these diagonal elements controls the population regression \(R_{i}^2\) of \(X_{it}\) on the factors. Here, \(R_{i}^2\) is the pre-specified regression \(R^2\) of the \(i\)-th series. The method of selecting \(R_{i}^2\) for \(i = 1, \ldots, N\) is drawn by Bai and Ng (2002)[5] that \(R_{i}^2\) is set to 0.5.

Second, we demonstrate the construction of post-break factor loadings \(\psi_i\), which requires to consider the existence of the structural instability. We set \(\psi_i = (1 - w)\lambda_i + w\lambda_i^*\), where \(\lambda_i^*\) has the same distribution as \(\lambda_i\) but they are independent. We can control the size of the instability by changing the value of the scalar \(w\). Particularly, \(w = 0\) corresponds to no break in the factor loadings.
CHAPTER 5. NUMERICAL RESULTS

The simulated time series has mean 0 and variance 1. We extract a maximum of $k = 8$ potential factors from the sample. For known break point, the estimators is based on the two-step PLS estimator described in Algorithm 1. We set $n_N = 5$ and $n_T = 10$, which means we have five cross-sectional partitions and ten time dimensional partitions. Normally, considering the cross-sectional dimension, the procedure has to be implemented on each partition of the sample, so it is time consuming. On the contrary, the analysis based on time partitions is fast.

For the unknown break point case, the estimation depends on adjusted version of Algorithms 1 described in Section 4.4. We consider a discrete set $\Pi_d$ which has the grid size $\tau = 0.001$. It is a shift by a quarter for a monthly data set of 300 periods. We set $\Pi_d = \{\pi_0 - 4\tau, \pi_0 - 3\tau, \ldots, \pi_0, \pi_0, \ldots, \pi_0 + 4\tau\}$, which is symmetric around the true break point $\pi_0$ and it stretches over a two-year interval in total. Moreover, the post-break subsample of the PMS estimator is attained by the least square estimator of the break point described in Section 4.3.

To evaluate the performance of the selected model, we also compute the mean-squared errors for out-of-sample forecasts (MSFE). The forecast series follows

\begin{align}
\text{Initial Value: } & y_1 = X_{1T} \\
\text{Pre-break: } & y_{t+1} = \varphi_a' F_t + \epsilon_{t+1}, \quad t = 1, \ldots, T_a, \\
\text{Post-break: } & y_{T_a+1} = \varphi_b' \bar{F}_t + \epsilon_{T_a+1}, \quad t = T_a + 1, \ldots, T_a + T_b. 
\end{align}

(5.3)

Suppose that the errors $\epsilon_1, \epsilon_2, \ldots, \epsilon_{T_a+T_b}$ are iid as $N(0,1)$. The loading vector is gen-
From the distribution \( \varphi_a \sim N(0, I_r) \). If there is no structural break, then we have unchanging factor loadings, which means \( \varphi_b = \varphi_a \). If we consider the structural instability defined in (2.3), then the post-break factor loading \( \varphi_b = (1 - w)\varphi_a + w\varphi^*_a \), where \( \varphi^*_a \) and \( \varphi_a \) are independent and have the same distribution.

To evaluate the MSFE of the out-of-sample forecasts, first, we select the model and the factors based on the \( X \) sample. Second, if there is no break points, we estimate \( \varphi_b = \varphi_a \) based on the full sample \( t = 1, \cdots, T_a + T_b - 1 \) and evaluate the MSFE associated with the prediction of \( y_{Ta+Tb+1} \). If there exists a break point, we estimate \( \varphi_b \) based on the subsample \( t = Ta + 1, \cdots, Ta + T_b - 1 \) and evaluate the MSFE associated with the prediction of \( y_{Ta+Tb+1} \).

\[
\text{MSFE}(\hat{y}_{Ta+Tb+1}) = \mathbb{E}\left[ (y_{Ta+Tb+1} - \hat{y}_{Ta+Tb+1})^2 \right].
\] (5.4)

In the simulation, we report the relative MSFE in (5.5), which is based on both the PMS estimator and the full-sample estimation. The full-sample estimator \( \hat{\Lambda} \) is defined as the first \( r \) columns of the full sample least squares estimator \( \tilde{\Lambda}_{LS} = T^{-1}X'\tilde{F} \), where \( r = r_a \) if \( B_0 = 0 \) (no break), \( r = r_a + r_b \) otherwise. If the value of the relative MSFE is less than 1, then the predictor based on the PMS estimator domains the full sample estimator.

\[
\text{Relative MSFE} = \frac{\text{MSFE}_{PMS}}{\text{MSFE}_{full}}
\] (5.5)
CHAPTER 5. NUMERICAL RESULTS

5.1.2 Results for Shrinkage Estimator

In this section, we present the results of three types of Monte Carlo experiments listed in Table 5.1. In the first experiment, we assume the break point is known as located in the middle of the sample ($\pi = 0.5$). The cross-sectional correlation is set to be modest that $\alpha$ and $\beta$ equal to 0.2. The second experiment has the same rate of correlation. In particular, we first consider that the break point is known and located nearly at the end of the sample ($\pi_0 = 0.8$). Second, we consider the scenario where the break point is unknown. Finally, besides the choice of a stronger cross-sectional correlation ($\alpha, \beta = 0.5$), the third experiment is similar to the first one. We suppose that the temporal correlation $\rho_a$ and $\rho_b$ are equal to 0.5 in all three cases, and we implement 1,000 Monte Carlo experiments to take the average result.

<table>
<thead>
<tr>
<th>Exp.</th>
<th>$\pi_0$</th>
<th>$\alpha, \beta$</th>
<th>Break Point</th>
</tr>
</thead>
<tbody>
<tr>
<td>1</td>
<td>0.5</td>
<td>0.2</td>
<td>Known</td>
</tr>
<tr>
<td>2</td>
<td>0.8</td>
<td>0.2</td>
<td>Known, Unknown</td>
</tr>
<tr>
<td>3</td>
<td>0.5</td>
<td>0.5</td>
<td>Known</td>
</tr>
</tbody>
</table>

Notes: $\pi_0$: break point fraction; $\alpha, \beta$: cross-sectional correlation; temporal correlation $\rho_a = \rho_b = 0.5$.

First, we display the Monte Carlo results of Experiment 1 in Table 5.2. The Panel A and B correspond to the situations of no break and the existence of instability respectively. We set the scalar of the instability $w$ as 0.5 and 0.7 and we include various choices of $N$ and $T$ in the experiment to test the asymptotic behaviour. We report the probability of correctly determining $B_0$, $r_a$ and $r_b$. The last column gives the MSFE of the predictor based on the PMS estimator relative to the predictor based on a forecast of full-sample least-squares estimator, where the number of factors is
set to $r_a$ for Panel A, and to $r_a + r_b$ for Panel B. The values less than 1 mean that the proposed PMS predictor is more accurate.

<table>
<thead>
<tr>
<th>$r_a$</th>
<th>$r_b$</th>
<th>$w$</th>
<th>$N$</th>
<th>$T$</th>
<th>$\Pr(\hat{B} = B_0)$</th>
<th>$\Pr(\hat{r}_a = r_a)$</th>
<th>$\Pr(\hat{r}_b = r_b)$</th>
<th>Relative MSFE</th>
</tr>
</thead>
<tbody>
<tr>
<td>3</td>
<td>3</td>
<td>0</td>
<td>120</td>
<td>120</td>
<td>1.00</td>
<td>1.00</td>
<td>1.00</td>
<td>1.00</td>
</tr>
<tr>
<td>3</td>
<td>3</td>
<td>0</td>
<td>140</td>
<td>140</td>
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<td>1.00</td>
<td>1.00</td>
</tr>
<tr>
<td>3</td>
<td>3</td>
<td>0</td>
<td>180</td>
<td>180</td>
<td>1.00</td>
<td>1.00</td>
<td>1.00</td>
<td>1.00</td>
</tr>
<tr>
<td>3</td>
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<td>0.5</td>
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</tr>
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<td>1.00</td>
<td>1.00</td>
</tr>
<tr>
<td>3</td>
<td>3</td>
<td>0.7</td>
<td>120</td>
<td>120</td>
<td>1.00</td>
<td>1.00</td>
<td>1.00</td>
<td>0.39</td>
</tr>
<tr>
<td>3</td>
<td>3</td>
<td>0.7</td>
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<td>1.00</td>
<td>0.80</td>
</tr>
<tr>
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<td>1.06</td>
</tr>
</tbody>
</table>

**Notes:** Cross-sectional correlation $\alpha = \beta = 0.2$; temporal correlation $\rho_a = \rho_b = 0.5$.

The procedure is very accurate in estimating the estimators $\hat{B}$, $\hat{r}_a$ and $\hat{r}_b$ in all cases with chosen value of $w$. The possible reason is that the selection procedure has strong power which can deal with big changes in the loadings. When the break in the factor loadings is rather big ($w = 0.5$ or $0.7$), it identifies the break point and the number of factors with probability 1. In terms of the relative MSFEs, it should be noticed that the proposed PMS predictor are almost accurate since the relative MSFEs is approximately equal to or less than 1. It means that the PMS predictor dominates the full-sample predictor.

Second, we move to the known break point case of Experiment 2. Table 5.3 gives the probability of $\hat{B} = B_0$, $\hat{r}_a = r_a$ and $\hat{r}_b = r_b$ with instability scalar 0 (no break), 0.7 and 1, as well as the MSFE.
Table 5.3: Known Break Point, $\pi_0 = 0.8$

<table>
<thead>
<tr>
<th>$r_a$</th>
<th>$r_b$</th>
<th>$w$</th>
<th>$N$</th>
<th>$T$</th>
<th>$\Pr(\hat{B} = B_0)$</th>
<th>$\Pr(\hat{r}_a = r_a)$</th>
<th>$\Pr(\hat{r}_b = r_b)$</th>
<th>MSFE</th>
</tr>
</thead>
<tbody>
<tr>
<td>3</td>
<td>3</td>
<td>0</td>
<td>120</td>
<td>240</td>
<td>1.00</td>
<td>1.00</td>
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<td>1.00</td>
<td>1.00</td>
<td>1.00</td>
</tr>
</tbody>
</table>

Panel A. No Break

<table>
<thead>
<tr>
<th>$r_a$</th>
<th>$r_b$</th>
<th>$w$</th>
<th>$N$</th>
<th>$T$</th>
<th>$\Pr(\hat{B} = B_0)$</th>
<th>$\Pr(\hat{r}_a = r_a)$</th>
<th>$\Pr(\hat{r}_b = r_b)$</th>
<th>MSFE</th>
</tr>
</thead>
<tbody>
<tr>
<td>3</td>
<td>3</td>
<td>0.7</td>
<td>120</td>
<td>240</td>
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<td>1.00</td>
<td>1.00</td>
<td>1.05</td>
</tr>
<tr>
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<td>1.00</td>
<td>1.00</td>
<td>0.69</td>
</tr>
<tr>
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<td>1.00</td>
<td>1.00</td>
<td>1.39</td>
</tr>
<tr>
<td>3</td>
<td>3</td>
<td>1</td>
<td>120</td>
<td>240</td>
<td>1.00</td>
<td>1.00</td>
<td>1.00</td>
<td>0.55</td>
</tr>
<tr>
<td>3</td>
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<td>300</td>
<td>1.00</td>
<td>1.00</td>
<td>1.00</td>
<td>0.36</td>
</tr>
<tr>
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<td>3</td>
<td>1</td>
<td>150</td>
<td>300</td>
<td>1.00</td>
<td>1.00</td>
<td>1.00</td>
<td>0.61</td>
</tr>
</tbody>
</table>

Panel B. Instability

Notes: Cross-sectional correlation $\alpha = \beta = 0.2$; temporal correlation $\rho_a = \rho_b = 0.5$.

When there is no break point, experiment 2 turns out that the procedure is overall accurate to determine $B_0$, $r_a$ and $r_b$. With a variety value of $N$ and $T$, the estimators of $B$, $r_a$ and $r_b$ are almost accurate with probability 1. Moreover, the PMS estimator weakly dominates the full-sample predictor since the value of MSFE are slightly less than 1. Compared with the instability scale $w = 0.7$, when it is set to be 1, the proposed PMS predictor does better than the full-sample least squares estimator. We can tell that a larger value of $w$ is necessary for the estimation to be accurate.

Table 5.4 is the output of Experiment 2 which shows that an unknown break point leads the result to be less accurate. Under the no-change scenario, the procedure correctly determines the estimators for all three sample sizes. However, to take into account the structural instability, it seems that a larger break in the loadings ($w = 1$ instead of $w = 0.7$) and larger sample size ($N > 120$ and $T > 240$) is preferred to increase the probability of detecting the break. Besides that, the result of estimating
the numbers of factors for all cases is perfect. According to the last column of Table 5.4, the PMS predictor performs well under the no-break scenario since the MSFEs equal to 1, which means that the PMS predictor is equivalent to the full-sample predictor. When we consider the instability, for both cases $w = 0.7$ and $w = 1$, the proposed PMS estimator dominates the full-sample predictor.

Table 5.4: Unknown Break Point, $\pi_0 = 0.8$

<table>
<thead>
<tr>
<th>$r_a$</th>
<th>$r_b$</th>
<th>$w$</th>
<th>$N$</th>
<th>$T$</th>
<th>$\Pr(\hat{B} = B_0)$</th>
<th>$\Pr(\hat{r}_a = r_a)$</th>
<th>$\Pr(\hat{r}_b = r_b)$</th>
<th>MSFE</th>
</tr>
</thead>
<tbody>
<tr>
<td></td>
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<td></td>
<td>Panel A. No Break</td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>3</td>
<td>3</td>
<td>0</td>
<td>120</td>
<td>240</td>
<td>1.00</td>
<td>1.00</td>
<td>1.00</td>
<td>1.00</td>
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<tr>
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<td></td>
<td></td>
<td></td>
<td></td>
<td>Panel B. Instability</td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>3</td>
<td>3</td>
<td>0.7</td>
<td>120</td>
<td>240</td>
<td>0.00</td>
<td>1.00</td>
<td>1.00</td>
<td>0.79</td>
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<td>240</td>
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<td>1.00</td>
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<td>1.17</td>
</tr>
<tr>
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<td>0.18</td>
</tr>
</tbody>
</table>

*Notes*: Cross-sectional correlation $\alpha = \beta = 0.2$; temporal correlation $\rho_a = \rho_b = 0.5$.

Finally, we present the result of experiment 3, which is very similar to experiment 1 (Table 5.1) but only with stronger dependence and correlation in errors ($\alpha = \beta = 0.5$). The results are reported in Table 5.5. In general, the procedure correctly estimates $B_0$, $r_a$ and $r_b$. All the value of MSFEs are less than 1, so the results favor the proposed PMS predictor. The noticeable thing is that compared to Table 5.1, we obtain less accurate detection of the break point and the estimation of the number of post-break factors in the case of $N, T = 140$ due to the increased cross-sectional correlation.
Table 5.5: Known Break Point, $\pi_0 = 0.5$

<table>
<thead>
<tr>
<th>$r_a$</th>
<th>$r_b$</th>
<th>$w$</th>
<th>$N$</th>
<th>$T$</th>
<th>Pr($\hat{B} = B_0$)</th>
<th>Pr($\hat{r}_a = r_a$)</th>
<th>Pr($\hat{r}_b = r_b$)</th>
<th>MSFE</th>
</tr>
</thead>
<tbody>
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<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
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<td>3</td>
<td>0</td>
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<td>1.00</td>
<td>1.00</td>
<td>1.00</td>
</tr>
<tr>
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<td>1.00</td>
<td>1.00</td>
</tr>
<tr>
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<td>1.00</td>
<td>1.00</td>
<td>1.00</td>
</tr>
<tr>
<td>Panel B. Instability</td>
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<td></td>
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<td></td>
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<td>0.10</td>
</tr>
<tr>
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<td>0.50</td>
<td>0.79</td>
</tr>
<tr>
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<td>1.00</td>
<td>1.00</td>
<td>0.28</td>
</tr>
</tbody>
</table>

Notes: Cross-sectional correlation $\alpha = \beta = 0.5$; temporal correlation $\rho_a = \rho_b = 0.5$.

5.2 Real Data Set: the Great Recession

The Great Recession in the United States started in December 2007 and finished in June 2009. It originates in an unusually dramatic financial crisis, which began with a financial collapse that erased more than half the capitalization of the stock market (Grusky et al. 2011)[19]. It is known as the longest post-war recession and its negative impact to the labor market is huge. From May 2007 to October 2009, the labor force lost 7.5 million jobs and more, and the unemployment rate ascended from 4.4% to 10.1% (Roberts 2009)[27]. From the global view, the world economy declined by 6% (Roberts 2009)[27]. We pick the Great Recession as an example for the empirical analysis because of its profound and large-scale influence.

Based on the data set of the Great Recession, we use the procedures developed in Chapters 2 to 4 to investigate the fluctuation of factor loadings and the emergence of
new factors. Section 5.2.1 gives some preliminary transformations and the analysis results are given in Section 5.2.2.

5.2.1 Some Preliminary Transformations

The data set of the Great Recession is from Stock and Watson (2012)[31], which contains a set of 200 macroeconomic and financial indicators. As in Cheng et al. (2016)[14]. After eliminating the duplicate series, we obtain the data set SW132 which includes 132 out of 200 indicators. We extend SW132 to 2012:M12, which represents the twelfth month of year 2012, using May 2013 data vintages. For consumption of durables, nondurables, services, nonresidential investment and 16 price series, we replace the quarterly series by their monthly counterparts. The rest quarterly series, which are not able to be converted to monthly series, are removed. Then, we include two new monthly components: change in private inventory and wage and salary disbursements. Following Stock and Watson (2012)[31], we remove local means from all series using a biweight kernel with a bandwidth of 100 months. The local means are approximately the same as the ones obtained by a centered moving average of ±70 months. The data set consists the observations of $N = 102$ series of monthly financial indicators, e.g. growth rates of participation of recreation services or monthly inflation of sales of retail stores. The sample begins after the Great Moderation and ranges from January 1985 to January 2013 which lasts 337 months ($T = 337$). A detailed list of the financial indicators is presented in the appendix.
5.2.2 Analysis Results

The two step estimation procedure described in Section 3.5 is applied. Moreover, when considering unknown break point, we also combine the procedure summarized in Section 4.4. We fix the number of potential factors \( k = 8 \) and use \( n_N = 5 \) and \( n_T = 10 \). The model selection results are presented in Table 5.6.

<table>
<thead>
<tr>
<th>Size</th>
<th>Interval (Month) Range</th>
<th>Factors</th>
<th>Break Points</th>
<th>Least Sq. (( \hat{\pi} ))</th>
<th>Revised</th>
</tr>
</thead>
<tbody>
<tr>
<td>0</td>
<td>2007:M12-2007:M12</td>
<td>1 ( \hat{r}_a )</td>
<td>1 ( \hat{r}_b - \hat{r}_a )</td>
<td>2007:M12</td>
<td>2007:M12</td>
</tr>
<tr>
<td>3</td>
<td>2007:M9-2008:M3</td>
<td>1 ( \hat{r}_a )</td>
<td>1 ( \hat{r}_b - \hat{r}_a )</td>
<td>2007:M9</td>
<td>2007:M12</td>
</tr>
<tr>
<td>6</td>
<td>2007:M6-2008:M6</td>
<td>1 ( \hat{r}_a )</td>
<td>1 ( \hat{r}_b - \hat{r}_a )</td>
<td>2007:M6</td>
<td>2007:M12</td>
</tr>
<tr>
<td>9</td>
<td>2007:M3-2008:M9</td>
<td>1 ( \hat{r}_a )</td>
<td>2 ( \hat{r}_b - \hat{r}_a )</td>
<td>2007:M3</td>
<td>2007:M12</td>
</tr>
</tbody>
</table>

Notes: We center the interval at 2007:M12 and use the averaging penalty functions \( P_1^*(\Lambda) \) and \( P_2^*(\Lambda) \) defined in (4.3) where the average is taken over the interval 2007:M12 \( \pm \) Size.

According to the U.S. National Bureau of Economic Research (the official arbiter of U.S. recessions), December 2007 is the beginning of the Great Recession. Therefore, the picked interval of potential break points are centered around \( T_c = 2007:M12 \) with four different interval size. For the first case in Table 5.6, the interval contains a single month (2007:M12), meaning that we consider the break point happens in 2007:M12. For interval size 9, the potential break point located in the months from 2007:M3 to 2008:M9. From the result, we attain one pre-break factor (\( \hat{r}_a = 1 \)) for all variety of interval sizes, and two or three post-break factors (\( \hat{r}_b = 2 \) or \( \hat{r}_b = 3 \)). There is a strong evidence of a change in the number of factors. In columns 4 and 5 of Table 5.6, we report the least squares estimation of the break point \( \hat{\pi} \) defined in (4.8). In the fifth column, we minimize the least square criterion over the chosen interval.
CHAPTER 5. NUMERICAL RESULTS

We can tell from Table 5.6 that we obtain less accurate result in estimating the correct break points by using least square method when we have intervals with nonzero size. It turns out that the minimum is always attained at the boundary. Moreover, we revise the break point by using the properties of $\hat{r}_a$ and $\hat{r}_b$. In Section 4.1, we derive from the study of Breitung and Eickmeier (2011)[10] that the sum of pre- and post-break factors $k$ is minimized at the true break point. Thus, for each given interval, we compute $\hat{r}_a + \hat{r}_b$ for each potential break point. Then we check whether the minimum of the sum over this interval is obtained at $T_c = 2007: M12$. If so, we set the revised break point equal to $T_c$. Otherwise, we adjust the revised break point as the closest point to $T_c$ where the minimum is attained. For all four cases in the analysis, the break points we obtained match the conjectured break point $T_c$ (Table 5.6). Therefore, the procedure has strong capability of locating unknown break point and estimating the number of factors.

Next, we applied bootstrap method based on the original dataset SW132. According to Efron and Tibshirani (1993)[16], Bootstrap is a recently developed simulation methodology using computer power to realize statistical inferences and assign measures of accuracy to statistical estimates. The process of bootstrap includes two steps. First, we draw a bootstrap sample, which has the same size as the initial observations, with replacement. Then, we repeat this process a large number of times to get bootstrap sample replicates which can be used to make inferences (Efron and Tibshirani, 1993)[16]. This technique requires fewer assumptions and offer greater accuracy and insight than standard methods in many fields (Stine, 1989)[30].
Concretely, we first generate 10000 bootstrap samples, then obtain the number of factors \( \hat{r}_a, \hat{r}_b \), and the occurrence of break point \( \hat{B} \) from each sample by applying the same procedure. We only focus on the situations that the break point is unknown. To calculate the probability of correctly estimating the number of factors, we round off the mean of \( \hat{r}_a \) and \( \hat{r}_a \) and set those value as \( r_a \) and \( r_b \) respectively. The value of MSFE shows the behaviour of the procedure. The result is displayed in Table 5.7 below.

<table>
<thead>
<tr>
<th>Model Configuration</th>
<th>Model Selection</th>
<th>Relative</th>
</tr>
</thead>
<tbody>
<tr>
<td>( r_a ) ( r_b )</td>
<td>( \Pr(\hat{B} = B_0) )</td>
<td>( \Pr(\hat{r}_a = r_a) )</td>
</tr>
<tr>
<td>3 3 102 337</td>
<td>0.85</td>
<td>0.41</td>
</tr>
</tbody>
</table>

From Table 5.7, it is obvious that we tend to properly detect the break point as the rate of success is higher that 85% for unknown break point. However, we have relatively lower probability of accessing the number of pre- and post- break factors which is around 0.4-0.5. In terms of the MSFE, the value is 0.85 which is less than 1. Therefore, for the bootstrap samples, the proposed PMS estimators domain the corresponding full sample estimators.
Chapter 6

Conclusion

In case of macroeconomics, one may be interested in studying the occurrence of the structural break, such as the 2007-2009 Great Recession, and its influence to the financial market and the whole economic environment. In order to eliminate the negative effect of the break point and maintain the economic stability, one needs to keep track of any new factors and the change of existing factors after the break point.

A high-dimensional model system is adopted to solve this problem. A shrinkage estimation procedure identifies the occurrence of the break point and estimates the numbers of pre- and post-break factors. The estimator works well as it is weakly dependent on the instabilities of unknown break point case. When the number of factors remains unchanged, the procedure can detect the changes in the factor loadings. We also find out that the occurrence of the break point can be identified by using a conventional least squares approach once we obtain the estimation of the numbers of pre- and post-break factors.
We applied Monte Carlo Simulation to analyze the behaviour of the procedure. It turns out that the shrinkage procedure has capability of determining the number of factors and testing for a break point for both known- and unknown-break-point case. In the empirical analysis based on the 2007-2009 Great Recession data set, the procedure can properly detect the break point in the most of the cases.
Appendix A

Some Statistical Background

In this appendix, we give some definitions, lemmas and theorems used in deriving the main results of this major paper. Most of the definitions, can be found in statistical textbooks, such as

**Definition A.1** (Casella and Berger (2002)[12]). A sequence of random variables \( \{X_n\}_{n=1}^{\infty} \) converges in probability to a random variable \( X \) if, for every \( \epsilon > 0 \),

\[
\lim_{n \to \infty} P(|X_n - X| \geq \epsilon) = 0.
\]

We denote it as \( X_n \overset{p}{\to} X \).

**Definition A.2** (Bickel and Doksum (2001)[8]). A sequence of random vectors \( Z_n = (Z_{n1}, Z_{n2}, \cdots, Z_{nm})' \) converges in probability to \( Z = (Z_1, Z_2, \cdots, Z_m)' \) iff \( Z_{nj} \overset{p}{\to} Z_j \) for \( 1 \leq j \leq m \). We denote it as \( Z_n \overset{p}{\to} Z \).
Definition A.3 (Fuller (1976)[18]). Let \( \{A_n\}_{n=1}^{\infty} \) be a sequence of random matrices and \( \{b_n\}_{n=1}^{\infty} \) be a sequence of positive real numbers.

(i) \( A_n = o_p(b_n) \) if for all \( i \) and \( j \) \( (A_n)_{ij} = o_p(b_n) \), that is, \( \forall \epsilon > 0 \), there exists \( k_\epsilon > 0 \) and an integer \( N_\epsilon \), such that \( P\left( |(A_n)_{ij}| > k_\epsilon b_n \right) < \epsilon \), \( \forall n > N_\epsilon \).

(ii) \( A_n = O_p(b_n) \) if for all \( i \) and \( j \) \( (A_n)_{ij} = O_p(b_n) \), that is, \( \left( \frac{(A_n)_{ij}}{b_n} \right) \xrightarrow{P} 0 \) as \( n \to \infty \).

Lemma A.1 (Strawderman (1994)[32]). Let \( A_n \) be a random sequence of symmetric nonnegative definite \( k \times k \) matrices where \( k < \infty \). If a positive definite symmetric \( k \times k \) matrix \( A \) with finite elements exists such that \( A_n \xrightarrow{P} A \) element-wise, then \( \|A_n - A\| \xrightarrow{P} 0 \), where \( \| \cdot \| \) denotes any proper norm on \( \mathbb{R}^{k \times k} \).

The proof of this lemma is given as Lemma 1 in Strawderman (1994)[32].
Appendix B

Some Proofs

In this appendix, we give the proof of one of the theorems used in the major paper.

**Proof of Theorem 3.4.** We need to consider two situations: there exists a break point with the instability of change and there is no change. In each situation, we have to prove that $\lim_{N,T \to \infty} \Pr(\hat{r}_a = r_a) = 1$, $\lim_{N,T \to \infty} \Pr(\hat{r}_b = r_b) = 1$ and $\lim_{N,T \to \infty} \Pr(\hat{B} = B_0) = 1$.

First, we consider the change with instability. Let $\Sigma_a = \Lambda^0 \Lambda^0 / N \in R^{r_a \times r_a}$, where $\Sigma_a^{1/2}$ is the Cholesky factor of $\Sigma_a$. Let $\Upsilon_a = [v_1 \ v_2 \ \cdots \ v_{r_a}]$ be a matrix of orthonormal eigenvectors such that $\Upsilon_a \Sigma_a^{1/2} \Sigma_a^{1/2} \Upsilon_a = V_a$. Suppose we have the eigenvalues of $(\Sigma_a^{1/2})' \Sigma_F \Sigma_a^{1/2}$ ordered from largest to smallest denoted by $\rho_\ell$, $\ell = 1, \cdots, r_a$. Therefore, by definition, $V_a$ is a diagonal matrix of eigenvalues of $(\Sigma_a^{1/2})' \Sigma_F \Sigma_a^{1/2}$ ordered from largest to smallest.

Further, one can verify that $\rho_\ell$, the $\ell$-th largest eigenvalue of $(\Sigma_a^{1/2})' \Sigma_F \Sigma_a^{1/2}$, is also the
\( \ell \)-th largest eigenvalue of matrix \( \Sigma_a \Sigma_F \). Then, we define the transformation matrix \( R_a \) and \( R_b \) in (3.3) such that \( R_a = \frac{1}{2} \Sigma_a^{1/2} \Upsilon_a \Sigma_a^{-1/2} \) and \( R_b = \frac{1}{2} \Sigma_b^{1/2} \Upsilon_b \Sigma_b^{-1/2} \). Moreover, we have defined that \( \Lambda^R = \Lambda_0 (R_a^{-1})' \in \mathbb{R}^{N \times r_a} \) and \( \Psi^R = \Psi_0 (R_b^{-1})' \in \mathbb{R}^{N \times r_b} \) in (3.3) and (3.9).

Then, we have

\[
\frac{\Lambda^R \Lambda^R}{N} = \frac{R_a^{-1} \Lambda_0' \Lambda_0 (R_a^{-1})'}{N} = V_a^{1/2} \Sigma_a^{-1/2} \Lambda_0' \Lambda_0 \Sigma_a^{-1/2} \Upsilon_a V_a^{1/2} = V_a \tag{B.1}
\]

By Assumption 3, we have \( \| \Sigma_a - \Sigma_\Lambda \| \to 0 \) as \( N \to \infty \). Since \( \Sigma_a \) is a sequence of symmetric positive definite matrices, we can apply Lemma 1 from Strawderman (1993)[32] to obtain \( \Sigma_a \overset{P}{\rightarrow}_{N \to \infty} \Sigma_\Lambda \). Then we multiply the positive definite matrix \( \Sigma_F \) on both sides to get \( \Sigma_a \Sigma_F \overset{P}{\rightarrow}_{N \to \infty} \Sigma_\Lambda \Sigma_F \). Since the convergence of matrices is entry-wise (de Boor, 2002)[9], we have for all \( \epsilon > 0 \),

\[
\lim_{N,T \to \infty} \Pr \left( \left| (\Sigma_a \Sigma_F)_{ij} - (\Sigma_\Lambda \Sigma_F)_{ij} \right| \leq \epsilon \right) = 1. \tag{B.2}
\]

where the \( ij \)-subscript represents the entry of the matrix located in \( i \)-th row and \( j \)-th column.

We apply the property that the eigenvalues of a matrix depend continuously to its entries (Alexanderian 2013)[2], we get

\[
\rho_\ell (\Sigma_a \Sigma_F) \overset{P}{\rightarrow}_{N \to \infty} \rho_\ell (\Sigma_\Lambda \Sigma_F). \tag{B.3}
\]
Therefore, each eigenvalue of $\Sigma_a \Sigma_F$ ordered from largest to smallest, which is also the $\ell$-th diagonal elements of $V_a$, converges in probability to the $\ell$-th largest eigenvalue of $\Sigma_a \Sigma_F$ as $N \to \infty$, denoted by $\rho_{\ell}(\Sigma_a \Sigma_F)$.

Let $a_{\ell}$ be a selection vector that selects the $\ell$-th column of a matrix. Similarly, $a'_{\ell}$ selects the $\ell$-th row of a matrix. Then, for $\ell = 1, \cdots, r_a$,

$$N^{-1}\|\Lambda^R_\ell\|^2 = N^{-1}(a'_\ell \Lambda^R \Lambda^R_\ell) = N^{-1}(a'_\ell \Lambda^R \Lambda^R_\ell) = a'_\ell (N^{-1} \Lambda^R \Lambda^R_\ell) a_{\ell} = a'_\ell V_a a_{\ell}$$

$$= \rho_{\ell}(\Sigma_a \Sigma_F) + o(1). \quad (B.4)$$

Here, $a'_\ell V_a a_{\ell}$ means the $\ell$-th diagonal element of $V_a$, which equals to the $\ell$-th largest eigenvalue of $\Sigma_a \Sigma_F$. Equation (B.4) is equivalent to

$$N^{-1/2}\|\Lambda^R_\ell\| = \left[\rho_{\ell}(\Sigma_a \Sigma_F)\right]^{1/2} + o(1). \quad (B.5)$$

From Theorem 3.3(i), we obtain

$$N^{-1/2}\|\hat{\Lambda}_\ell - \Lambda^R_\ell\| = O_p(\frac{1}{\sqrt{N_T}}) \text{ for } \ell = r_a. \quad (B.6)$$

By applying the triangle inequality on (B.6), we can get

$$\left|N^{-1/2}\|\hat{\Lambda}_\ell\| - N^{-1/2}\|\Lambda^R_\ell\|\right| \leq N^{-1/2}\|\hat{\Lambda}_\ell - \Lambda^R_\ell\| = O_p(\frac{1}{\sqrt{N_T}}). \quad (B.7)$$

We get rid of the absolute value and only retain the left side. Then, based on (B.5),
we obtain the following result

\[-N^{-1/2}\|\hat{\Lambda}_\ell - \Lambda^R_\ell\| \leq N^{-1/2}\|\hat{\Lambda}_\ell\| - N^{-1/2}\|\Lambda^R_\ell\|\]

\[N^{-1/2}\|\Lambda^R_\ell\| - N^{-1/2}\|\hat{\Lambda}_\ell - \Lambda^R_\ell\| \leq N^{-1/2}\|\hat{\Lambda}_\ell\|\]

\[\left[\rho_\ell(\Sigma_\Lambda \Sigma_F)\right]^{1/2} + o(1) \leq N^{-1/2}\|\hat{\Lambda}_\ell\| + O_p(C_{NT}^{-1}).\]  \hspace{1cm} (B.8)

Therefore, we conclude that for \(\ell = r_a\), \(\lim_{N,T\to\infty} \Pr(\|\hat{\Lambda}_\ell\| > 0) = 1\). Here, it shows that the \(r_a\)-th column of \(\hat{\Lambda}\) has value greater than 0 with probability approaching to 1. Recall the definition from (3.10), \(\hat{r}_a\) is the largest column number of \(\hat{\Lambda}\) which has value greater than 0. Thus,

\[\lim_{N,T\to\infty} \Pr(\hat{r}_a \geq r_a) = 1.\]  \hspace{1cm} (B.9)

From Theorem 3.3(ii), we know that

\[\lim_{N,T\to\infty} \Pr(\|\hat{\Lambda}_\ell\|^2 = 0) = 1\] for \(\ell = r_a + 1, \cdots, k\).  \hspace{1cm} (B.10)

Hence, we have the knowledge that the columns of \(\hat{\Lambda}\) from \(r_a + 1\) to \(k\) has value 0 with probability approaching to 1. In this case,

\[\lim_{N,T\to\infty} \Pr(\hat{r}_a \leq r_a) = 1.\]  \hspace{1cm} (B.11)

By (B.9) and (B.11), we obtain

\[\lim_{N,T\to\infty} \Pr(\hat{r}_a = r_a) = 1.\]  \hspace{1cm} (B.12)

Note that by definition in (3.10), we have \(\hat{r}_b \geq \hat{r}_a\). Together with (B.12) and the
APPENDIX B. SOME PROOFS

condition \( r_a = r_b \), we have

\[
\lim_{N,T \to \infty} \Pr(\hat{r}_b \geq r_b) = 1. \tag{B.13}
\]

Similar as (B.11), Theorem 3.3(v) implies that

\[
\lim_{N,T \to \infty} \Pr(\hat{r}_b \leq r_b) = 1. \tag{B.14}
\]

From (B.13) and (B.13) we conclude that

\[
\lim_{N,T \to \infty} \Pr(\hat{r}_b = r_b) = 1. \tag{B.15}
\]

For the second situation, with the existence of the instability, we have \( r_a = r_b \) and \( B_0 = 1 \). First note that if \( r_a = r_b \), we have

\[
N^{-1} \Gamma^R R = N^{-1} (\Psi^R - \Lambda^R)'(\Psi^R - \Lambda^R) = e' \Sigma^+_{\Lambda \Psi} e + o(1), \tag{B.16}
\]

where \( e = \lim_{N \to \infty} (R_a^{-1}, -R_b^{-1})' \) has full rank following Assumption 2 and 3. \( \Sigma^+_{\Lambda \Psi} \) is defined in (3.1). By a Cholesky decomposition, write \( \Sigma^+_{\Lambda \Psi} = (\Sigma^+_{\Lambda \Psi})^{1/2}(\Sigma^+_{\Lambda \Psi})^{1/2} \) with \( \text{rank}((\Sigma^+_{\Lambda \Psi})^{1/2}) = \text{rank}(\Sigma^+_{\Lambda \Psi}) > r_a \). For a \( 2r_a \times 2r_a \) matrix \( (\Sigma^+_{\Lambda \Psi})^{1/2} \), the rank of the null space of \( (\Sigma^+_{\Lambda \Psi})^{1/2} \) is smaller than \( r_a \). It follows that \( (\Sigma^+_{\Lambda \Psi})^{1/2} e \neq 0 \) because \( \text{rank}(e) = r_a \), and this immediately implies that with \( \Sigma_{\Gamma} = e' \Sigma^+_{\Lambda \Psi} e \neq 0, r_a > r_b \) and \( \text{rank}(\Sigma^+_{\Lambda \Psi}) > r_a \), then

\[
N^{-1} \Gamma^R R \to \Sigma_{\Gamma} \text{ for some } \Sigma_{\Gamma} \neq 0 \text{ as } N \to \infty. \tag{B.17}
\]
Then, we apply the selection vector again,

\[ N^{-1} \| \Gamma^R_\ell \|^2 = N^{-1} \| \Gamma^R a_\ell \|^2 = N^{-1} \| \Psi^R a_\ell - \Lambda^R a_\ell \|^2 = \left[ N^{-1/2} \| \Psi^R a_\ell - \Lambda^R a_\ell \| \right]^2 \]

\[ \geq \left( N^{-1/2} \| \Psi^R a_\ell \| - N^{-1/2} \| \Lambda^R a_\ell \| \right)^2 \]

\[ = \left[ (a'_\ell V_b a_\ell)^{1/2} - (a'_\ell V_a a_\ell)^{1/2} \right]^2 \]

\[ = \left[ (\rho_\ell (\Sigma_\Psi \Sigma_T))^1/2 - (\rho_\ell (\Sigma_\Lambda \Sigma_F))^1/2 \right]^2 + o(1). \]

\[(B.18)\]

where the inequality follows from the triangle inequality, and the last equality holds since we can draw the similar conclusion from (B.4). Therefore, we have

\[ N^{-1/2} \| \Gamma^R_\ell \| \geq \left| (\rho_\ell (\Sigma_\Psi \Sigma_T))^1/2 - (\rho_\ell (\Sigma_\Lambda \Sigma_F))^1/2 \right| + o(1) \text{ for } \ell = 1, \cdots, r_a. \]

\[(B.19)\]

Recall Theorem 3.3(iii), we known that for \( \ell = 1, \cdots, r_b \) and \( \Gamma^0 \neq 0, \)

\[ N^{-1/2} \| \hat{\Gamma}_\ell - \Gamma^R_\ell \| = O_p(C^{-1}_{NT}). \]

\[(B.20)\]

Then, based on (B.17), (B.19) and (B.20) and Assumption 1, we have for \( \ell = 1, \cdots, r_b, \)

\[ \left| N^{-1/2} \| \hat{\Gamma}_\ell \| - N^{-1/2} \| \Gamma^R_\ell \| \right| \leq N^{-1/2} \| \hat{\Gamma}_\ell - \Gamma^R_\ell \| = O_p(C^{-1}_{NT}), \]

\[ -N^{-1/2} \| \hat{\Gamma}_\ell - \Gamma^R_\ell \| \leq N^{-1/2} \| \hat{\Gamma}_\ell \| - N^{-1/2} \| \Gamma^R_\ell \|, \]

\[ N^{-1/2} \| \Gamma^R_\ell \| - N^{-1/2} \| \hat{\Gamma}_\ell - \Gamma^R_\ell \| \leq N^{-1/2} \| \hat{\Gamma}_\ell \|, \]

\[ \left| (\rho_\ell (\Sigma_\Psi \Sigma_T))^1/2 - (\rho_\ell (\Sigma_\Lambda \Sigma_F))^1/2 \right| + o(1) \leq N^{-1/2} \| \hat{\Gamma}_\ell \| + O_p(C^{-1}_{NT}). \]
Therefore,
\[
\lim_{N,T \to \infty} \Pr(\|\hat{\Gamma}_\ell\| > 0) = 1 \text{ for } \ell = 1, \cdots, r_b. \tag{B.21}
\]
Then, we consider the relation of two events \(\{\|\hat{\Gamma}_\ell\| > 0\} \subset \{\|\hat{\Gamma}\| > 0\}\). Then we have
\[
\Pr(\|\hat{\Gamma}_\ell\| > 0) \leq \Pr(\|\hat{\Gamma}\| > 0),
\]
then
\[
\lim_{N,T \to \infty} \Pr(\|\hat{\Gamma}_\ell\| > 0) \leq \lim_{N,T \to \infty} \Pr(\|\hat{\Gamma}\| > 0) \leq 1,
\]
and then
\[
1 \leq \lim_{N,T \to \infty} \Pr(\|\hat{\Gamma}\| > 0) \leq 1,
\]
this gives
\[
\lim_{N,T \to \infty} \Pr(\|\hat{\Gamma}\| > 0) = 1. \tag{B.22}
\]
Moreover, since \(\Pr(\|\hat{\Gamma}\| > 0) + \Pr(\|\hat{\Gamma}\| = 0) = 1\) and the two events \(\{\|\hat{\Gamma}\| > 0\}\) and \(\{\|\hat{\Gamma}\| = 0\}\) are disjoint, we apply the law of total probability, (B.21) and (B.22) to obtain
\[
\Pr(I_{\{\|\hat{\Gamma}_\ell\| > 0\}} = 1) = \Pr(\{I_{\{\|\hat{\Gamma}_\ell\| > 0\}} = 1\} \cap \{\|\hat{\Gamma}\| > 0\})
+ \Pr(\{I_{\{\|\hat{\Gamma}_\ell\| > 0\}} = 1\} \cap \{\|\hat{\Gamma}\| = 0\}),
\]
\[
\Pr(I_{\{\|\hat{\Gamma}_\ell\| > 0\}} = 1) = \Pr(\{\|\hat{\Gamma}_\ell\| > 0\} \cap \{\|\hat{\Gamma}\| > 0\})
+ \Pr(\{\|\hat{\Gamma}_\ell\| > 0\} \cap \{\|\hat{\Gamma}\| = 0\}),
\]
\[
\Pr(I_{\{\|\hat{\Gamma}_\ell\| > 0\}} = 1) = \Pr(\{\|\hat{\Gamma}_\ell\| > 0\}) + \Pr(\{\|\hat{\Gamma}_\ell\| > 0\} \cap \{\|\hat{\Gamma}\| = 0\})
+ \Pr(\{\|\hat{\Gamma}_\ell\| > 0\} \cap \{\|\hat{\Gamma}\| = 0\}).
\]
\[ \lim_{N,T \to \infty} \Pr(\mathcal{I}_{\{\|\hat{\Gamma}_{\ell}\| > 0\}}) = 1 = \lim_{N,T \to \infty} \Pr\left(\{\|\hat{\Gamma}_{\ell}\| > 0\}\right) \]
\[ + \lim_{N,T \to \infty} \Pr\left(\{\|\hat{\Gamma}_{\ell}\| > 0\} \cap \{\|\hat{\Gamma}\| = 0\}\right), \]

Therefore,

\[ \lim_{N,T \to \infty} \Pr(\mathcal{I}_{\{\|\hat{\Gamma}_{\ell}\| > 0\}}) = 1. \quad (B.23) \]

Similarly, we have \( \{\mathcal{I}_{\{\|\hat{\Gamma}\| > 0\}} = 1\} \subset \{\mathcal{I}_{\{\|\hat{\Gamma}\| > 0\}} = 1\} \) as well as (B.23), which imply that

\[ \Pr(\mathcal{I}_{\{\|\hat{\Gamma}\| > 0\}} = 1) \leq \Pr(\mathcal{I}_{\{\|\hat{\Gamma}\| > 0\}} = 1) \leq 1 \]
\[ \lim_{N,T \to \infty} \Pr(\mathcal{I}_{\{\|\hat{\Gamma}\| > 0\}} = 1) \leq \lim_{N,T \to \infty} \Pr(\mathcal{I}_{\{\|\hat{\Gamma}\| > 0\}} = 1) \leq 1 \]
\[ 1 \leq \lim_{N,T \to \infty} \Pr(\mathcal{I}_{\{\|\hat{\Gamma}\| > 0\}} = 1) \leq 1. \]

Therefore,

\[ \lim_{N,T \to \infty} \Pr(\mathcal{I}_{\{\|\hat{\Gamma}\| > 0\}} = 1) = 1. \quad (B.24) \]

By (B.24) and the definition of \( \hat{\mathcal{B}} \) in (3.8), we conclude that

\[ \lim_{N,T \to \infty} \Pr(\hat{\mathcal{B}} = \mathcal{B}_0 = 1) = 1. \quad (B.25) \]

Hence, by taking into account for the break point and the instability, we have Theorem 3.3 proved by (B.12), (B.25) and (B.15).

Now, for the case where there is no change, i.e. \( r_a = r_b \) and \( \mathcal{B}_0 = 0 \), Theorem 3.3(vi)
and (v) imply that

\[ \lim_{N,T \to \infty} \Pr(\|\hat{\Gamma}_\ell\| = 0) = 1 \text{ for } \ell = 1, \ldots, k, \]  

(B.26)

By applying the same procedure from (B.23) and (B.24), we conclude that

\[ \lim_{N,T \to \infty} \Pr(I_{\{\|\hat{\Gamma}\| > 0\}} = 0) = 1. \]  

(B.27)

Thus, by the definition of \( \hat{\mathcal{B}} \) in (3.8) and the fact that \( \mathcal{B}_0 = 0 \), we have

\[ \lim_{N,T \to \infty} \Pr(\hat{\mathcal{B}} = \mathcal{B}_0 = 0) = 1. \]  

(B.28)

Then, the definition of \( \hat{r}_b \) in (3.10), the fact that \( r_a = r_b \) and (B.12) imply that

\[ \lim_{N,T \to \infty} \Pr(\hat{r}_b = r_b) = 1. \]  

(B.29)

Similarly, when there is no change, we have Theorem 3.4 proved based on (B.12), (B.28) and (B.29).
Appendix C

Supplemental Tables for the Great Recession Data Set

For the convenience of the reader, we report Table C.1 to C.4 below, which give the financial indicators of the Great Recession dataset analysed in Section 5.2. The reported tables are the same as given in Cheng et al. (2016, see Supplemental Appendix Tables S3-S5[14]).

Table C.1: List of Indicators - Part I

<table>
<thead>
<tr>
<th>Name</th>
<th>Long Description</th>
</tr>
</thead>
<tbody>
<tr>
<td>Cons: Dur</td>
<td>Real Personal Consumption Expenditures: Durable Goods</td>
</tr>
<tr>
<td>Cons: Svc</td>
<td>Real Personal Consumption Expenditures: Services</td>
</tr>
<tr>
<td>Cons: NonDur</td>
<td>Real Personal Consumption Expenditures: Non-durable Goods</td>
</tr>
<tr>
<td>Real InvtCh</td>
<td>Component for Change in Private Inventories, deflated by JCXFE</td>
</tr>
<tr>
<td>Real WageG</td>
<td>Component for Government GDP: Wage and Salary Disbursements by Industry, Government, deflated by JCXFE</td>
</tr>
</tbody>
</table>
### APPENDIX C. SUPPLEMENTAL TABLES FOR THE GREAT RECESSION DATA SET

Table C.2: List of Indicators - Part II

<table>
<thead>
<tr>
<th>Name</th>
<th>Long Description</th>
</tr>
</thead>
<tbody>
<tr>
<td>IP: DurGds materials</td>
<td>Industrial Production: Durable Materials</td>
</tr>
<tr>
<td>IP: NondurGds materials</td>
<td>Industrial Production: Nondurable Materials</td>
</tr>
<tr>
<td>IP: Auto</td>
<td>IP: Automotive products</td>
</tr>
<tr>
<td>IP: NonDurConsGoods</td>
<td>Industrial Production: Nondurable Consumer Goods</td>
</tr>
<tr>
<td>IP: BusEquip</td>
<td>Industrial Production: Business Equipment</td>
</tr>
<tr>
<td>CapU Tot</td>
<td>Capacity Utilization: Total Industry</td>
</tr>
<tr>
<td>CapU Man</td>
<td>Capacity Utilization: Manufacturing (FRED past 1972)</td>
</tr>
<tr>
<td>Emp: DurGoods</td>
<td>All Employees: Durable Goods Manufacturing</td>
</tr>
<tr>
<td>Emp: Const</td>
<td>All Employees: Construction</td>
</tr>
<tr>
<td>Emp: Edu &amp; Health</td>
<td>All Employees: Education &amp; Health Services</td>
</tr>
<tr>
<td>Emp: Finance</td>
<td>All Employees: Financial Activities</td>
</tr>
<tr>
<td>Emp: Infor</td>
<td>All Employees: Information Services</td>
</tr>
<tr>
<td>Emp: Bus Serv</td>
<td>All Employees: Professional &amp; Business Services</td>
</tr>
<tr>
<td>Emp: Leisure</td>
<td>All Employees: Leisure &amp; Hospitality</td>
</tr>
<tr>
<td>Emp: OtherSvcs</td>
<td>All Employees: Other Services</td>
</tr>
<tr>
<td>Emp: Mining/NatRes</td>
<td>All Employees: Natural Resources &amp; Mining</td>
</tr>
<tr>
<td>Emp: Trade&amp;Trans</td>
<td>All Employees: Trade, Transportation &amp; Utilities</td>
</tr>
<tr>
<td>Emp: Retail</td>
<td>All Employees: Retail Trade</td>
</tr>
<tr>
<td>Emp: Wholesale</td>
<td>All Employees: Wholesale Trade</td>
</tr>
<tr>
<td>Emp: Gov(Fed)</td>
<td>All Employees: Government: Federal</td>
</tr>
<tr>
<td>Emp: Gov (State)</td>
<td>All Employees: Government: State Government</td>
</tr>
<tr>
<td>Emp: Gov (Local)</td>
<td>All Employees: Government: Local Government</td>
</tr>
<tr>
<td>URate: Age16-19</td>
<td>Unemployment Rate - 16-19 yrs</td>
</tr>
<tr>
<td>URate: Age &gt; 20 Men</td>
<td>Unemployment Rate - 20 yrs. &amp; over, Men</td>
</tr>
<tr>
<td>URate: Age &gt; 20 Women</td>
<td>Unemployment Rate - 20 yrs. &amp; over, Women</td>
</tr>
<tr>
<td>U: Dur &lt; 5wks</td>
<td>Number Unemployed for Less than 5 Weeks</td>
</tr>
<tr>
<td>U: Dur 5-14wks</td>
<td>Number Unemployed for 5-14 Weeks</td>
</tr>
<tr>
<td>U: Dur &gt; 15-26wks</td>
<td>Civilians Unemployed for 15-26 Weeks</td>
</tr>
<tr>
<td>U: Dur &gt; 27wks</td>
<td>Number Unemployed for 27 Weeks &amp; over</td>
</tr>
<tr>
<td>U: JobLosers</td>
<td>Unemployment Level - Job Losers</td>
</tr>
<tr>
<td>U: LF Reentry</td>
<td>Unemployment Level - Reentrants to Labor Force</td>
</tr>
<tr>
<td>U: JobLeavers</td>
<td>Unemployment Level - Job Leavers</td>
</tr>
<tr>
<td>U: NewEntrants</td>
<td>Unemployment Level - New Entrants</td>
</tr>
<tr>
<td>Emp: SlackWk</td>
<td>Employment Level - Part-Time for Economic Reasons, All Industries</td>
</tr>
<tr>
<td>AWH Man</td>
<td>Average Weekly Hours: Manufacturing</td>
</tr>
<tr>
<td>AWH Privat</td>
<td>Average Weekly Hours: Total Private Industrie</td>
</tr>
<tr>
<td>AWH Overtime</td>
<td>Average Weekly Hours: Overtime: Manufacturing</td>
</tr>
<tr>
<td>Indicator</td>
<td>Description</td>
</tr>
<tr>
<td>-----------------------------------------------</td>
<td>------------------------------------------------------------------------------</td>
</tr>
<tr>
<td>HPermits</td>
<td>New Private Housing Units Authorized by Building Permit</td>
</tr>
<tr>
<td>Hstarts: MW</td>
<td>Housing Starts in Midwest Census Region</td>
</tr>
<tr>
<td>Hstarts: NE</td>
<td>Housing Starts in Northeast Census Region</td>
</tr>
<tr>
<td>Hstarts: S</td>
<td>Housing Starts in South Census Region</td>
</tr>
<tr>
<td>Hstarts: W</td>
<td>Housing Starts in West Census Region</td>
</tr>
<tr>
<td>Constr. Contracts</td>
<td>Construction contracts (mil. sq. ft.) (Copyright, McGraw-Hill)</td>
</tr>
<tr>
<td>IP: DurConsGoods</td>
<td>Industrial Production: Durable Consumer Goods</td>
</tr>
<tr>
<td>Ret. Sale</td>
<td>Sales of retail stores (mil. Chain 2000 $)</td>
</tr>
<tr>
<td>Orders (DurMfg)</td>
<td>Mfrs new orders durable goods industries (bil. chain 2000 $)</td>
</tr>
<tr>
<td>Orders (ConsumerGoods/Mat.)</td>
<td>Mfrs new orders, consumer goods and materials (mil. 1982 $)</td>
</tr>
<tr>
<td>UnfOrders (DurGds)</td>
<td>Mfrs unfilled orders durable goods indus. (bil. chain 2000 $)</td>
</tr>
<tr>
<td>Orders (NonDefCap)</td>
<td>Mfrs new orders, nondefense capital goods (mil. 1982 $)</td>
</tr>
<tr>
<td>VendPerf</td>
<td>Index of supplier deliveries vendor performance (pct.)</td>
</tr>
<tr>
<td>MT Invent</td>
<td>Manufacturing and trade inventories (bil. Chain 2005 $)</td>
</tr>
<tr>
<td>PCED-MotorVec</td>
<td>Motor vehicles and parts</td>
</tr>
<tr>
<td>PCED-DurHousehold</td>
<td>Furnishings and durable household equipment</td>
</tr>
<tr>
<td>PCED-Recreation</td>
<td>Recreational goods and vehicles</td>
</tr>
<tr>
<td>PCED-OthDurGds</td>
<td>Other durable goods</td>
</tr>
<tr>
<td>PCED-Food-Bev</td>
<td>Food and beverages purchased for off-premises consumption</td>
</tr>
<tr>
<td>PCED-Clothing</td>
<td>Clothing and footwear</td>
</tr>
<tr>
<td>PCED-Gas-Enrgy</td>
<td>Gasoline and other energy goods</td>
</tr>
<tr>
<td>PCED-OthNDurGds</td>
<td>Other nondurable goods</td>
</tr>
<tr>
<td>PCED-Housing-Utilities</td>
<td>Housing and utilities</td>
</tr>
<tr>
<td>PCED-HealthCare</td>
<td>Health care</td>
</tr>
<tr>
<td>PCED-TransSvc</td>
<td>Transportation services</td>
</tr>
<tr>
<td>PCED-RecServices</td>
<td>Recreation services</td>
</tr>
<tr>
<td>PCED-FoodServ-Acc.</td>
<td>Food services and accommodations</td>
</tr>
<tr>
<td>PCED-FIRE</td>
<td>Financial services and insurance</td>
</tr>
<tr>
<td>PCED-OtherServices</td>
<td>Other services</td>
</tr>
<tr>
<td>PPI: FinConsGds</td>
<td>Producer Price Index: Finished Consumer Goods</td>
</tr>
<tr>
<td>PPI: FinConsGds(Food)</td>
<td>Producer Price Index: Finished Consumer Foods</td>
</tr>
<tr>
<td>PPI: IndCom</td>
<td>Producer Price Index: Industrial Commodities</td>
</tr>
<tr>
<td>PPI: IntMat</td>
<td>Producer Price Index: Intermediate Materials: Supplies &amp; Components</td>
</tr>
<tr>
<td>NAPM ComPrice</td>
<td>NAPM COMMODITY PRICES INDEX (PERCENT)</td>
</tr>
<tr>
<td>Real Price: NatGas</td>
<td>PPI: Natural Gas, deflated by PCEPILFE</td>
</tr>
<tr>
<td>Real Price: Oil</td>
<td>PPI: Crude Petroleum, deflated by PCEPILFE</td>
</tr>
<tr>
<td>FedFunds</td>
<td>Effective Federal Funds Rate</td>
</tr>
</tbody>
</table>
Table C.4: List of Indicators - Part IV

<table>
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<th>Indicator</th>
<th>Description</th>
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<td>TB-3Mth</td>
<td>3-Month Treasury Bill: Secondary Market Rate</td>
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<tr>
<td>BAA-GS10</td>
<td>BAA-GS10 Spread</td>
</tr>
<tr>
<td>MRTG-GS10</td>
<td>Mortg-GS10 Spread</td>
</tr>
<tr>
<td>TB6m-TB3m</td>
<td>tb6m-tb3m</td>
</tr>
<tr>
<td>GS1-TB3m</td>
<td>GS1-Tb3m</td>
</tr>
<tr>
<td>GS10-TB3m</td>
<td>GS10-Tb3m</td>
</tr>
<tr>
<td>CP-TB Spread</td>
<td>CP-Tbill Spread: CP3FM-TB3MS</td>
</tr>
<tr>
<td>Ted-Spread</td>
<td>MED3-TB3MS (Version of TED Spread)</td>
</tr>
<tr>
<td>Real C&amp;I Loan</td>
<td>Commercial and Industrial Loans at All Commercial Banks Defl by PCEPILFE</td>
</tr>
<tr>
<td>Real ConsLoans</td>
<td>Consumer (Individual) Loans at All Commercial Banks Outlier Code because of change in data in April 2010 see FRB H8 Reles Defl by PCEPILFE</td>
</tr>
<tr>
<td>Real NonRevCredit</td>
<td>Total Nonrevolving Credit Owned and Securitized, Outstanding Defl by PCEPILFE</td>
</tr>
<tr>
<td>Real LoansRealEst</td>
<td>Real Estate Loans at All Commercial Banks Defl by PCEPILFE</td>
</tr>
<tr>
<td>Real RevolvCredit</td>
<td>Total Revolving Credit Outstanding Defl by PCEPILFE</td>
</tr>
<tr>
<td>S&amp;P500</td>
<td>S&amp;PS COMMON STOCK PRICE INDEX: COMPOSITE (1941-43=10)</td>
</tr>
<tr>
<td>DJIA</td>
<td>COMMON STOCK PRICES: DOW JONES INDUSTRIAL AVERAGE</td>
</tr>
<tr>
<td>VXO</td>
<td>VXO (Linked by N. Bloom) .. Average daily VIX from 2009</td>
</tr>
<tr>
<td>Ex rate: Major</td>
<td>FRB Nominal Major Currencies Dollar Index (Linked to EXRUS in 1973:1)</td>
</tr>
<tr>
<td>Ex rate: Switz</td>
<td>FOREIGN EXCHANGE RATE: SWITZERLAND (SWISS FRANC PER USD)</td>
</tr>
<tr>
<td>Ex rate: Japan</td>
<td>FOREIGN EXCHANGE RATE: JAPAN (YEN PER USD)</td>
</tr>
<tr>
<td>Ex rate: UK</td>
<td>FOREIGN EXCHANGE RATE: UNITED KINGDOM (CENTS PER POUND)</td>
</tr>
<tr>
<td>EX rate: Canada</td>
<td>FOREIGN EXCHANGE RATE: CANADA (CAD PER USD)</td>
</tr>
</tbody>
</table>
Bibliography


Vita Auctoris

Miss. Wen (Wendy) Gao was born in 1994, Dalian, Liaoning, China. She graduated from Dalian No.1 High School in June 2012. From there she went to University of California at Davis, where she obtained a Bachelor of Science degree in Statistics. She is currently a candidate for the Master of Science degree in Statistics at the University of Windsor. She hopes to graduate in September 2018.