Note

Things to do:

1. Literature review and introduction. Need more references, need to mention synthetic interventions in the introduction or the discussion.
2. Need to mention possible generalizations in the Discussion.
3. Need to add weighted adjustment estimator to the data analysis.
4. Need to add simulations.
5. Update abstract.

Updates in this version

1. The synthetic intervention paper and comment are added in the introduction. Other parts haven't been revised.
2. SCM.R for SCM method and sim.R for simulation are added in Github.
3. Data analysis references are fixed. The structure is organized better.

Minimizing post shock forecasting error using disparate information

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Abstract

We develop a forecasting methodology for time series data that has undergone a shock. We still can provide credible forecasts for a time series in the presence of such systematic shocks by drawing from disparate time series that have undergone similar shocks for which post-shock outcome data is recorded. These disparate time series are assumed to have mechanistic similarities to the time series under study but are otherwise independent (Granger noncausal). The inferential goal of our forecasting methodology is to supplement observed time series data with post-shock data from the disparate time series in order to minimize average forecast risk.

1 Introduction

In this article we provide forecasting adjustment techniques with the goal of lowering overall forecast error when the time series under study has undergone a structural shock. It is unlikely that any forecast that previously gave successful predictions for the time series of interest will be able to accommodate the structural shock. However, all is not lost in this setting, one can integrate information from disparate time series that have previously undergone similar structural shocks to estimate the shock effect of the time series under study. One can then combine these past similar shock effects and add them to the present forecast to reduce the overall forecast error.

Improving forecasts through forecast combination has a rich history [Bates and Granger, 1969, Mundlak, 1978, Timmermann, 2006, Granger and Newbold, 2014]. The classical setting for the forecast combination problem is when there are competing forecasts for a single time series. **The following list of methods needs to fall in the forecast combination literature, not the time series pooling literature**. In this setting there are a plethora of methods for combining forecasts, e.g., (1) model averaging [Newbold and Harvey, 2002, Timmermann, 2006, Hansen, 2008], (2) model selection [Lee and Phillips, 2015, Greenaway-McGrevy, 2020], (3) time-series pooling [Mundlak, 1978, Zellner et al., 1991, Lee et al., 2020, Plessen, 2020]. Model averaging typically selects weights for models based on minimizing various loss functions whereas model selection chooses the model through minimization of those loss functions, see for example, Fosten and Greenaway-McGrevy [2019] **I do not think that this reference belongs here**. Classical forecast combination may fail when forecasting in the presence of structural shocks.

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In our post-shock setting we combine estimated quantities from different time series with the aim of lowering forecast error for a single time series under study. Our techniques are similar to those in time-series pooling and data integration. The literature of time-series pooling is mainly related to pooling cross-sectional panel data [Mundlak, 1978, Zellner et al., 1991, Fosten and Greenaway-McGrevy, 2019]. The issues about whether to assume homogeneity or heterogeneity of slope coefficients across individual units are confounded. Baltagi [2008] showed that homogeneity approach often outperform heterogeneity one in mean squared forecast error; while heterogeneity approach is more general to accommodate differences among units. **How does this relevant to what we are doing?**

Data integration for forecasting is a broad area of research including ideas from many areas. Lee et al. [2020] constructed a Bayesian hierarchical model embracing data integration to improve predictive precision of COVID-19 infection trajectories for different countries. A similar setup may be beneficial for post-shock prediction but may be too dependent upon model specification for the shock distribution. Plessen [2020] employed a data-mining approach to combine COVID-19 data from different countries as input to predict global net daily infections and deaths of COVID-19 using clustering. However, there is a tremendous amount of volatility in this form of COVID-19 data, and the fit of this prediction method may be improved with modeling structure or preprocessing of the donor pool. From a machine learning perspective, Agarwal et al. [2020] proposed a model-free synthetic intervention method to predict unobserved potential outcomes after different interventions given a donor pool of observed outcomes with given interventions. Among those methodologies, they share the implicit assumption that candidates from donor pool should be similar. In other words, caution should be dedicated to the construction of donor pool; otherwise, the result will be expected not to be robust to the noise introduced by the insimilar ones. *More to be added ...*

Need more of a transition to what we do here and why it is good. We develop and compare aggregation techniques in this post-shock setting and investigate settings for when they do and do not decrease mean squared prediction error. We assume a simple auto regressive data generating process similar to that in Blundell and Bond [1998] with a general random effects structure. The main idea is to first average the estimated shock effects from the disparate time series and then add the averaged estimated shock effect to the present forecast. When these time series are independent and the mean of shock effect distribution is large relative to its variance then this technique will reduce mean squared prediction error under the assumed model. Note that this methodology is not motivated with the goal of unbiased, asymptotically unbiased, or consistent estimation for the shock-effect of the time series under study. We consider three aggregation techniques: simple averaging, inverse-variance weighted averaging, and similarity weighting. The latter technique is similar to the weighting in synthetic control methodology [Abadie et al., 2010].

Need to discuss our example and our simulation results.

2 Setting

We will suppose that an analyst has time series data $(y_{i,t}, \mathbf{x}_{i,t})$, $t = 1, \ldots, T_i$, $i = 1, \ldots, n+1$, where $y_{i,t}$ is a scalar response and $\mathbf{x}_{i,t}$ is a vector of covariates that are revealed to the analyst prior to the observation of $y_{1,t}$. Suppose that the analyst is interested in forecasting $y_{1,t}$, the first time series in the collection. To gauge the performance of a procedure that produces forecasts $\{\hat{y}_{1,t}, t = 1, 2, \ldots\}$ given time horizon T_1, we consider forecast risk in the form of root mean squared error (RMSE),

$$R_T = \sqrt{\frac{1}{T}\sum_{t=1}^{T} \mathrm{E}(\hat{y}_{1,t} - y_{1,t})^2},$$

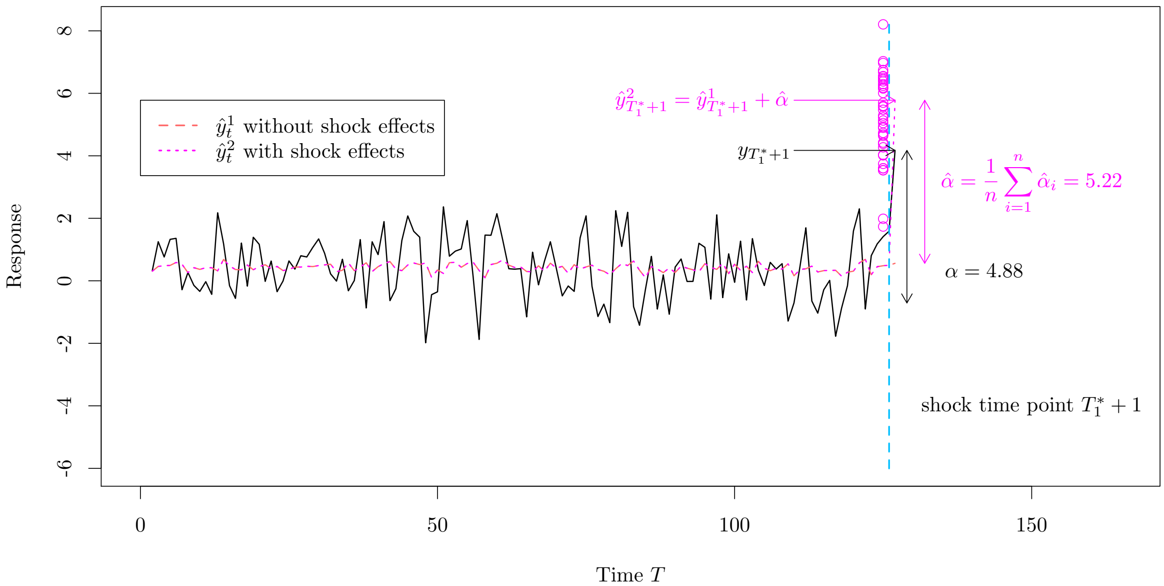

Figure 1. The time series experience a shock at $T_1^* + 1 = 126$ with true shock effect $\alpha = 4.88$. The figure is a comparison between forecast without considering shock effects and the one uses simple averaging given $n = 40$ disparate time series, and that the shock time is at $T_1^* + 1 = 126$. The magenta dots represent least square estimate $\hat{\alpha}_i$ from disparate time series. The prediction of $\hat{y}^2_{T_1^*+1}$ and $\hat{y}^1_{T_1^*+1}$ differs only by an adjustment $\hat{\alpha} = 5.22$. It is clear that $\hat{y}^2_{T_1^*+1}$ performs better than $\hat{y}^1_{T_1^*+1}$.

in our analyses. **Let's switch to root mean squared error, this metric is often used in forecasting articles.** In this article, we consider a dynamic panel data model with autoregressive structure similar to that in Blundell and Bond [1998]. Our dynamic panel model includes an additional shock effect whose presence or absence is given by the binary variable $D_{i,t}$, the details of this model are in the next section.

Figure 1 provides simple intuition of the practical usefulness of our proposed methodology. This figure depicts a time-series that experienced a "shock" at time point $T_1^* + 1 = 126$. It is supposed that the researcher does not have any information beyond $T_1^* + 1$, but does have observations of forty disparate time series that have previously undergone a similar shock for which post-shock responses are recorded. Similarity in this context means that the shock effects are random variables that from a common distribution. In this example, the mean of the estimated shock effects is taken as a shock-effect estimator for the time series under study. Forecasts are then made by adding this shock-effect estimator to the estimated response values obtained from the process that ignores the shock. It is apparent from Figure 1 that adjusting forecasts in this manner 1) leads to a reduction in forecasting risk; 2) does not fully recover the true shock-effect. We evaluate the performance of this post-shock prediction methodology throughout this article; we outline situations for when it is expected to work and when it is not.

2.1 Model Setup

In this section, we will describe the assumed dynamic panel models for which post-shock aggregated estimators are provided. The basic structure of these models are the same, the differences between them lie in the setup of the shock effect distribution.

The model $\mathcal{M}_1$ is defined as

$$\mathcal{M}_1 \colon y_{i,t} = \eta_i + \alpha_i D_{i,t} + \phi_i y_{i,t-1} + \theta_i' \mathbf{x}_{i,t} + \beta_i' \mathbf{x}_{i,t-1} + \varepsilon_{i,t} \tag{1}$$

for $t = 1, \ldots, T_i$ and $i = 1, \ldots, n+1$, where $D_{i,t} = 1(t = T_i^* + 1)$, $T_i^* < T_i$ and $\mathbf{x}_{i,t} \in \mathbb{R}^p$, $p \geq 1$. We assume that the $\mathbf{x}_{i,t}$'s are fixed and T_i^*s are known. The random effects structure for $\mathcal{M}_1$ is:

$$
\begin{aligned}
\eta_i &\overset{iid}{\sim} \eta, \text{ where } \mathrm{E}(\eta) = 0, \mathrm{Var}(\eta) = \sigma_\eta^2, \qquad i = 1, \ldots, n+1,\\
\phi_i &\overset{iid}{\sim} \phi, \text{ where } |\phi| < 1, \qquad i = 1, \ldots, n+1,\\
\theta_i &\overset{iid}{\sim} \theta, \text{ where } \mathrm{E}(\theta) = \mu_\theta, \mathrm{Var}(\theta) = \Sigma_\theta^2, \qquad i = 1, \ldots, n+1,\\
\beta_i &\overset{iid}{\sim} \beta, \text{ where } \mathrm{E}(\beta) = \mu_\beta, \mathrm{Var}(\beta) = \Sigma_\beta^2, \qquad i = 1, \ldots, n+1,\\
\varepsilon_{i,t} &\overset{iid}{\sim} \mathcal{N}(0, \sigma^2), \qquad t = 1, \ldots, T_i,\ i = 1, \ldots, n+1,\\
\alpha_i &\overset{iid}{\sim} \mathcal{N}(\mu_\alpha, \sigma_\alpha^2), \qquad i = 1, \ldots, n+1;\\
\eta &\perp\!\!\!\perp \alpha_i \perp\!\!\!\perp \phi \perp\!\!\!\perp \theta \perp\!\!\!\perp \varepsilon_{i,t}.
\end{aligned}
$$

Notice that $\mathcal{M}_1$ assumes that α_i are iid with $\mathrm{E}(\alpha_i) = \mu_\alpha$ for $i = 1, \ldots, n+1$. We also consider a model where the shock effects are linear functions of covariates and lagged covariates with an additional additive mean-zero error. The random effects structure for this model (model $\mathcal{M}_2$) is:

$$
\mathcal{M}_2\colon \begin{array}{l} y_{i,t} = \eta_i + \alpha_i D_{i,t} + \phi_i y_{i,t-1} + \theta_i' \mathbf{x}_{i,t} + \beta_i' \mathbf{x}_{i,t-1} + \varepsilon_{i,t} \\ \alpha_i = \mu_\alpha + \delta_i' \mathbf{x}_{i,T_i^*+1} + \gamma_i' \mathbf{x}_{i,T_i^*} + \tilde{\varepsilon}_i, \end{array} \tag{2}
$$

for $i = 1, \ldots, n+1$, where the added random effects are

$$
\begin{aligned}
\tilde{\varepsilon}_i &\overset{iid}{\sim} \mathcal{N}(0, \sigma_\alpha^2), \qquad i = 1, \ldots, n+1;\\
\eta &\perp\!\!\!\perp \alpha_i \perp\!\!\!\perp \phi \perp\!\!\!\perp \theta \perp\!\!\!\perp \varepsilon_{i,t} \perp\!\!\!\perp \tilde{\varepsilon}_i.
\end{aligned}
$$

We further define $\tilde{\alpha}_i = \mu_\alpha + \delta_i' \mathbf{x}_{i,T_i^*+1} + \gamma_i' \mathbf{x}_{i,T_i^*}$. We will investigate post-shock aggregated estimators in $\mathcal{M}_2$ in settings where δ_i and γ_i are either fixed or random. We let $\mathcal{M}_{21}$ denote model $\mathcal{M}_2$ with $\gamma_i = \gamma$ and $\delta_i = \delta$ for $i = 1, \ldots, n+1$, where γ and δ are fixed unknown parameters. We let $\mathcal{M}_{22}$ denote model $\mathcal{M}_2$ with the following random effects structure for γ and δ:

$$
\begin{array}{l} \gamma_i \overset{iid}{\sim} \mathrm{E}(\gamma) = \mu_\gamma, \mathrm{Var}(\gamma) = \Sigma_\gamma \\ \delta_i \overset{iid}{\sim} \mathrm{E}(\delta) = \mu_\delta, \mathrm{Var}(\delta) = \Sigma_\delta \end{array} \quad \text{with} \quad \delta_i \perp\!\!\!\perp \tilde{\varepsilon}_i \quad \text{and} \quad \gamma_i \perp\!\!\!\perp \tilde{\varepsilon}_i.
$$

Note that δ_i and γ_i may be dependent. We further define the parameter sets

$$
\begin{aligned}
\Theta &= \{(\eta_i, \phi_i, \theta_i, \beta_i, \alpha_i, \mathbf{x}_{i,t}, y_{i,t-1}, \delta_i, \gamma_i)\colon t = 1, \ldots, T_i, i = 2, \ldots, n+1\}.\\
\Theta_1 &= \{(\eta_i, \phi_i, \theta_i, \beta_i, \alpha_i, \mathbf{x}_{i,t}, y_{i,t-1}, \delta_i, \gamma_i)\colon t = 1, \ldots, T_i, i = 1\}.,
\end{aligned} \tag{3}
$$

where Θ and Θ_1 can adapt to $\mathcal{M}_1$ by dropping δ_i and γ_i. We assume this for notational simplicity.

2.2 Forecast

In this section we show how post-shock aggregate estimators improve upon standard forecasts that do not account for the shock effect. More formally, we will consider the following candidate forecasts:

$$
\begin{aligned}
&\text{Forecast 1} : \hat{y}^1_{1,T_1^*+1} = \hat{\eta}_1 + \hat{\phi}_1 y_{1,T_1^*} + \hat{\theta}_1' \mathbf{x}_{1,T_1^*+1} + \hat{\beta}_1' \mathbf{x}_{1,T_1^*},\\
&\text{Forecast 2} : \hat{y}^2_{1,T_1^*+1} = \hat{\eta}_1 + \hat{\phi}_1 y_{1,T_1^*} + \hat{\theta}_1' \mathbf{x}_{1,T_1^*+1} + \hat{\beta}_1' \mathbf{x}_{1,T_1^*} + \hat{\alpha},
\end{aligned}
$$

where $\hat{\eta}_1$, $\hat{\phi}_1$, $\hat{\theta}_1$, and $\hat{\beta}_1$ are all OLS estimators of η_1, ϕ_1, θ_1, and β_1 respectively, and $\hat{\alpha}$ is some form of estimator for the shock effect of time series of interest, i.e., α_1. The first forecast ignores the presence of α_1 while the second forecast incorporates an estimate of α_1 that is obtained from the other independent forecasts under study.

Note that the two forecasts do not differ in their predictions for $y_{1,t}$, $t = 1, \dots T_1^*$, they only differ in predicting y_{1,T_1^*+1}. Throughout the rest of this article we show that the collection of disparate time series $\{y_{i,t}, t = 2, \dots, T_i, i = 1, \dots, n\}$ has the potential to improve the forecasts for $y_{1,t}$ when $t > T_1^*$ under different circumstances for the dynamic panel model $\mathcal{M}_1$, $\mathcal{M}_{21}$, and $\mathcal{M}_{22}$. It is important to note that in general $\hat{\alpha}$ is not a consistent estimator of the unobserved α_1 nor does it converge to α_1. Despite these inferential shortcomings, adjustment of the forecast for y_{1,T_1^*+1} through the addition of $\hat{\alpha}$ has the potential to lower forecast risk under several conditions corresponding to different estimators of α_1.

2.3 Construction of shock effects estimators

We now construct the aggregate estimators of the shock effects that appear in Forecast 2. We use these to forecast response values $y_{1,t}$ when $t > T_1^*$, i.e., the time series of interest after the shock time where we assume that T_1^* is known. First, we introduce the procedures of parameter estimation for $\mathcal{M}_1$, $\mathcal{M}_{21}$, and $\mathcal{M}_{22}$ (see Section 2.1). Conditional on all regression parameters, previous responses, and covariates, the response variable $y_{i,t}$ in $\mathcal{M}_1$, $\mathcal{M}_{21}$, and $\mathcal{M}_{22}$ has distribution

$$y_{i,t} \sim N(\eta_i + \alpha_i D_{i,t} + \phi_i y_{i,t-1} + \theta_i' \mathbf{x}_{i,t} + \beta_i' \mathbf{x}_{i,t-1}, \sigma^2).$$

For $i = 2, \dots, n$, all parameters in this model will be estimated with ordinary least squares (OLS) using historical data of $t = 1, \dots, n_i$. For $i = 1$, we estimate all the parameters but α_1 using OLS procedures for $t = 1, \dots, T_1^*$. In particular, let $\hat{\alpha}_i$, $i = 2, \dots, n+1$ be the OLS estimate of α_i. Note that parameter estimation for $\mathcal{M}_1$ is identically the same as $\mathcal{M}_{21}$ and $\mathcal{M}_{22}$.

Second, we introduce the candidate estimators for α_1. Define the *adjustment estimator* for time series $i = 1$ by,

$$\hat{\alpha}_{\text{adj}} = \frac{1}{n} \sum_{i=2}^{n+1} \hat{\alpha}_i, \tag{4}$$

where the $\hat{\alpha}_i$s in (4) are OLS estimators of all of the α_is. We can use $\hat{\alpha}_{\text{adj}}$ as an estimator for the unknown α_1 term for which no meaningful estimation information otherwise exists. It is intuitive that $\hat{\alpha}_{\text{adj}}$ should perform well under $\mathcal{M}_1$ where we assume that α_i's share the same mean for $i = 1, \dots, n+1$. However, it can also be shown that $\hat{\alpha}_{\text{adj}}$ may be less favorable in $\mathcal{M}_{21}$ and $\mathcal{M}_{22}$, which will be discussed in detail in Section 3.

We also consider the *inverse-variance weighted estimator* in practical settings where the T_i's and T_i^*'s vary greatly across i. The inverse-variance weighted estimator is defined as

$$\hat{\alpha}_{\text{IVW}} = \frac{\sum_{i=2}^{n+1} \hat{\alpha}_i / \hat{\sigma}_{i\alpha}^2}{\sum_{i=2}^{n+1} 1/\hat{\sigma}_{i\alpha}^2}, \qquad \text{where} \qquad \hat{\sigma}_{i\alpha}^2 = \hat{\sigma}_i^2 (\mathbf{U}_i' \mathbf{U}_i)_{22}^{-1},$$

$\hat{\alpha}_i$ is the OLS estimator of α_i, $\hat{\sigma}_i$ is the residual standard error from OLS estimation, and $\mathbf{U}_i$ is the design matrix for OLS with respect to time series for $i = 2, \dots, n+1$. Note that since σ is unknown, estimation is required and the numerator and denominator terms are dependent in general. However, $\hat{\alpha}_{\text{IVW}}$ can be a reasonable estimator in practical settings. We do not provide closed form expressions for $\text{E}(\hat{\alpha}_{\text{IVW}})$ and $\text{Var}(\hat{\alpha}_{\text{IVW}})$, empirical performance of $\hat{\alpha}_{\text{IVW}}$ is assessed via Monte Carlo simulation (see Section 4).

We now motivate a *weighted-adjustment estimator* for model $\mathcal{M}_{21}$ and $\mathcal{M}_{22}$. Our weighted-adjustment estimator is inspired by the weighting techniques in synthetic control methodology (SCM) developed in Abadie et al. [2010]. However, our weighted-adjustment estimator is not a causal estimator and our estimation premise is a reversal of that in SCM. Our objective is in predicting a post-shock response y_{1,T_1^*+1} that is not yet observed using disparate time series whose post-shock responses are observed.

We use similar notation as that in Abadie et al. [2010] to motivate our weighted-adjustment estimator. Consider a $n \times 1$ weight vector $\mathbf{W} = (w_2, \ldots, w_{n+1})$, where $w_i \in [0, 1]$ for all $i = 2, \ldots, n+1$. Construct

$$\mathbf{X}_1 = \begin{pmatrix} \mathbf{x}_{1,T_1^*} \\ \mathbf{x}_{1,T_1^*+1} \end{pmatrix} \quad \text{and} \quad \hat{\mathbf{X}}_1(\mathbf{W}) = w_2 \begin{pmatrix} \mathbf{x}_{2,T_2^*} \\ \mathbf{x}_{2,T_2^*+1} \end{pmatrix} + \cdots + w_{n+1} \begin{pmatrix} \mathbf{x}_{n+1,T_{n+1}^*} \\ \mathbf{x}_{n+1,T_{n+1}^*+1} \end{pmatrix}.$$

where $\mathbf{X}_1$ and $\hat{\mathbf{X}}_1(\mathbf{W})$ are $2 \times p$. Define $\mathcal{W} = \{\mathbf{W} \in [0,1]^n \colon w_2 + \cdots + w_{n+1} = 1\}$. Suppose there exists $\mathbf{W}^* \in \mathcal{W}$ with $\mathbf{W}^* = (w_2^*, \ldots, w_{n+1}^*)$ such that

$$\mathbf{X}_1 = \hat{\mathbf{X}}_1(\mathbf{W}^*) \quad i.e., \quad \mathbf{x}_{1,T_1^*} = \sum_{i=2}^{n+1} w_i^* \mathbf{x}_{i,T_i^*} \text{ and } \mathbf{x}_{1,T_1^*+1} = \sum_{i=2}^{n+1} w_i^* \mathbf{x}_{i,T_i^*+1}. \tag{5}$$

Notice that $\mathbf{W}^*$ exists as long as $\mathbf{X}_1$ falls in the convex hull of

$$\left\{ \begin{pmatrix} \mathbf{x}_{2,T_2^*} \\ \mathbf{x}_{2,T_2^*+1} \end{pmatrix}, \ldots, \begin{pmatrix} \mathbf{x}_{n+1,T_{n+1}^*} \\ \mathbf{x}_{n+1,T_{n+1}^*+1} \end{pmatrix} \right\}.$$

Our weighted-adjustment estimator will therefore perform well when the pool of disparate time series posses similar covariates to the time series for which no post-shock responses are observed. We compute $\mathbf{W}^*$ as

$$\mathbf{W}^* = \underset{\mathbf{W} \in \mathcal{W}}{\arg\min} \left\| \mathrm{vec}\big(\mathbf{X}_1 - \hat{\mathbf{X}}_1(\mathbf{W})\big) \right\|_{2p}. \tag{6}$$

Abadie et al. [2010] commented that we can select $\mathbf{W}^*$ so that (5) holds approximately and that weighted-adjustment estimation techniques of this form are not appropriate when the fit is poor. Note that $\mathbf{W}^*$ is not random since the covariates are assumed to be fixed. Since $\mathcal{W}$ is a closed and bounded subset of $\mathbb{R}^n$, $\mathcal{W}$ is compact. Because the objective function is continuous in $\mathbf{W}$, $\mathbf{W}^*$ will always exist. Our weighted-adjustment estimator for the shock effect α_1 is

$$\hat{\alpha}_{\text{wadj}} = \sum_{i=2}^{n+1} w_i^* \hat{\alpha}_i \quad \text{for} \quad \mathbf{W}^* = \begin{pmatrix} w_2^* & \cdots & w_{n+1}^* \end{pmatrix}.$$

Estimation properties of $\hat{\alpha}_{\text{adj}}$, $\hat{\alpha}_{\text{IVW}}$, and $\hat{\alpha}_{\text{wadj}}$ are discussed in the remaining sections.

Remark 1. In Section 2.1 we specify that $\mathbf{x}_{i,t}, \theta, \beta \in \mathbb{R}^p$. However, it is not necessary that the all p covariates are important for every time series under study. The regression coefficients θ and β are nuisance parameters that are not of primary importance. It will be understood that structural 0s in $\mathbf{x}_{i,t}$ correspond to variables that are unimportant.

3 Forecast risk and properties of shock-effects estimators

In this section, we discuss the properties that are related to forecast-risk reduction. In discussion of risk, it is useful to derive expressions for expectation and variance of the adjustment estimator $\hat{\alpha}_{\text{adj}}$ and weighted-adjustment estimator. The expression for the expectations are as follow,

(i) Under $\mathcal{M}_1$, $\mathrm{E}(\hat{\alpha}_{\text{adj}}) = \mathrm{E}(\hat{\alpha}_{\text{wadj}}) = \mu_\alpha$.

(ii) Under $\mathcal{M}_{21}$,

$$\mathrm{E}(\hat{\alpha}_{\text{adj}}) = \mu_\alpha + \frac{1}{2}\sum_{i=2}^{n+1} \delta' \mathbf{x}_{i,T_i^*+1} + \frac{1}{n}\sum_{i=2}^{n+2} \gamma' \mathbf{x}_{i,T_i^*} \quad \text{and} \quad \mathrm{E}(\hat{\alpha}_{\text{wadj}}) = \mu_\alpha + \delta' \mathbf{x}_{1,T_1^*+1} + \gamma' \mathbf{x}_{1,T_1^*}.$$

(iii) Under $\mathcal{M}_{22}$,

$$\mathrm{E}(\hat{\alpha}_{\text{adj}}) = \mu_\alpha + \frac{1}{2}\sum_{i=2}^{n+1} \mu_\delta' \mathbf{x}_{i,T_i^*+1} + \frac{1}{n}\sum_{i=2}^{n+2} \mu_\gamma' \mathbf{x}_{i,T_i^*} \quad \text{and} \quad \mathrm{E}(\hat{\alpha}_{\text{wadj}}) = \mu_\alpha + \mu_\delta' \mathbf{x}_{1,T_1^*+1} + \mu_\gamma' \mathbf{x}_{1,T_1^*}.$$

Formal justification for these results can be found in Appendix. Note that $\hat{\alpha}_{\text{adj}}$, $\hat{\alpha}_{\text{wadj}}$, and $\hat{\alpha}_{\text{IVW}}$ are not unbiased estimators for α_1. Notice that under $\mathcal{M}_1$, $\hat{\alpha}_{\text{adj}}$ and $\hat{\alpha}_{\text{adj}}$ are unbiased estimators for $\mathrm{E}(\alpha_1) = \mu_\alpha$ (see distributional details of α_1 in Section 2.1). Nevertheless, $\hat{\alpha}_{\text{adj}}$ is a biased estimator for $\mathrm{E}(\alpha_1)$ but $\hat{\alpha}_{\text{wadj}}$ is an unbiased estimator for $\mathrm{E}(\alpha_1)$ under both $\mathcal{M}_{21}$ and $\mathcal{M}_{22}$. Thus, we collect these results as the following proposition.

Proposition 1.

(i) Under $\mathcal{M}_1$, $\hat{\alpha}_{\text{adj}}$ is an unbiased estimator of $\mathrm{E}(\alpha_1)$. Under $\mathcal{M}_{21}$ and $\mathcal{M}_{22}$, $\hat{\alpha}_{\text{adj}}$ is a biased estimator of $\mathrm{E}(\alpha_1)$ in general.

(ii) Suppose that $\mathbf{W}^*$ satisfies (5). Under $\mathcal{M}_1$, $\mathcal{M}_{21}$ and $\mathcal{M}_{22}$, $\hat{\alpha}_{\text{wadj}}$ is an unbiased estimator of $\mathrm{E}(\alpha_1)$.

Unbiasedness properties for $\mathrm{E}(\alpha_1)$ of $\hat{\alpha}_{\text{adj}}$ and $\hat{\alpha}_{\text{wadj}}$ allow for simple risk-reduction conditions and invoke a method of comparison, although our primary interest is in reducing forecast risk. These conditions will be discussed in Section 3.1 and Section 3.2. Next, we present the variance expressions for $\hat{\alpha}_{\text{adj}}$ and $\hat{\alpha}_{\text{wadj}}$ as below

(i) Under $\mathcal{M}_1$ and $\mathcal{M}_{21}$,

$$\mathrm{Var}(\hat{\alpha}_{\text{adj}}) = \frac{\sigma^2}{n^2}\sum_{i=2}^{n+1} \mathrm{E}\big\{(\mathbf{U}_i'\mathbf{U}_i)^{-1}_{22}\big\} + \frac{\sigma_\alpha^2}{n^2}$$

$$\mathrm{Var}(\hat{\alpha}_{\text{wadj}}) = \sigma^2 \sum_{i=2}^{n+1} (w_i^*)^2 \mathrm{E}\big\{(\mathbf{U}_i'\mathbf{U}_i)^{-1}_{22}\big\} + \sigma_\alpha^2 \sum_{i=2}^{n+1} (w_i^*)^2$$

(ii) Under $\mathcal{M}_{22}$,

$$\mathrm{Var}(\hat{\alpha}_{\text{adj}}) = \frac{\sigma^2}{n^2}\sum_{i=2}^{n+1} \mathrm{E}\big\{(\mathbf{U}_i'\mathbf{U}_i)^{-1}_{22}\big\} + \frac{1}{n^2}\mathrm{Var}(\alpha_i)$$

$$\mathrm{Var}(\hat{\alpha}_{\text{wadj}}) = \sigma^2 \sum_{i=2}^{n+1} (w_i^*)^2 \mathrm{E}\big\{(\mathbf{U}_i'\mathbf{U}_i)^{-1}_{22}\big\} + \sum_{i=2}^{n+1} (w_i^*)^2 \mathrm{Var}(\alpha_i).$$

Formal justification for these results can be found in Appendix. Note that the variances are not comparable in closed-form because of the term $\mathrm{E}\{(\mathbf{U}_i'\mathbf{U}_i)^{-1}_{22}\}$. This term exists because of the inclusion of the random lagged response in our auto regressive model formulation. Under $\mathcal{M}_{22}$, the expression for $\mathrm{Var}(\alpha_i)$ is not of closed form because γ_i and δ_i may be dependent when they are placed in a random-effects model. We investigate comparisons between the variability of these estimators in Section 3.2.

As Section 3.1 and 3.2 detailed the conditions for risk-reduction and comparisons, they usually involve fixed quantities related to variance and expectation. To make use of those properties in practice, estimation is required. Section 3.3 will introduce a general procedure of parametric bootstrap under the context of the problem to attain this purpose.

3.1 Conditions for risk-reduction for shock-effects estimators

In this section we will discuss the conditions for risk reduction for individual shock-effects estimators under $\mathcal{M}_1$, $\mathcal{M}_{21}$, and $\mathcal{M}_{22}$.

3.1.1 Conditions under $\mathcal{M}_1$

Recall that Proposition 1 implies that the adjustment estimator $\hat{\alpha}_{\mathrm{adj}}$ and weighted-adjustment estimator $\hat{\alpha}_{\mathrm{wadj}}$ are unbiased for $\mathrm{E}(\alpha_1)$ under $\mathcal{M}_1$. With this result, we will have the following propositions that specify the conditions that are necessary for risk reduction.

Proposition 2. *Under $\mathcal{M}_1$,*

(i) $R_{T_1^*+1,2} < R_{T_1^*+1,1}$ *when* $\mathrm{Var}(\hat{\alpha}_{\mathrm{adj}}) < \mu_\alpha^2$.

(ii) if $\mathbf{W}^*$ *satisfies (5),* $R_{T_1^*+1,2} < R_{T_1^*+1,1}$ *when* $\mathrm{Var}(\hat{\alpha}_{\mathrm{wadj}}) < \mu_\alpha^2$.

Proposition 2 tells that under $\mathcal{M}_1$ if the variance of the estimator is smaller than the squared mean of α_1, those estimators will enjoy the risk reduction properties. Recalling from variance expression at the beginning of Section 3, Proposition 2 shows that the risk-reduction condition is

$$\mathrm{Var}(\hat{\alpha}_{\mathrm{adj}}) = \frac{\sigma^2}{n^2}\sum_{i=2}^{n+1}\mathrm{E}\{(\mathbf{U}_i'\mathbf{U}_i)^{-1}_{22}\} + \frac{\sigma_\alpha^2}{n^2} < \mu_\alpha^2 \tag{7}$$

In terms of the adjustment estimator, $\hat{\alpha}_{\mathrm{adj}}$, (7) implies two facts: (1) Forecast 2 is preferable to Forecast 1 asymptotically in n whenever $\mu_\alpha \neq 0$; (2) In finite pool of time series, Forecast 2 is preferable to Forecast 1 when the μ_α is large relative to its variability and overall regression variability.

For the weighted-adjustment estimator $\hat{\alpha}_{\mathrm{wadj}}$, if $\mathbf{W}^*$ does not satisfy (5), its unbiased properties for $\mathrm{E}(\alpha_1)$ should hold approximately when the fit in (6) is appropriate as commented in Section 2.3. From Proposition 2 and variance expression of $\hat{\alpha}_{\mathrm{wadj}}$, the following is the risk-reduction condition for $\hat{\alpha}_{\mathrm{wadj}}$.

$$\mathrm{Var}(\hat{\alpha}_{\mathrm{wadj}}) = \sigma^2\sum_{i=2}^{n+1}(w_i^*)^2\mathrm{E}\{(\mathbf{U}_i'\mathbf{U}_i)^{-1}_{22}\} + \sigma_\alpha^2\sum_{i=2}^{n+1}(w_i^*)^2 < \mu_\alpha^2.$$

In this case, Forecast 2 is preferable to Forecast 1 when μ_α is large relative to the *weighted sum of variances for shock effects for other time series* and overall regression variability. However, the above criteria are generally difficult to evaluate in practice due to the term $\hat{\alpha}_{\mathrm{wadj}}$. Section 3.3 will provide a detailed treatment about how to deal with these technical inequalities in practice.

3.1.2 Conditions under $\mathcal{M}_{21}$ and $\mathcal{M}_{22}$

The α_is have different means under $\mathcal{M}_{21}$ and $\mathcal{M}_{22}$ unlike under $\mathcal{M}_1$. However, Proposition 1 implies that $\hat{\alpha}_{\text{wadj}}$ is an unbiased estimator of $\mathrm{E}(\alpha_1)$. We now state conditions for risk reduction.

Proposition 3. *If* $\mathbf{W}^*$ *satisfies (5), under* $\mathcal{M}_{21}$ *and* $\mathcal{M}_{22}$, $R_{T_1^*+1,2} < R_{T_1^*+1,1}$ *when* $\mathrm{Var}(\hat{\alpha}_{\text{wadj}}) < (\mathrm{E}(\alpha_1))^2$.

Based on Proposition 3, we can obtain a similar inequality as in Section 3.1.1 as below

$$\mathrm{Var}(\hat{\alpha}_{\text{wadj}}) = \sigma^2 \sum_{i=2}^{n+1} (w_i^*)^2 \mathrm{E}\big\{(\mathbf{U}_i'\mathbf{U}_i)_{22}^{-1}\big\} + \sum_{i=2}^{n+1} (w_i^*)^2 \mathrm{Var}(\alpha_i) < (\mathrm{E}(\alpha_1))^2,$$

where $\mathrm{Var}(\alpha_i)$ may be replaced with σ_α^2 in $\mathcal{M}_{21}$. The conclusions and intuitions will be identically the same as what we have in Section 3.1.1.

Proposition 1 shows that $\hat{\alpha}_{\text{adj}}$ is a biased estimator of $\mathrm{E}(\alpha_1)$ under $\mathcal{M}_{21}$ and $\mathcal{M}_{22}$ generally. Hence, Proposition 2 no longer holds for $\hat{\alpha}_{\text{adj}}$ under $\mathcal{M}_{21}$ and $\mathcal{M}_{22}$. But, as an alternative, we can derive similar conditions as below. By Lemma 1 (see Section 7.1) and risk decomposition, we will achieve risk-reduction as long as

$$\begin{aligned}
\mathrm{E}(\alpha_1^2) = \mathrm{Var}(\alpha_1) + (\mathrm{E}(\alpha_1))^2 &> \mathrm{E}(\hat{\alpha}_{\text{adj}} - \alpha_1)^2 \\
&= \mathrm{Var}(\hat{\alpha}_{\text{adj}}) + (\mathrm{E}(\hat{\alpha}_{\text{adj}}) - \alpha_1)^2 \\
&= \mathrm{Var}(\hat{\alpha}_{\text{adj}}) + \mathrm{Var}(\alpha_1) + (\mathrm{E}(\hat{\alpha}_{\text{adj}}) - \mathrm{E}(\alpha_1))^2
\end{aligned}$$

Therefore, the above inequality will simply to

$$(\mathrm{E}(\alpha_1))^2 > \mathrm{Var}(\hat{\alpha}_{\text{adj}}) + (\mathrm{E}(\hat{\alpha}_{\text{adj}}) - \mathrm{E}(\alpha_1))^2.$$

Note that since $\hat{\alpha}_{\text{adj}}$ is biased for $\mathrm{E}(\alpha_1)$, the bias term $(\mathrm{E}(\hat{\alpha}_{\text{adj}}) - \mathrm{E}(\alpha_1))^2$ will become complicated and simplification yields no insightful results.

As mentioned in Section 2.3, it is difficult to evaluate the expectation and variance of $\hat{\alpha}_{\text{IVW}}$. In other words, $\hat{\alpha}_{\text{IVW}}$ is generally biased for $\mathrm{E}(\alpha_1)$. That is to say we can adapt the above proof to derive the risk-reduction conditions for $\hat{\alpha}_{\text{IVW}}$: under $\mathcal{M}_1$, $\mathcal{M}_{21}$, and $\mathcal{M}_{22}$, $R_{T_1^*+1,2} < R_{T_1^*+1,1}$ when $\mathrm{Var}(\hat{\alpha}_{\text{IVW}}) + (\mathrm{E}(\hat{\alpha}_{\text{IVW}}) - \mathrm{E}(\alpha_1))^2 < (\mathrm{E}(\alpha_1))^2$.

Topics of evaluation of these inequalities in practice can be found in Section 3.3. We will discuss comparisons of adjustment estimators in the the next Section.

3.2 Comparisons among estimators

In comparing shock-effects estimators, we would assume that the risk-reduction conditions are satisfied as in Section 3.1.

Denote the risk-reduction quantity for the adjustment estimator as Δ_{adj}, the one for inverse-weighted estimator as Δ_{IVW}, and the one for weighted-adjustment estimator as Δ_{wadj}. As long as the risk-reduction of one estimator is greater than those of others, we will vote it as the best estimator among our pool of estimators for consideration. For example, if we find that $\Delta_{\text{wadj}} > \Delta_{\text{adj}}$ and $\Delta_{\text{wadj}} > \Delta_{\text{IVW}}$, the weighted-adjustment estimator $\hat{\alpha}_{\text{wadj}}$ is the most favorable.

According to discussion in Section 3.1.2, we know that under $\mathcal{M}_1$, $\mathcal{M}_{21}$, and $\mathcal{M}_{22}$, the risk-reduction quantity for $\hat{\alpha}_{\text{IVW}}$ is

$$\Delta_{\text{IVW}} = (\mathrm{E}(\alpha_1))^2 - \mathrm{Var}(\hat{\alpha}_{\text{IVW}}) - (\mathrm{E}(\hat{\alpha}_{\text{IVW}}) - \mathrm{E}(\alpha_1))^2.$$

From discussions in Section 3.1, we know that the risk-reduction quantities for $\hat{\alpha}_{\text{adj}}$ and $\hat{\alpha}_{\text{wadj}}$ differ across models, we will discuss in different cases accordingly.

3.2.1 Under $\mathcal{M}_1$

From Proposition 2, we know that the risk-reduction quantities for $\hat{\alpha}_{\text{adj}}$ and $\hat{\alpha}_{\text{wadj}}$ are

$$\Delta_{\text{adj}} = \mu_\alpha^2 - \text{Var}(\hat{\alpha}_{\text{adj}}) \quad \text{and} \quad \Delta_{\text{wadj}} = \mu_\alpha^2 - \text{Var}(\hat{\alpha}_{\text{wadj}}).$$

Under the framework of $\mathcal{M}_1$, the risk-reduction quantity for $\hat{\alpha}_{\text{IVW}}$ is

$$\Delta_{\text{IVW}} = \mu_\alpha^2 - \text{Var}(\hat{\alpha}_{\text{IVW}}) - (\text{E}(\hat{\alpha}_{\text{IVW}}) - \mu_\alpha)^2.$$

In other words, when $\text{Var}(\hat{\alpha}_{\text{wadj}}) < \text{Var}(\hat{\alpha}_{\text{adj}})$ and $\hat{\alpha}_{\text{wadj}} < \text{Var}(\hat{\alpha}_{\text{IVW}}) + (\text{E}(\hat{\alpha}_{\text{IVW}}) - \mu_\alpha)^2$, we would prefer $\hat{\alpha}_{\text{wadj}}$ as the best estimator. Other conditions for voting the other estimators as the best one follow similarly.

3.2.2 Under $\mathcal{M}_{21}$ and $\mathcal{M}_{22}$

According to Proposition 3 and the discussion in Section 3.1.2, the risk-reduction quantities $\hat{\alpha}_{\text{adj}}$ and $\hat{\alpha}_{\text{wadj}}$ are

$$\Delta_{\text{adj}} = (\text{E}(\alpha_1))^2 - \text{Var}(\hat{\alpha}_{\text{adj}}) - (\text{E}(\hat{\alpha}_{\text{adj}}) - \text{E}(\alpha_1))^2 \quad \text{and} \quad \Delta_{\text{wadj}} = (\text{E}(\alpha_1))^2 - \text{Var}(\hat{\alpha}_{\text{wadj}}).$$

In this case, the risk-reduction quantity for $\hat{\alpha}_{\text{adj}}$ is similar to that of $\hat{\alpha}_{\text{IVW}}$ since they are both biased for $\text{E}(\alpha_1)$. Thus,

$$\Delta_{\text{IVW}} = (\text{E}(\alpha_1))^2 - \text{Var}(\hat{\alpha}_{\text{IVW}}) - (\text{E}(\hat{\alpha}_{\text{IVW}}) - \text{E}(\alpha_1))^2$$

For the case of $\hat{\alpha}_{\text{adj}}$ and $\hat{\alpha}_{\text{wadj}}$, we can derive the following inequality for $\hat{\alpha}_{\text{wadj}}$ to be favored over $\hat{\alpha}_{\text{adj}}$.

$$\text{Var}(\hat{\alpha}_{\text{adj}}) - \text{Var}(\hat{\alpha}_{\text{wadj}}) + \big(\text{E}(\hat{\alpha}_{\text{adj}}) - \text{E}(\alpha_1)\big)^2 > 0.$$

We analyze this inequality from two perspectives.

1. If it turns out to be fact that the variance of the weighted-adjustment estimator is greater than that of adjustment estimator, we should be aware that the compromise for variance because of using $\hat{\alpha}_{\text{wadj}}$ shouldn't exceed the squared bias, i.e., $\big(\text{E}(\hat{\alpha}_{\text{adj}}) - \text{E}(\alpha_1)\big)^2$.

2. If instead the variance of $\hat{\alpha}_{\text{wadj}}$ is smaller than that of $\hat{\alpha}_{\text{adj}}$, the above inequality should always hold because $\big(\text{E}(\hat{\alpha}_{\text{adj}}) - \text{E}(\alpha_1)\big)^2 > 0$ under $\mathcal{M}_{21}$ and $\mathcal{M}_{22}$.

These are some analytical results for comparison studies among estimators of α_1. Next, we will detail a framework for estimation of risk-reduction quantities using a parametric bootstrap routine. Therefore, the above inequalities can be analyzed numerically in practice.

3.3 Bootstrap for risk-reduction evaluation problems

In this section, we present a bootstrap procedure that approximates the distribution of our shock-effect estimators and check the underlying conditions of our risk reduction results in practice. Our procedure involves the resampling of residuals in the separate OLS fits. This procedure has its origins in Section 6 of Efron and Tibshirani [1986], and it involves the resampling the residuals which are assumed to be the realizations of an iid processes. Bose [1988] showed that the asymptotic accuracy for OLS parameter estimation can be further improved from $O(T^{-1/2})$ to $o(T^{-1/2})$

almost surely under some regularity conditions. However, the pseudo time series generated by this procedure are not stationary. (**Is this necessary to state? There needs to be a rationale fro why this is not a problem.**)

I think that our bootstrap procedure should first sample the disparate time series without replacement, and then estimate the parameters in each resampled time series. The formal steps of our bootstrap procedures are outlined in the Supplementary Materials, the intuition for our procedure is as follows: let B be the bootstrap sample size and initialize $y_{i,0}$ for all $i = 2, \ldots, n+1$. At iteration b, resample the residuals and then obtain shock-effect estimators for each of the disparate time series for all $i = 2, \ldots, n+1$. Then construct and store any of the adjustment estimators $\hat{\alpha}^{(b)}_{\text{adj}}$, $\hat{\alpha}^{(b)}_{\text{wadj}}$, and $\hat{\alpha}^{(b)}_{\text{IVW}}$. We can then estimate distributional quantities of our shock-effect estimators with the bootstrap sample of $\hat{\alpha}^{(b)}_{\text{adj}}$, $\hat{\alpha}^{(b)}_{\text{wadj}}$, and $\hat{\alpha}^{(b)}_{\text{IVW}}$, for $b = 1, \ldots, B$. Once can then use the bootstrapped shock-effects estimated by our residual bootstrap to provide an approximation for parameters involved in the risk-reduction conditions in Propositions 2 and 3.

Let $\overline{\hat{\alpha}_{\text{aadj}}} = \frac{1}{B}\sum_{b=1}^{B}\sum_{i=2}^{n+1}\hat{\alpha}^{(b)}_i$, $\overline{\hat{\alpha}_{\text{wadj}}} = \frac{1}{B}\sum_{b=1}^{B}\sum_{i=2}^{n+1}w^*_i\hat{\alpha}^{(b)}_i$, and $\overline{\hat{\alpha}_{\text{IVW}}} = \frac{1}{B}\sum_{b=1}^{B}\sum_{i=2}^{n+1}\hat{\alpha}^{(b)}_{\text{IVW},i}$. The almost-sure convergence results for $\hat{\alpha}_{\text{adj}}$ and $\hat{\alpha}_{\text{wadj}}$ follow naturally, provided that the regularity conditions outlined in Bose [1988] hold. The proof for the case of $\overline{\hat{\alpha}_{\text{wadj}}}$, the sample mean of the bootstrapped weighted adjustment estimator, can be presented as below. As $B \to \infty$,

$$\overline{\hat{\alpha}_{\text{wadj}}} = \frac{1}{B}\sum_{b=1}^{B}\sum_{i=2}^{n+1}w^*_i\hat{\alpha}^{(b)}_i = \sum_{i=2}^{n+1}w^*_i\frac{1}{B}\sum_{b=1}^{B}\hat{\alpha}^{(b)}_i \xrightarrow{a.s.} \sum_{i=2}^{n+1}w^*_i\alpha_i = \mathrm{E}(\hat{\alpha}_{\text{wadj}}|\Theta),$$

The case for the sample variance of the bootstrapped weighted adjustment estimator is similar. The same rationale holds for the adjustment estimator as well since it is a linear combination of OLS estimates, and the weights are not random conditioned on Θ. However, the case for $\hat{\alpha}_{\text{IVW}}$ is slightly different because it is a linear combination of OLS estimates with random weights. It is not clear with respect to whether similar consistency results holds for $\hat{\alpha}_{\text{IVW}}$. We claim that our bootstrap procedure provides an approximation for the case of $\hat{\alpha}_{\text{IVW}}$.

We stress that the above approximation is conditioned on Θ and that bootstrapping cannot alleviate the inherent bias of using our adjustment estimators as estimates for α_1. Simulation for justification of the parametric bootstrap is provided in Section 4.1.

4 Simulation

4.1 Parametric bootstrap simulation

5 Forecasting Conoco Phillips stock in the presence of shocks

In this example we forecast Conoco Phillips stock prices in the midst of the coronavirus recession. Specific interest is in predictions made after March 6th, 2020, the Friday before the stock market crash on March 9th, 2020. We will detail how we combine knowledge from disparate time series to improve the forecast of Conoco Phillips stock price that would be made without adjustment for the shock. The forecast consists of four steps: (1) pick a model, (2) selection of covariates, (3) choices of donor pool, and (4) forecast.

Conoco Phillips is chosen for this analysis because it is a large oil and gas resources company [ConocoPhillips, 2020]. Focus on the oil sector is because oil prices have been shown to exhibit a cointegrating behavior with economic indices [He et al., 2010], and our chosen time frame represents the onset of a significant economic down turn, coupled with a Russia and OPEC battle for global oil price control the Sunday before trading resumes on Monday, March 9th [Sukhankin, 2020]. Furthermore, fear of and action in response to the coronavirus pandemic began to uptick dramatically between Friday, March 6th and Monday, March 9th. Major events include the SXSW festival being cancelled as trading closed on March 6th [Wang et al., 2020]. New York declared a state of emergency on March 7th [New York State Government, 2020], and by Sunday, March 8th, eight states have declared a state of emergency [Alonso, 2020] while Italy placed 16 million people in quarantine [Sjödin et al., 2020].

Economic indicators forecasted our recession before the coronavirus pandemic began. The current recession followed an inversion of the yield curve that first happened back in March, 2019 [Tokic, 2019]. An inversion of the yield curve is an event that signals that recessions are more likely [Andolfatto and Spewak, 2018, Bauer and Mertens, 2018]. In this analysis we investigate the performance of oil companies in previous recessions that followed an inversion of the yield curve to obtain a suitable Conoco Phillips donor pool for estimating the March 9th shock effect on Conoco Phillips oil stock. We also consider previous OPEC oil supply shocks [Mensi et al., 2014]. We will borrow from the literature on oil price forecasting to establish appropriate time horizons and forecasting models. Recessions that occurred before 1973 are disregarded since oil price forecasts cannot be represented by standard time series models before 1973 [Alquist et al., 2013]. In this analysis we make the following considerations:

(1) **AR(1) model and time window**. We will use a simple AR(1) model to forecast Conoco Phillips stock price. This model has been shown to beat no-change forecasts when predicting oil prices over time horizons of 1 and 3 months [Alquist et al., 2013]. We will consider 30 pre-shock trading days and we will forecast the immediate shock effect and the shock effect over a future five trading day window. All estimates will be adjusted for inflation. The model setup for AR(1) is exactly the same as what is stated in Section 2.1 with addition of shock effects. All the parameters are estimated using OLS.

(2) **Selection of covariates**. We perform our analyses incorporating daily S&P 500 index prices and West Texas Intermediate (WTI) crude oil prices as covariates.

(3) **Construction of donor pool**. Our donor pool will consist of Conoco Phillips shock effects observed after September 11th, 2001, several events in September, 2008, and November 27, 2014. The first two shock effects were observed during recessions that were predicated by an inversion of the yield curve [Bauer and Mertens, 2018], and the third was an OPEC induced supply side shock effect [Huppmann and Holz, 2015]. The reasons for those three shocks are:

    (a) On September 11th, 2001, Islamic extremist al-Qaeda committed a series of terrorist attacks against the United States of America [Braniff and Moghadam, 2011]. In addition to the tragic loss of life, this triggered negative repercussions on the U.S. and world economies with falling stock prices and a falling U.S. dollar [Floyd and Fuerbringer, 2001]. Moreover, trading was closed until September 17, 2001 [Johnston and Nedelescu, 2006].

    (b) In early September 2008, time series of oil prices experienced a sudden increase in volatility simultaneously due to turmoil in financial markets. The political, economic, social or environmental events may coincide with these shocks [Ewing and Malik, 2013]. Notable shock effects followed the placement of Fannie May and Freddie Mac in conservatorship on September 7th (shock effect on the 8th), Lehman Brothers filing for bankruptcy on September 15th, and the Office of Thrift Supervision closes Washington Mutual Bank on September 25th [Dwyer and Tkac, 2009, Longstaff, 2010].

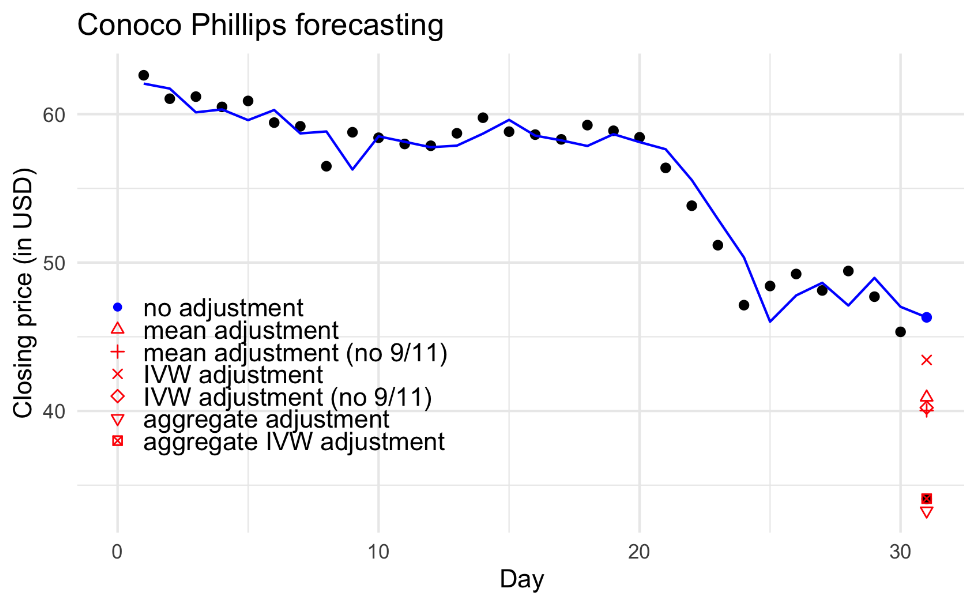

Figure 2. March 9th, 2020 post-shock forecasts for Conoco Phillips stock price. The blue line displays the

(c) On November 27th, 2014, it is documented that oil prices fall as OPEC opts not to cut production [Huppmann and Holz, 2015]. During the Great Recession when economic activity clearly declined, both oil and stock prices fell which points to demand factors. During the second half of 2014, oil prices plummeted but equity prices generally increased, suggesting that supply factors were the key driver [Baffes et al., 2015, Page 19].

(4) **Forecast. Jilei, please fill this in. I do not know what you have in mind.**

Jilei, can you add the weighted adjustment estimator to this analysis? We estimate the five shock effects mentioned above using OLS and then construct $\hat{\alpha}_{\text{adj}}$ and $\hat{\alpha}_{\text{IVW}}$ from these estimates. We also consider estimation without the September 11th, 2001 shock effect since there was a week long delay in trading due to the terrorist attack. We also study additive shock effect estimators where the shock effects corresponding to separate supply and demand shocks are added to estimate the unknown shock effect. The supply shock donor pool consists of the November 27th, 2014 shock effect and the demand shock donor pool consists of the remaining shock effects.

We can see from Figure 2 that $\hat{\alpha}_{\text{adj}}$ and $\hat{\alpha}_{\text{IVW}}$ perform decently well, they do not recover the magnitude of the shock effect but are much better than unadjusted forecasts that do not account for shock effects. The unadjusted forecast had an RMSE of 12.24, the use of $\hat{\alpha}_{\text{adj}}$ had an RMSE of 6.85, the use of $\hat{\alpha}_{\text{adj}}$ without September 11th had an RMSE of 5.96, the use of $\hat{\alpha}_{\text{IVW}}$ had an RMSE of 9.37, and the use of $\hat{\alpha}_{\text{IVW}}$ without September 11th had an RMSE of 6.15. The additive adjustment estimator computed by adding the $\hat{\alpha}_{\text{adj}}$ estimators for the demand and supply shock effects without September 11th has an RMSE of 0.80. The additive adjustment estimator computed by adding the $\hat{\alpha}_{\text{IVW}}$ estimators for the demand and supply shock effects without September 11th has an RMSE of 0.023.

We also see that additive estimators do extremely well. However, these additive estimators are retrospective but their is some apriori justification for their use. Kilian [2009] study the effect that different supply and demand shocks have on oil prices through a vector auto regressive model. Their model postulates an additive nature of shock effects, although the additivity parameters requires estimation in their context.

6 Discussion

Our bootstrap procedure can be extended to approximate the distribution of shock effect estimators from more general time series.

7 Appendix

7.1 Proofs

7.1.1 Justification of Expectation of $\hat{\alpha}_{\mathrm{adj}}$ and $\hat{\alpha}_{\mathrm{wadj}}$

The building block for the following proof is the fact that least squares is conditionally unbiased conditioned on Θ.

Case I: under $\mathcal{M}_1$: It follows that under $\mathcal{M}_1$ (see Section 2.1),

$$\mathrm{E}(\hat{\alpha}_{\mathrm{adj}}) = \frac{1}{n}\sum_{i=2}^{n+1}\mathrm{E}(\mathrm{E}(\hat{\alpha}_i|\Theta)) = \mu_\alpha \quad \text{and} \quad \mathrm{E}(\hat{\alpha}_{\mathrm{wadj}}) = \sum_{i=2}^{n+1} w_i^*\mathrm{E}(\mathrm{E}(\hat{\alpha}_i|\Theta)) = \sum_{i=2}^{n+1} w_i^*\mu_\alpha = \mu_\alpha.$$

where we used the fact that $\sum_{i=2}^{n+1} w_i = 1$.

Case II: under $\mathcal{M}_{21}$ and $\mathcal{M}_{22}$: Since $\mathrm{E}(\tilde{\varepsilon}_{i,T_i}) = 0$, $\mathrm{E}(\hat{\tilde{\alpha}}_i) = \mathrm{E}(\tilde{\alpha}_i) = \mathrm{E}(\alpha_i)$, it follows that

$$\begin{aligned}
\mathrm{E}(\hat{\alpha}_{\mathrm{wadj}}) = \mathrm{E}\left\{\mathrm{E}\left(\sum_{i=2}^{n+1} w_i^*\hat{\alpha}_i|\Theta\right)\right\} &= \mathrm{E}\left(\sum_{i=2}^{n+1} w_i^*\alpha_i\right)\\
&= \mathrm{E}\Bigg\{\sum_{i=2}^{n+1} w_i^*\left[\mu_\alpha + \delta_i'\mathbf{x}_{i,T_i^*+1} + \gamma_i'\mathbf{x}_{i,T_i^*}\right]\Bigg\}\\
&= \mu_\alpha + \mathrm{E}\Bigg\{\sum_{i=2}^{n+1} w_i^*\left[\delta_i'\mathbf{x}_{i,T_i^*+1} + \gamma_i'\mathbf{x}_{i,T_i^*}\right]\Bigg\}. \qquad (\mathbf{W} \in \mathcal{W})
\end{aligned}$$

Similarly,

$$\mathrm{E}(\hat{\alpha}_{\mathrm{adj}}) = \mu_\alpha + \frac{1}{n}\sum_{i=2}^{n+1}\mathrm{E}(\delta_i'\mathbf{x}_{i,T_i^*+1} + \gamma_i'\mathbf{x}_{i,T_i^*}).$$

7.1.2 Justification of Variance of $\hat{\alpha}_{\mathrm{adj}}$ and $\hat{\alpha}_{\mathrm{wadj}}$

Notice that under the setting of OLS, the design matrix for $\mathcal{M}_2$ is the same as the one for $\mathcal{M}_1$. Therefore, it follows that

$$\begin{aligned}
\mathrm{Var}(\hat{\alpha}_{\mathrm{wadj}}) &= \mathrm{E}(\mathrm{Var}(\hat{\alpha}_{\mathrm{wadj}}|\Theta)) + \mathrm{Var}(\mathrm{E}(\hat{\alpha}_{\mathrm{wadj}}|\Theta))\\
&= \mathrm{E}\left\{\mathrm{Var}\left(\sum_{i=2}^{n+1} w_i^*\hat{\alpha}_i|\Theta\right)\right\} + \mathrm{Var}\left(\sum_{i=2}^{n+1} w_i^*\alpha_i\right)
\end{aligned}$$

Under $\mathcal{M}_{21}$ where $\delta_i = \delta$ and $\gamma_i = \gamma$ are fixed unknown parameters, we will have

$$\begin{aligned}
\mathrm{Var}(\hat{\alpha}_{\mathrm{wadj}}) &= \mathrm{E}\left\{\sum_{i=2}^{n+1}(w_i^*)^2(\sigma^2(\mathbf{U}_i'\mathbf{U}_i)^{-1}_{22})\right\} + \sigma_\alpha^2\sum_{i=2}^{n+1}(w_i^*)^2 \\
&= \sigma^2\sum_{i=2}^{n+1}(w_i^*)^2\mathrm{E}\big\{(\mathbf{U}_i'\mathbf{U}_i)^{-1}_{22}\big\} + \sigma_\alpha^2\sum_{i=2}^{n+1}(w_i^*)^2. \qquad (8)
\end{aligned}$$

Similarly, under $\mathcal{M}_{22}$ where we assume $\delta_i \perp\!\!\!\perp \gamma_i \perp\!\!\!\perp \varepsilon_{i,t}$, we have

$$\mathrm{Var}(\hat{\alpha}_{\mathrm{wadj}}) = \sigma^2\sum_{i=2}^{n+1}(w_i^*)^2\mathrm{E}\big\{(\mathbf{U}_i'\mathbf{U}_i)^{-1}_{22}\big\} + \sum_{i=2}^{n+1}(w_i^*)^2\mathrm{Var}(\alpha_i)$$

For the adjustment estimator, we simply replace $\mathbf{W}^*$ with $1/n\mathbf{1}_n$. Thus, under $\mathcal{M}_{21}$ we have

$$\mathrm{Var}(\hat{\alpha}_{\mathrm{adj}}) = \frac{\sigma^2}{n^2}\sum_{i=2}^{n+1}\mathrm{E}\big\{(\mathbf{U}_i'\mathbf{U}_i)^{-1}_{22}\big\} + \frac{\sigma_\alpha^2}{n^2}$$

Under $\mathcal{M}_{22}$, we shall have

$$\mathrm{Var}(\hat{\alpha}_{\mathrm{adj}}) = \frac{\sigma^2}{n^2}\sum_{i=2}^{n+1}\mathrm{E}\big\{(\mathbf{U}_i'\mathbf{U}_i)^{-1}_{22}\big\} + \frac{1}{n^2}\mathrm{Var}(\alpha_i).$$

Notice that $\mathcal{M}_1$ differs from $\mathcal{M}_{21}$ only by its mean parameterization of α (see Section 2.1). In other words, the variances of $\hat{\alpha}_{\mathrm{adj}}$ and $\hat{\alpha}_{\mathrm{wadj}}$ under $\mathcal{M}_1$ are the same for those under $\mathcal{M}_{21}$.

7.2 Proofs for lemmas and propositions

Proof of Proposition 1 The proof for unbiasedness follows immediately from discussions related to expectation in Section 3. For the biasedness of $\hat{\alpha}_{\mathrm{adj}}$ under $\mathcal{M}_{21}$ and $\mathcal{M}_{22}$, we write the bias term for $\hat{\alpha}_{\mathrm{adj}}$ as below.

$$\mathrm{Bias}(\hat{\alpha}_{\mathrm{adj}}) = \begin{cases} \frac{1}{n}\sum_{i=2}^{n+1}\delta'(\mathbf{x}_{i,T_i^*+1} - n\mathbf{x}_{1,T_1^*+1}) + \frac{1}{n}\sum_{i=2}^{n+1}\gamma'(\mathbf{x}_{i,T_i^*} - n\mathbf{x}_{1,T_1^*}) & \text{for } \mathcal{M}_{21} \\ \frac{1}{n}\sum_{i=2}^{n+1}\mu_\delta'(\mathbf{x}_{i,T_i^*+1} - n\mathbf{x}_{1,T_1^*+1}) + \frac{1}{n}\sum_{i=2}^{n+1}\mu_\gamma'(\mathbf{x}_{i,T_i^*} - n\mathbf{x}_{1,T_1^*}) & \text{for } \mathcal{M}_{22} \end{cases}.$$

But it may be unbiased in some special circumstances when the above bias turns out to be 0. $\square$

Lemma 1. *The forecast risk difference is* $R_{T_1^*+1,1} - R_{T_1^*+1,2} = \mathrm{E}(\alpha_1^2) - \mathrm{E}(\hat{\alpha}-\alpha_1)^2$ *for all estimators of* α_1 *that are independent of* Θ_1 *(see Section 2.1).*

Proof of Lemma 1 Define

$$C(\Theta_1) = \hat{\eta}_1 + \hat{\phi}_1 y_{1,T_1^*} + \hat{\theta}_1'\mathbf{x}_{1,T_1^*+1} + \hat{\beta}_1'\mathbf{x}_{1,T_1^*} - (\eta_1 + \phi_1 y_{1,T_1^*} + \theta_1'\mathbf{x}_{1,T_1^*+1} + \beta_1'\mathbf{x}_{1,T_1^*}),$$

where Θ_1 is as defined in (3). Notice that

$$R_{T_1^*+1,1} = \mathrm{E}\big\{\big(C(\Theta_1) - \alpha_1\big)^2\big\} \qquad \text{and} \qquad R_{T_1^*+1,2} = \mathrm{E}\big\{\big(C(\Theta_1) + \hat{\alpha} - \alpha_1\big)^2\big\}.$$

It follows that

$$R_{T_1^*+1,1} - R_{T_1^*+1,2} = \mathrm{E}(\alpha_1^2) - 2\mathrm{E}(C(\Theta_1)\hat{\alpha}) - \mathrm{E}(\hat{\alpha} - \alpha_1)^2.$$

Assuming $\mathbf{S} = (\mathbf{1}_n, \mathbf{y}_{1,t-1}, \mathbf{x}_1, \mathbf{x}_{1,t-1})$ has full rank, under OLS setting, $\hat{\eta}_1$, $\hat{\phi}_1$, $\hat{\theta}_1$, and $\hat{\beta}_1$ are unbiased estimators of η_1, ϕ_1, θ_1, and β_1, respectively under conditioning of Θ_1. Since we assume $\hat{\alpha}$ is independent of Θ_1, through the method of iterated expectation,

$$\mathrm{E}(C(\Theta_1)\hat{\alpha}) = \mathrm{E}\big\{\hat{\alpha} \cdot \mathrm{E}(C(\Theta_1) \mid \Theta_1)\big\} = 0.$$

It follows that

$$R_{T_1^*+1,1} - R_{T_1^*+1,2} = \mathrm{E}(\alpha_1^2) - \mathrm{E}(\hat{\alpha} - \alpha_1)^2,$$

which finishes the proof. □

Proof of Proposition 2 The proofs are arranged into two separate parts as below.

Proof for statement (i): Under $\mathcal{M}_1$, $\hat{\alpha}_{\mathrm{adj}}$ is an unbiased estimator of $\mathrm{E}(\alpha_1)$ because

$$\begin{aligned} E\left(\frac{1}{n}\sum_{i=2}^{n+1}\hat{\alpha}_i\right) = \frac{1}{n}\sum_{i=2}^{n+1}\mathrm{E}(\hat{\alpha}_i) &= \frac{1}{n}\sum_{i=2}^{n+1}\mathrm{E}(\mathrm{E}(\hat{\alpha}_i \mid \Theta)) \\ &= \frac{1}{n}\sum_{i=2}^{n+1}\mathrm{E}(\alpha_i) = \mu_\alpha = \mathrm{E}(\alpha_1), \end{aligned}$$

where we used the fact that OLS estimator is unbiased when the design matrix $\mathbf{U}_i$ is of full rank for all $i = 2, \ldots, n+1$. Because $\alpha_1 \perp\!\!\!\perp \varepsilon_{i,t}$, $\mathrm{E}(\hat{\alpha}_{\mathrm{adj}}\alpha_1) = \mathrm{E}(\hat{\alpha}_{\mathrm{adj}})\mathrm{E}(\alpha_1) = (\mathrm{E}(\hat{\alpha}_{\mathrm{adj}}))^2$. By Lemma 1,

$$\begin{aligned} R_{T_1^*+1,1} - R_{T_1^*+1,2} &= \mathrm{E}(\alpha_1^2) - \mathrm{E}(\hat{\alpha}_{\mathrm{adj}} - \alpha_1)^2 \\ &= \mathrm{E}(\alpha_1^2) - \mathrm{E}(\alpha_1^2) - \mathrm{E}(\hat{\alpha}_{\mathrm{adj}}^2) + 2\mathrm{E}(\hat{\alpha}_{\mathrm{adj}}\alpha_1) \\ &= \mu_\alpha^2 - \mathrm{Var}(\hat{\alpha}_{\mathrm{adj}}) \end{aligned}$$

Therefore, as long as we have $\mathrm{Var}(\hat{\alpha}_{\mathrm{adj}}) < \mu_\alpha^2$, we will achieve the risk reduction.

Proof for statement (ii): By Proposition 1, the property that $\hat{\alpha}_{\mathrm{wadj}}$ is an unbiased estimator of μ_α holds for $\mathcal{M}_1$. The remainder of the proof follows a similar argument to the proof of statement (i). □

Proof of Proposition 3 By Proposition 1, the property that $\hat{\alpha}_{\mathrm{wadj}}$ is an unbiased estimator of $\mathrm{E}(\alpha_1)$ holds for $\mathcal{M}_{21}$ and $\mathcal{M}_{22}$. The remainder of the proof follows a similar argument to the proof of Proposition 2. □

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