

# Identification Robust Inference for Risk Prices in Structural Stochastic Volatility Models

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## Abstract

In structural stochastic volatility asset pricing models, changes in volatility affect risk premia through two channels: (1) the investor’s willingness to bear high volatility in order to get high expected returns as measured by the market risk price, and (2) the investor’s direct aversion to changes in future volatility as measured by the volatility risk price. Disentangling these channels is difficult and poses a subtle identification problem that invalidates standard inference. We adopt the discrete-time exponentially affine model of Han, Khrapov, and Renault (2018), which links the identification of volatility risk price to the leverage effect. In particular, we develop a minimum distance criterion that links the market risk price, the volatility risk price, and the leverage effect to the well-behaved reduced-form parameters governing the return and volatility’s joint distribution. The link functions are almost flat if the leverage effect is close to zero, making estimating the volatility risk price difficult. We apply the conditional quasi-likelihood ratio test Andrews and Mikusheva (2016) develop in a nonlinear GMM framework to a minimum distance framework. The resulting conditional quasi-likelihood ratio test is uniformly valid. We invert this test to derive robust confidence sets that provide correct coverage for the risk prices regardless of the leverage effect’s magnitude.

**JEL Codes:** C12, C14, C38, C58, G12

**Keywords:** weak identification, robust inference, stochastic volatility, leverage, market risk premium, volatility risk premium, risk price, confidence set, asymptotic size

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# 1 Introduction

A fundamental question in finance is how investors optimally trade off risk and return. Economic theories predict investors demand a higher return as compensation for bearing more risk. Hence, we should expect a positive relationship between the mean and volatility of returns. Some seminal early papers proposed a static trade-off between risk and expected return, most notably the capital asset pricing model (CAPM) of Sharpe (1964) and Lintner (1965). In practice, volatility varies over time. Consequently, a significant strand of the recent literature examines the dynamic tradeoff between volatility and returns, including structural stochastic volatility models (Bansal et al. (2014) and Dew-Becker et al. (2017)). In these models, investors care not just about how an asset's returns co-move with the volatility but also care how they co-move with changes in volatility.

In structural stochastic volatility models, changes in volatility affect risk premia through two channels: (1) the investor's willingness to tolerate high volatility in order to get high expected returns as measured by the market risk price, and (2) the investor's direct aversion to changes in future volatility as measured by the volatility risk price. We adopt the discrete-time exponentially affine model of Han, Khrapov, and Renault (2018), who represent the market risk price and the volatility risk price by two structural parameters. In this model, Han, Khrapov, and Renault (2018) establish the significant result that the identification of the volatility risk price depends on a substantial leverage effect, which is the negative correlation between return and change in volatility.

Although the leverage effect is theoretically less than zero, it is difficult to quantify empirically, and its estimate usually is small (Aït-Sahalia, Fan, and Li 2013). When the leverage effect is mild, the data only provide a limited amount of information about the volatility risk, compared to the finite-sample noise in the data. This low signal-to-noise ratio, modeled by weak identification, invalidates standard inference based on the generalized method of moments (GMM) estimator, see Stock and Wright (2000) and Andrews and Cheng (2012).

We provide identification-robust confidence set (CS) for the structural parameters that measure the market risk price, the volatility risk price, and the leverage effect. The robust CS provides correct asymptotic coverage, uniformly over a large set of models that allow for any amount of leverage effect. This uniform validity is crucial for the CS to have good finite-sample coverage (Mikusheva 2007; Andrews and Guggenberger 2010). In contrast, standard confidence sets based on the GMM estimator and its asymptotic normality do not have uniform validity in the presence of small leverage effect. This issue affects all of the structural parameters because they are estimated simultaneously.

We achieve robust inference in two steps. First, we establish a minimum distance criterion using link functions between the structural parameters and a set of reduced-form parameters that determine the joint distribution of the return and volatility. The structural model implies that the link functions are zero when evaluated at the true values of the structural parameters and the reduced-form parameters. Identification and estimation of these reduced form parameters are standard and not affected by small leverage effect. However, the link function is almost flat in the structural parameters when the leverage effect is small, resulting in weak identification. Second,

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given this minimum distance criterion, we invert the conditional quasi-likelihood ratio (QLR) test by Andrews and Mikusheva (2016) to construct a robust confidence set. The key feature of this test is that it treats the flat link function as an infinite dimensional nuisance parameter. The critical value is constructed by conditioning on a sufficient statistic for this nuisance parameter, and it is shown to yield a valid test regardless of the nuisance parameter. Andrews and Mikusheva (2016) develop this test in a GMM framework. We show it works in the minimum distance context here and provide conditions for its asymptotic validity. For practitioners, we provide a detailed algorithm for the construction of this simulation-based robust confidence set.

Our paper relates to the empirical analysis of the effect of volatility on risk premia. As Lettau and Ludvigson (2010) mention, the evidence here is inconclusive. Bollerslev, Engle, and Wooldridge (1988), Harvey (1989), Ghysels, Santa-Clara, and Valkanov (2005), Bali and Peng (2006), and Ludvigson and Ng (2007) find a positive relationship, while Campbell (1987), Breen, Glosten, and Jagannathan (1989), Pagan and Hong (1991), Whitelaw (1994), and Brandt and Kang (2004) find a negative relationship. Also, some papers use both a market risk factor and a variance risk factor to explain the risk premia dynamics, including Christoffersen, Heston, and Jacobs (2013), Feunou et al. (2014), and Dew-Becker et al. (2017). A substantial positive variance risk premium is documented by Bollerslev, Law, and Tauchen (2008) and Drechsler and Yaron (2011). We contribute to this literature by providing a new way to make inference for the market risk price and the volatility risk price. This new CS not only allows for both effects but also takes into account the potential identification issue.

The weak identification issue studied in this paper is relevant in many economic applications, ranging from linear instrumental variable models (Staiger and Stock 1997) to nonlinear structural models (Mavroeidis, Plagborg-Møller, and Stock 2014; Andrews and Mikusheva 2015). This paper is the first one to study this issue in structural asset pricing models with stochastic volatility. Moreira (2003) introduced the conditional inference approach to the linear instrumental variable model, and Kleibergen (2005) applied it to the nonlinear GMM problem. Andrews and Mikusheva (2016) propose conditional inference for nonlinear GMM problems with infinitely dimensional nuisance parameter. This paper applies it to a minimum distance criterion and extends the scope of its application to a new type of asset pricing model.

The rest of the paper is organized as follows. Section 2 provides the model and its parameterization. Section 3 provides model-implied restrictions and use them to derive the link function. Section 4 provides the asymptotic distribution of the reduced-form parameter and robust inference for the structural parameter. A detailed algorithm to construct the robust confidence set is given in Section 4.3. Proofs are given in the appendix.

## 2 The Model

This section provides a parametric structural model with stochastic volatility, following Han, Khrapov, and Renault (2018). They extend the discrete-time exponentially-affine model of Darolles,

looks like double space.

Gouriéroux, and Jasiak (2006), and their model is a natural discrete-time analog of the Heston (1993) model. We specify this model using a stochastic discount factor (SDF), also called the pricing kernel, and the physical measure, which gives the joint distribution of the return and volatility dynamics.<sup>1</sup> We first define the SDF and parameterize it as an exponential affine function with unknown parameters. Then we provide parametric distribution for the physical measure.

Let  $P_t$  be the price of the asset. Let  $r_{t+1} = \log(P_{t+1}/P_t) - r_f$  denote the log excess return minus the risk-free rate and  $\sigma_{t+1}^2$  denote its volatility. The observed data is  $W_t = (r_t, \sigma_t^2)$  for  $t = 1, \dots, T$ . Let  $\mathcal{F}_t$  be the representative investor's information set at time  $t$ .

## 2.1 Stochastic Discount Factor and Its Parameterization

SDF

The prices of all assets satisfy the following asset pricing equation in terms of the stochastic discount factor (SDF):

$$P_t = \mathbb{E} [M_{t,t+1} \exp(-r_f) P_{t+1} | \mathcal{F}_t]. \quad (1)$$

Following the definition of  $r_{t+1}$ , the pricing equation implies that for all assets

$$1 = \mathbb{E} [M_{t,t+1} \exp(r_{t+1}) | \mathcal{F}_t]. \quad (2)$$

We start by parameterizing the SDF by the exponential affine model. Let  $\pi$  be the price of volatility risk and  $\kappa$  be the price of market risk. They are both considered as structural parameters.

**Definition 1.** Parameterize the Stochastic Discount Factor

$$M_{t,t+1}(\pi, \kappa) = \exp(m_0 + m_1 \sigma_t^2 - \pi \sigma_{t+1}^2 - \kappa r_{t+1}). \quad (3)$$

Throughout we assume that the two risks that command nonzero prices are the market risk price and the volatility risk price. Consequently, we only use variation in the first two moments of the data to estimate these parameters.

## 2.2 Parameterizing the Volatility and Return Dynamics

Next, we parameterize the joint distribution of  $\{W_t : t = 1, \dots, T\}$ . Following Han, Khrapov, and Renault (2018), we make the following assumptions. First, the return  $r_t$  and volatility  $\sigma_t^2$  are first-order Markov. Second, there is no Granger-causality from the return to the volatility. Third, returns are independent across time given the volatility. We do allow  $\sigma_t^2$  and  $r_t$  to be contemporaneously correlated, as they are in the data.

We adopt the conditional autoregressive gamma process as in Gouriéroux and Jasiak (2006) and Han, Khrapov, and Renault (2018) for the volatility process. The model is parameterized in terms of the Laplace transform:

$$\mathbb{E} [\exp(-x \sigma_{t+1}^2) | \mathcal{F}_t] = \exp(-A(x) \sigma_t^2 - B(x)) \quad (4)$$

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1. The risk-neutral measure is unobserved due to the lack of option data.

for all  $x \in \mathbb{R}$ . The function  $A(x)$  and  $B(x)$  are parameterized as follows.

**Definition 2.** Parameterize the Volatility Dynamics

$$A(x) := \frac{\rho x}{1 + cx}, \quad (5)$$

$$B(x) := \delta \log(1 + cx), \quad (6)$$

with  $\rho \in [0, 1 - \epsilon]$ ,  $c > \epsilon$ ,  $\delta > \epsilon$  for some  $\epsilon > 0$ .

In this specification,  $\rho$  is a persistence parameter,  $c$  is a scale parameter, and  $\delta$  is a level parameter. We can see this clearly in the following conditional mean and variance formulas for  $\sigma_{t+1}^2$ .

**Remark 1** (Volatility Moment Conditions).

$$\mathbb{E} [\sigma_{t+1}^2 \mid \sigma_t^2] = \rho \sigma_t^2 + c\delta, \quad (7)$$

$$\mathbb{V}\text{ar} [\sigma_{t+1}^2 \mid \sigma_t^2] = 2c\rho\sigma_t^2 + c^2\delta. \quad (8)$$

~~All parameters are identified in these two moments conditions.~~

Next, we model the return dynamics. Similar to the volatility dynamics, the distribution of  $r_t$  given  $\sigma_{t+1}^2$  and  $\sigma_t^2$  is specified in terms of the Laplace transform:

$$\mathbb{E} [\exp(-xr_{t+1}) \mid \mathcal{F}_t, \sigma_{t+1}^2] = \exp(-C(x)\sigma_{t+1}^2 - D(x)\sigma_t^2 - E(x)) \quad (9)$$

for all  $x \in \mathbb{R}$ . The function  $C(x)$ ,  $D(x)$ , and  $E(x)$  are parameterized as follows such that the return has a conditional Gaussian distribution.

**Definition 3.** Parameterize the Return Dynamics

$$C(x) := \psi x - \frac{1 - \phi^2}{2} x^2, \quad (10)$$

$$D(x) := \beta x, \quad (11)$$

$$E(x) := \gamma x \quad (12)$$

with  $\phi \in [-1 + \epsilon, 0]$  for some  $\epsilon > 0$ .

Under this specification, we have the following representation of the conditional mean and variance for  $r_{t+1}$ .

**Remark 2** (Return Moment Conditions).

$$\mathbb{E} [r_{t+1} \mid \sigma_t^2, \sigma_{t+1}^2] = \psi \sigma_{t+1}^2 + \beta \sigma_t^2 + \gamma, \quad (13)$$

$$\mathbb{V}\text{ar} [r_{t+1} \mid \sigma_t^2, \sigma_{t+1}^2] = (1 - \phi^2) \sigma_{t+1}^2. \quad (14)$$

→ should add  $\phi$  is leverage effect parameter it measures var red -

since we don't discuss below (13) & (14), we shouldn't do it here either.

to be consistent with (6)

### 3 Link Functions

So far, we have introduced the following parameters:  $(m_0, m_1, \kappa, \pi)$  in SDF,  $(\rho, c, \delta)$  in the volatility dynamic, and  $(\psi, \beta, \gamma, \phi)$  in the return dynamic. Next, we explore restrictions among these parameters that are consistent with this model. In other words, not all of these parameters can change freely under the structural model.

We use these restrictions to construct link functions between a set of reduced-form parameters and a set of structural parameters. These link functions play an important role on separating the regularly behaved reduced-form parameters from the structural parameters. They also are used to conduct identification robust inference for the structural parameters based on a minimum distance criterion. All of these restrictions are also imposed in the GMM estimation in Han, Khrapov, and Renault (2018). However, because the volatility risk price is weakly identified, they calibrate it instead of estimating it. Given this calibrated value, they proceed to estimate all other parameters with GMM.

#### 3.1 Pricing Equation Restrictions

We first explore restrictions implied by the pricing equation  $\mathbb{E}[M_{t,t+1} \exp(r_{t+1}) | \mathcal{F}_t] = 1$ . We start with a simple result stating that the constants  $m_0$  and  $m_1$  are normalization constants implied by all the other parameters. Thus,  $m_0$  and  $m_1$  are not free parameters to be estimated. Instead, they should take the value given below, once other parameters are specified. These restrictions on  $m_0$  and  $m_1$  are obtained by applying the restriction  $\mathbb{E}[M_{t,t+1} \exp(r_{t+1}) | \mathcal{F}_t] = 1$  to the risk free asset. Applying the same argument to any other asset, we also obtain another set of two restrictions, which can be written in terms of the coefficients  $\beta$  and  $\gamma$  under the linear form of  $D(x)$  and  $E(x)$ .

**Lemma 1.** *Given the parameterization in the model, the pricing equation  $\mathbb{E}[M_{t,t+1} \exp(r_{t+1}) | \mathcal{F}_t] = 1$  implies that*

$$\begin{aligned} m_0 &= E(\kappa) + B(\pi + C(\kappa)), \\ m_1 &= D(\kappa) + A(\pi + C(\kappa)), \end{aligned}$$

and

$$\begin{aligned} \gamma &= B(\pi + C(\kappa - 1)) - B(\pi + C(\kappa)), \\ \beta &= A(\pi + C(\kappa - 1)) - A(\pi + C(\kappa)). \end{aligned}$$

The two equalities on  $\beta$  and  $\gamma$  link them to the market risk price,  $\kappa$ , and the volatility risk,  $\pi$ , through the functions  $A(\cdot), B(\cdot), C(\cdot)$ , which also involve parameters  $(\rho, c, \delta, \psi, \phi)$ . We treat these two equalities as link functions in the minimum distance criterion specified below.

### 3.2 Leverage Effect Restrictions

Following Han, Khrapov, and Renault (2018), we parameterize  $\psi$  as

$$\phi/\sqrt{2c}$$

$$\psi := \frac{\phi}{\sqrt{2c}} - \frac{1 - \phi^2}{2} \sigma_{t+1}^2 + (1 - \phi^2) \kappa \sigma_{t+1}^2.$$

This should be regular " = "  
This is not a definition.  
 $\psi$  was defined  
in  $C(x)$   
above (15)

The first part  $\frac{\phi}{\sqrt{2c}}$  measures the leverage effect arising from the instantaneous correlation between  $r_{t+1}$  and  $\sigma_{t+1}^2$ . The second part is the traditional Jensen effect term that arises from taking expectation of a log-Gaussian random variable. The third term arises from risk-aversion, which is why it is proportional to  $\kappa$ .

we never mentioned  $\phi$  is the leverage effect para.

### 3.3 Structural and Reduced-Form Parameters

Because  $\phi$  is the leverage effect parameter, we group it together with market risk price  $\kappa$  and the volatility risk price  $\pi$  and call  $\theta = (\kappa, \pi, \phi)'$  structural parameters. These structural parameters are estimated by restrictions from this structural model. In contrast, the other parameters in the conditional mean and variance of the return and volatility, see Remark 1 and Remark 2, are simply estimated using these moments, without any model restrictions. As such, we call them the reduced-form parameters. Because  $1 - \phi^2$  shows up in the conditional variance of  $r_{t+1}$ , see Equation (14), we define  $\zeta = 1 - \phi^2$  as a reduced-form parameter and link it to the structural parameter  $\phi$  through this relationship. To sum up, the reduced-form parameters are  $\omega = (\rho, c, \delta, \psi, \beta, \gamma, \zeta)'$ .

my discussion below (14)

Using  $\zeta$  as a reduced-form parameter has the additional benefit of avoiding the estimation of  $\phi$  directly. Estimating  $\phi$  when its true value is close to 0 results in an estimator with a non-standard asymptotic distribution due to the boundary constraint. The inference procedure below does not require estimation of  $\phi$  and is uniform over  $\phi$  even if its true value is on or close to the boundary 0.

$\phi$  was deleted. we should keep it there, as a prep for this.

The link functions between the structural parameter  $\theta$  and the reduced-form parameter  $\omega$  are collected together in For the inference problem studied below, we know  $g(\theta_0, \omega_0) = 0$  when evaluated at the true value of  $\theta$  and  $\omega$ .

add it back

### 3.4 Identification

One of the important contributions of Han, Khrapov, and Renault (2018) is to establish the relationship between the identification of the volatility risk price and the leverage effect. In particular, they show that when the leverage effect parameter  $\phi = 0$ , the volatility risk price  $\pi$  is not identified. To see this result, note that the only source of identification information on  $\pi$  are the first two link functions in  $g(\theta_0, \omega_0) = 0$ , which come from Lemma 1. Clearly, these two equations are independent of  $\pi$  if  $C(\kappa) = C(\kappa - 1)$ . Of course, general, there is no reason to assume that this is the case. Using the definition of  $C(\cdot)$  and Equation (15), we have

$$C(\kappa) - C(\kappa - 1) = \psi - (1 - \phi^2) \left( \kappa - \frac{1}{2} \right) = k\phi.$$

by  $\frac{1}{\sqrt{2c}}$   
k is replaced

delete this!  
we work on the case when this is the case.



Clearly, the strength of identification is governed by the strength of the leverage effect. In other words, we need  $\phi \neq 0$  to identify the volatility risk price  $\pi$ .

Even if  $\phi \neq 0$ , we do not know it. In practice, with a finite-sample size and different types of noise in the data, such as measurement errors and omitted variables, a much more substantial leverage effect is required to obtain a standard identification situation where the noise in the data is negligible compared to the information to identify  $\pi$ . However, if only a small leverage effect is found, as in Bandi and Renò (2012) and Aït-Sahalia, Fan, and Li (2013), or the magnitude of the leverage effect is completely unknown, an identification robust procedure is needed to conduct inference in this problem. We provide such a procedure now.

## 4 Robust Inference for Risk Prices

### 4.1 Asymptotic Distribution of the Reduced-Form Parameter

Write  $\omega := (\omega_1, \omega_2, \omega_3)'$ , where  $\omega_1 = (\rho, c, \delta) \in O_1$ ,  $\omega_2 = (\gamma, \beta, \psi) \in O_2$ , and  $\omega_3 = \zeta \in O_3$ . The parameter space for  $\omega$  is  $O = O_1 \times O_2 \times O_3 \subset R^{d_\omega}$ . The true value of  $\omega$  is assumed to be in the interior of the parameter space.

Below we describe the estimator  $\hat{\omega} := (\hat{\omega}_1, \hat{\omega}_2, \hat{\omega}_3)'$  and provide its asymptotic distribution. We estimate these parameters separately because  $\omega_1$  only shows up in the conditional mean and variance of  $\sigma_{t+1}^2$ ,  $\omega_2$  only shows up in the conditional mean of  $r_{t+1}$ , and  $\omega_3$  only shows up in the conditional variance of  $r_{t+1}$ .

We first estimate  $\omega_1 = (\rho, c)$  based on the conditional mean and variance of  $\sigma_{t+1}^2$ , which can be equivalently written as

$$\begin{aligned} E[\sigma_{t+1}^2 | \sigma_t^2] &= A \text{ and } E[\sigma_{t+1}^4 | \sigma_t^2] = B, \text{ where} \\ A &= \rho\sigma_t^2 + c\delta \text{ and } B = A^2 + (2c\rho\sigma_t^2 + c^2\delta). \end{aligned} \quad (16)$$

Because the conditional mean of  $\sigma_{t+1}^2$  and  $\sigma_{t+1}^4$  are linear and quadratic functions, respectively, of the conditioning variable  $\sigma_t^2$ , without loss of efficiency, they can be transformed to the unconditional moments

$$E[h_t(\omega_{10})] = 0, \text{ where } h_t(\omega_1) = [(1, \sigma_t^2) \otimes (\sigma_{t+1}^2 - A), (1, \sigma_t^2, \sigma_t^4) \otimes (\sigma_{t+1}^4 - B)]', \quad (17)$$

and  $\omega_{10}$  represents the true value of  $\omega_1$ . The two-step GMM estimator of  $\omega_1$  is

$$\hat{\omega}_1 = \arg \min_{\omega_1 \in O_1} \left( T^{-1} \sum_{t=1}^T h_t(\omega_1) \right)' \hat{V}_1 \left( T^{-1} \sum_{t=1}^T h_t(\omega_1) \right), \quad (18)$$

where  $\hat{V}_1$  is a consistent estimator of  $V_1 = \sum_{m=-\infty}^{\infty} \text{Cov}[h_t(\omega_{10}), h_{t+m}(\omega_{10})]$ .

We estimate  $\omega_2$  by the generalized least squares (GLS) estimator because the conditional mean of  $r_{t+1}$  is a linear function of the conditioning variable  $\sigma_t^2$  and  $\sigma_{t+1}^2$  and the conditional variance is



proportional to  $\sigma_{t+1}^2$ . The GLS estimator of  $\omega_2$  is

$$\hat{\omega}_2 = \left( \sum_{t=1}^T x_t x_t' \right)^{-1} \sum_{t=1}^T x_t y_t, \text{ where} \\ x_t = \sigma_{t+1}^{-1} (1, \sigma_t^2, \sigma_{t+1}^2)' \text{ and } y_t = \sigma_{t+1}^{-1} r_{t+1}. \quad (19)$$

We estimate  $\omega_3$  by the sample variance estimator

$$\hat{\omega}_3 = T^{-1} \sum_{t=1}^T (y_t - \hat{y}_t)^2, \text{ where } \hat{y}_t = x_t' \hat{\omega}_2. \quad (20)$$

Let  $P$  denote the distribution of the data  $\{W_t = (r_{t+1}, \sigma_{t+1}^2) : t \geq 1\}$  and  $\mathcal{P}$  denote the parameter space of  $P$ . Note that the true values of the structural parameter and the reduced-form parameters are all determined by  $P$ . We allow  $P$  to change with  $T$ . For notational simplicity, the dependence on  $P$  and  $T$  is suppressed.

Let

$$f_t(\omega) = \begin{pmatrix} h_t(\omega_1) \\ x_t(y_t - x_t' \omega_2) \\ (y_t - x_t' \omega_2)^2 \end{pmatrix} \in R^{d_f} \text{ and } V = \sum_{m=-\infty}^{\infty} \text{Cov}[f_t(\omega_0), f_{t+m}(\omega_0)]'. \quad (21)$$

The estimator  $\hat{\omega}$  defined above is based on the first moment of  $f_t(\omega)$ . Thus, the limiting distribution of  $\hat{\omega}$  relates to the limiting distribution of  $T^{-1/2} \sum_{t=1}^T (f_t(\omega_0) - \mathbb{E}[f_t(\omega_0)])$  following from the central limit theorem. Furthermore, because  $\omega_1$  is the GMM estimator based on some nonlinear moment conditions, we need uniform convergence of the sample moments and their derivatives to show the consistency and asymptotic normality of  $\hat{\omega}_1$ . These uniform convergence follows from the uniform law of large numbers. Because  $\hat{\omega}_2$  is a simple OLS estimator by regressing  $y_t$  and  $x_t$ , we need the regressors to not exhibit multicollinearity. We make the necessary assumptions below. All of them are easily verifiable with weakly dependent time series data.

Let  $\hat{V}$  denote a heteroskedasticity and autocorrelation consistent (HAC) estimator of  $V$ . The estimator  $\hat{V}_1$  is a submatrix of  $\hat{V}$  associate with  $V_1$ . Let  $H_t(\omega_1) = \partial h_t(\omega_1) / \partial \omega_1'$ .

**Assumption R.** The following conditions hold uniformly over  $P \in \mathcal{P}$ , for some fixed  $0 < C < \infty$ .

1.  $T^{-1} \sum_{t=1}^T (h_t(\omega_1) - \mathbb{E}[h_t(\omega_1)]) \rightarrow_p 0$  and  $T^{-1} \sum_{t=1}^T (H_t(\omega_1) - \mathbb{E}[H_t(\omega_1)]) \rightarrow_p 0$ ,  $\mathbb{E}[H_t(\omega_1)]$  is continuous in  $\omega_1$ , all uniformly over the parameter space of  $\omega_1$ .
2.  $T^{-1} \sum_{t=1}^T (x_t x_t' - \mathbb{E}[x_t x_t']) \rightarrow_p 0$ .
3.  $V^{-1/2} \{T^{-1/2} (\sum_{t=1}^T f_t(\omega_0) - \mathbb{E}[f_t(\omega_0)])\} \rightarrow_d N(0, I)$  and  $\hat{V} - V \rightarrow_p 0$ .
4.  $C^{-1} \leq \lambda_{\min}(A) \leq \lambda_{\max}(A) \leq C$  for  $A = V, \mathbb{E}[H_t(\omega_{1,0})' H_t(\omega_{1,0})], \mathbb{E}[x_t x_t'], \mathbb{E}[z_t z_t']$ , where  $z_t = (1, \sigma_t^2, \sigma_t^4)'$ .

Let  $H(\omega_1) = \mathbb{E}[H_t(\omega_1)]$  and  $\bar{H}(\omega_1) = T^{-1} \sum_{t=1}^T H_t(\omega_1)$ . Define

$$\begin{aligned}\mathcal{B} &= \text{diag}\{[H(\omega_{10})V_1^{-1}H(\omega_{10})]^{-1}H(\omega_{10})V_1^{-1}, \mathbb{E}[x_t x_t']^{-1}, 1\}, \\ \hat{\mathcal{B}} &= \text{diag}\{[\bar{H}(\hat{\omega}_1)' \hat{V}_1^{-1} \bar{H}(\hat{\omega}_1)]^{-1} \bar{H}(\hat{\omega}_1)' \hat{V}_1^{-1}, [T^{-1} \sum_{t=1}^T x_t x_t']^{-1}, 1\}.\end{aligned}\quad (22)$$

The following lemma provides the asymptotic distribution of the reduced-form parameter and a consistent estimator of its asymptotic covariance. Note that we put the asymptotic covariance on the left side of the convergence to allow the distribution of the data to change with sample size  $T$ .

**Lemma 2.** Suppose *Assumption R* holds. The following results hold uniformly over  $P \in \mathcal{P}$ .

$$\xi_T := \Omega^{-1/2} T^{-1/2} (\hat{\omega} - \omega_0) \rightarrow_d \xi \sim N(0, I), \text{ where } \Omega = \mathcal{B} V \mathcal{B}', \quad (23)$$

and

$$\hat{\Omega} - \Omega \rightarrow_p 0, \text{ where } \hat{\Omega} = \hat{\mathcal{B}} \hat{V} \hat{\mathcal{B}}'. \quad (24)$$

## 4.2 Weak Identification

The true value of the structural parameter  $\theta$  and the reduced-form parameter  $\omega$  satisfy the link function  $g(\theta_0, \omega_0) = 0$ . In a standard problem without any identification issues, we can estimate  $\theta_0$  by the minimum distance estimator,  $\hat{\theta} = (\hat{\kappa}, \hat{\pi}, \hat{\phi})'$ , which minimizes  $Q_T(\theta) = g(\theta, \hat{\omega})' W_T g(\theta, \hat{\omega})$  for some weighting matrix  $W_T$  and construct tests and confidence sets for  $\theta_0$  using an asymptotically normal approximation for  $T^{1/2}(\hat{\theta} - \theta_0)$ . However, this standard method does not work in the present problem when  $\pi_0$  is only weakly identified. In this case,  $g(\theta, \hat{\omega})$  is almost flat in  $\pi$  and the minimum distance estimator of  $\hat{\pi}$  is not even consistent. To make the problem even more complicated, the inconsistency of  $\hat{\pi}$  has a spillover effect on  $\hat{\kappa}$  and  $\hat{\phi}$ , making the distribution of  $\hat{\kappa}$  and  $\hat{\phi}$  non-normal even in large samples.

Before presenting the robust CS, we first introduce some useful quantities and provide some heuristic discussions of the identification problem and its consequences. Let  $G(\theta, \omega)$  denote the partial derivative of  $g(\theta, \omega)$  with respect to (w.r.t.)  $\omega$ . Let  $g_0(\theta) = g(\theta, \omega_0)$  and  $G_0(\theta) = G(\theta, \omega_0)$  be the link function and its derivative evaluated at  $\omega_0$  and  $\hat{g}(\theta) = g(\theta, \hat{\omega})$  and  $\hat{G}(\theta) = G(\theta, \hat{\omega})$  be the same quantities evaluated at the estimator  $\hat{\omega}$ . The delta method gives

$$\eta_T(\theta) := T^{1/2} [\hat{g}(\theta) - g_0(\theta)] = G_0(\theta) \Omega^{1/2} \cdot \xi_T + o_p(1), \quad (25)$$

where  $\xi_T \rightarrow_d N(0, I)$  following *Lemma 2*. Thus,  $\eta_T(\cdot)$  weakly converges to a Gaussian process  $\eta(\cdot)$  with covariance function  $\Sigma(\theta_1, \theta_2) = G_0(\theta_1) \Omega G_0(\theta_2)'$ .

Following *Equation (25)*, we can write  $T^{1/2} \hat{g}(\theta) = \eta_T(\theta) + T^{1/2} g_0(\theta)$ , where  $\eta_T(\theta)$  is the noise from the reduced-form parameter estimation and  $T^{1/2} g_0(\theta)$  is the signal from the link function. Under weak identification,  $g_0(\theta)$  is almost flat in  $\theta$ , modeled as the signal  $T^{1/2} g_0(\theta)$  being finite even for  $\theta \neq \theta_0$  and  $T \rightarrow \infty$ . Thus, the signal and the noise are of the same order of magnitude, yielding

an inconsistent minimum distance estimator  $\hat{\theta}$ . This is in contrast with the strong identification scenario, where  $T^{1/2}g_0(\theta) \rightarrow \infty$  for  $\theta \neq \theta_0$  as  $T \rightarrow \infty$  and  $g_0(\theta_0) = 0$ . In this case, the signal is so strong that the minimum distance estimator is consistent.

The identification strength of  $\theta_0$  is determined by the function  $T^{1/2}g_0(\theta)$ . However, this function is unknown and cannot be consistently estimated (due to  $T^{1/2}$ ). Thus, we take the conditional inference procedure as in Andrews and Mikusheva (2016) and view  $T^{1/2}g_0(\theta)$  as an infinite dimensional nuisance parameter for the inference of  $\theta_0$ . The goal is to construct robust CS for  $\theta_0$  that has correct size asymptotically regardless of this unknown nuisance parameter.

confidence set

### 4.3 Conditional QLR Test

We construct a CS for  $\theta \in \Theta := [0, M_1] \times [-M_2, 0] \times [1 - \epsilon, 0]$  by inverting the test  $H_0 : \theta = \theta_0$  vs  $H_1 : \theta \neq \theta_0$ , where  $M_1$  and  $M_2$  are large constants and  $\epsilon$  is a small constant. The test statistic is a QLR statistic that takes the form

positive positive

$$QLR(\theta_0) := T\hat{g}(\theta_0)' \hat{\Sigma}(\theta_0, \theta_0)^{-1} \hat{g}(\theta_0) - \min_{\theta \in \Theta} T\hat{g}(\theta)' \hat{\Sigma}(\theta, \theta)^{-1} \hat{g}(\theta), \quad (26)$$

where  $\hat{\Sigma}(\theta_1, \theta_2) = \hat{G}(\theta_1) \hat{\Omega} \hat{G}(\theta_2)'$  and  $\hat{\Omega}$  is the consistent estimator of  $\Omega$  defined above.

Andrews and Mikusheva (2016) provide the conditional QLR test in a nonlinear GMM problem, where  $\hat{g}(\theta)$  is replaced by a sample moment. The same method can be applied to the present nonlinear minimum distance problem. Following Andrews and Mikusheva (2016), we first project  $\hat{g}(\theta)$  onto  $\hat{g}(\theta_0)$  and construct a residual process

$$\hat{r}(\theta) = \hat{g}(\theta) - \hat{\Sigma}(\theta, \theta_0) \hat{\Sigma}(\theta_0, \theta_0)^{-1} \hat{g}(\theta_0). \quad (27)$$

The limiting distribution of  $\hat{r}(\theta)$  and  $\hat{g}(\theta_0)$  are Gaussian and independent. Thus, conditional on  $\hat{r}(\theta)$ , the asymptotic distribution of  $\hat{g}(\theta)$  no longer depends on the nuisance parameter,  $T^{1/2}g_0(\theta)$ , making the procedure robust to all identification strengths.

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Specifically, we obtain the  $1 - \alpha$  conditional quantile of the QLR statistic, denoted by  $c_{1-\alpha}(r, \theta_0)$ , as follows. For  $b = 1, \dots, B$ , we take independent draws  $\eta_b^* \sim N(0, \hat{\Sigma}(\theta_0, \theta_0))$  and produce a simulated process,

remove

$$g_b^*(\theta) := \hat{r}(\theta) + \hat{\Sigma}(\theta, \theta_0) \hat{\Sigma}(\theta_0, \theta_0)^{-1} \eta_b^*, \quad (28)$$

and a simulated statistic:

remove

$$QLR_b^*(\theta_0) := T\hat{g}(\theta_0)' \hat{\Sigma}(\theta_0, \theta_0)^{-1} \hat{g}(\theta_0) - \min_{\theta \in \Theta} Tg_b^*(\theta)' \hat{\Sigma}(\theta, \theta)^{-1} g_b^*(\theta). \quad (29)$$

Let  $b_0 = \lceil (1 - \alpha)B \rceil$ , the smallest integer greater than or equal to  $(1 - \alpha)B$ . Then the critical value  $c_{1-\alpha}(r, \theta_0)$  is the  $b_0^{th}$  smallest value among  $\{QLR_b^*, b = 1, \dots, B\}$ . We execute the steps reported in Algorithm 1 to form a robust confidence set for  $\theta$ .

To obtain confidence intervals for each element of  $\theta_0$ , one simple solution is to project the

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How about all identification scenarios

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**Algorithm 1** Construing the Confidence Set
 

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1. Estimate the reduced-form parameter  $\hat{\omega} = (\hat{\omega}_1, \hat{\omega}_2, \hat{\omega}_3)'$  following the estimators defined in Equation (18), Equation (19), and Equation (20).
2. Obtain a consistent estimator of its asymptotic covariance  $\hat{\Omega} = \hat{\mathcal{B}}\hat{V}\hat{\mathcal{B}}'$ , where  $\hat{\mathcal{B}}$  is defined in Equation (22) and  $\hat{V}$  is a HAC estimator of  $V$ .
3. For  $\theta_0 \in \Theta$ ,
  - (a) Construct the QLR statistic  $QLR(\theta_0)$  in Equation (26) using  $g(\theta, \omega)$ ,  $G(\theta, \omega)$ ,  $\hat{\omega}$ , and  $\hat{\Omega}$ .
  - (b) Compute the residual process  $\hat{r}(\theta)$  in Equation (27).
  - (c) Given  $\hat{r}(\theta)$ , compute the critical value  $c_{1-\alpha}(r, \theta_0)$  described above.
4. Repeat these steps for different values of  $\theta_0$ . Construct a confidence set by collecting the null values that are not rejected, i.e., the nominal level  $1 - \alpha$  confidence set for  $\theta_0$  is

$$CS_T = \{\theta_0 : QLR_T(\theta_0) \leq c_{1-\alpha}(r, \theta_0)\}. \quad (30)$$


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CS constructed above to each axis. The resulting confidence interval also has correct coverage. An alternative solution is to first concentrate out the nuisance parameters before applying the conditional inference approach above, see Andrews and Mikusheva (2016, Section 5). However, this concentration approach only works when the nuisance parameter is strongly identified. In the present set-up, this approach does not work for  $\kappa$  and  $\phi$  because the nuisance parameter  $\pi$  is weakly identified.

**Assumption S.** The following conditions hold over  $P \in \mathcal{P}$ , for any  $\theta$  in its parameter space, and any  $\omega$  in some fixed neighborhood around its true value, for some fixed  $0 < C < \infty$ .

1.  $g(\theta, \omega)$  is partially differentiable in  $\omega$ , with partial derivative  $G(\theta, \omega)$  that satisfies  $\|G(\theta_1, \omega) - G(\theta_2, \omega)\| \leq C\|\theta_1 - \theta_2\|$  and  $\|G(\theta, \omega_1) - G(\theta, \omega_2)\| \leq C\|\omega_1 - \omega_2\|$ .
2.  $C^{-1} \leq \lambda_{\min}(G(\theta, \omega)'G(\theta, \omega)) \leq \lambda_{\max}(G(\theta, \omega)'G(\theta, \omega)) \leq C$ .

**Theorem 3.** Suppose Assumption R and Assumption S hold. Then,

$$\liminf_{T \rightarrow \infty} \inf_{P \in \mathcal{P}} \Pr(\theta_0 \in CS_T) \geq 1 - \alpha.$$

This Lemma states that the confidence set constructed by the conditional QLR test has correct asymptotic size. Uniformity is important for this confidence set to cover the true parameter with a probability close to  $1 - \alpha$  in finite-samples. This uniform result is established over a parameter space  $\mathcal{P}$  that allows for the weak identification of the structural parameter  $\theta$ .

## 5 Simulations and Empirical Results

To be added.

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## Appendix A Proofs

### A.1 Proof of **Lemma 1**

*Proof.* For the risk free asset,  $r_{t+1} = 0$ . Therefore, we have

$$\begin{aligned}
 1 &= E \left[ \exp \left( m_0 + m_1 \sigma_t^2 - \pi \sigma_{t+1}^2 - \kappa r_{t+1} \right) \mid \mathcal{F}_t \right] \\
 &= \exp(m_0 + m_1 \sigma_t^2) E \left[ \exp \left( -\pi \sigma_{t+1}^2 \right) E \left[ \exp \left( -\kappa r_{t+1} \right) \mid \mathcal{F}_t, \sigma_{t+1}^2 \right] \mid \mathcal{F}_t \right] \\
 &= \exp(m_0 - E(\kappa) + m_1 \sigma_t^2 - D(\kappa) \sigma_t^2) E \left[ \exp \left( -\pi \sigma_{t+1}^2 - C(\kappa) \sigma_{t+1}^2 \right) \mid \mathcal{F}_t \right] \\
 &= \exp(m_0 - E(\kappa) + m_1 \sigma_t^2 - D(\kappa) \sigma_t^2 - A(\pi + C(\kappa)) \sigma_t^2 - B(\pi + C(\kappa))),
 \end{aligned}$$

where the first equality follows from the pricing equation, the second equality follows from the law of iterated expectations, the third equation uses the Laplace transform for  $r_{t+1}$  in **Equation (9)**, and the last equality follows from the Laplace transform for  $\sigma_{t+1}^2$  in **Equation (4)**. Since  $M_{t,t+1}$



must integrate to 1, the constant term and coefficient for  $\sigma_t^2$  must equal 0, which gives the claimed result for  $m_0$  and  $m_1$ .

We can apply the same argument above to any asset  $r_{t+1}$ . This gives the same result, except  $\kappa$  is replaced by  $\kappa - 1$  throughout. This implies that the two equalities for  $m_0$  and  $m_1$  also hold with  $\kappa$  replaced by  $\kappa - 1$ . Therefore,

$$\begin{aligned} E(\kappa - 1) + B(C(\kappa - 1) + \pi) &= E(\kappa) + B(C(\kappa) + \pi), \\ D(\kappa - 1) + A(C(\kappa - 1) + \pi) &= D(\kappa) + A(C(\kappa) + \pi). \end{aligned}$$

The claimed results for  $\gamma$  and  $\beta$  follow from  $\gamma = \mathbb{E}(\kappa) - E(\kappa - 1)$  and  $\beta = D(\kappa) - D(\kappa - 1)$  under the linear specification of  $E(x) = \gamma x$  and  $D(x) = \beta x$ . □

## A.2 Proof of Lemma 2

*Proof.* Under the assumption that (i)  $\mathbb{E}(z_t z_t')$  has the smallest eigenvalue bounded away from 0 and (ii)  $c > \varepsilon$  and  $\delta > \varepsilon$  for some  $\varepsilon > 0$ , we not only have  $\omega_{10}$  as a uniquely minimizer of  $\|\mathbb{E}[h_t(\omega_1)]\|$  but also have a uniform positive lower bound for  $\|\mathbb{E}[h_t(\omega_1)]\|$  for  $\|\omega_1 - \omega_{10}\| \geq \varepsilon$ . Thus, consistency of  $\hat{\omega}_1$  follows from standard arguments for the consistency of a GMM estimator under a uniform convergence of the criterion under [Assumption R](#) (1) and (2).

Let  $\bar{h}(\omega_1) = T^{-1} \sum_{t=1}^T h_t(\omega_1)$  and  $\bar{H}(\omega) = T^{-1} \sum_{t=1}^T H_t(\omega_1)$ . By construction, the estimator satisfies the first-order condition

$$\begin{aligned} 0 &= \begin{pmatrix} \bar{H}(\hat{\omega}_1)' \hat{V}_1^{-1} \bar{h}(\hat{\omega}_1) \\ T^{-1} \sum_{t=1}^T x_t(y_t - x_t' \hat{\omega}_2) \\ \hat{\omega}_3 - T^{-1} \sum_{t=1}^T (y_t - \hat{y}_t)^2 \end{pmatrix} \\ &= \begin{pmatrix} \bar{H}(\hat{\omega}_1)' \hat{V}_1^{-1} \bar{h}(\omega_{10}) + \bar{H}(\hat{\omega}_1)' \hat{V}_1^{-1} \bar{H}(\tilde{\omega}_1)(\hat{\omega}_1 - \omega_{10}) \\ T^{-1} \sum_{t=1}^T x_t(y_t - x_t' \omega_{20}) - T^{-1} \sum_{t=1}^T x_t x_t' (\hat{\omega}_2 - \omega_{20}) \\ (\hat{\omega}_3 - \omega_3) + \omega_3 - T^{-1} \sum_{t=1}^T (y_t - x_t \hat{\omega}_2)^2 \end{pmatrix}, \end{aligned} \quad (31)$$

where the second equality follows from a mean value expansion of  $\bar{h}(\hat{\omega}_1)$  around  $\omega_{10}$ , with  $\tilde{\omega}_1$  between  $\omega_{10}$  and  $\hat{\omega}_1$ . Let

$$\tilde{\mathcal{B}} = \text{diag} \left\{ [\bar{H}(\hat{\omega}_1)' \hat{V}_1^{-1} \bar{H}(\tilde{\omega}_1)]^{-1} \bar{H}(\hat{\omega}_1)' \hat{V}_1^{-1}, [T^{-1} \sum_{t=1}^T x_t x_t']^{-1}, 1 \right\}. \quad (32)$$

Then solving for  $\hat{\omega} - \omega$ , [Equation \(31\)](#) implies that

$$T^{1/2} (\hat{\omega} - \omega) = \tilde{\mathcal{B}} \cdot T^{-1/2} \sum_{t=1}^T \begin{pmatrix} -h_t(\omega_{10}) \\ x_t(y_t - x_t' \omega_{20}) \\ (y_t - x_t \hat{\omega}_2)^2 - \omega_3 \end{pmatrix}$$

(using  $\omega_3 = \mathbb{E}[(y_t - x'_t \omega_{20})^2]$ )

$$= \tilde{\mathcal{B}} \cdot T^{-1/2} \sum_{t=1}^T \begin{pmatrix} -h_t(\omega_{10}) \\ x_t(y_t - x'_t \omega_{20}) \\ (y_t - x'_t \omega_{20})^2 - \mathbb{E}[(y_t - x'_t \omega_{20})^2] \end{pmatrix} + \begin{pmatrix} 0 \\ 0 \\ \varepsilon_T \end{pmatrix}, \quad (33)$$

We can solve for  $\varepsilon$ :

$$\begin{aligned} \varepsilon_T &= T^{-1/2} \sum_{t=1}^T \left[ (y_t - x'_t \hat{\omega}_2)^2 - (y_t - x'_t \omega_{20})^2 \right] \\ &= 2T^{-1} \sum_{t=1}^T (y_t - x'_t \omega_{20}) x'_t \left[ T^{1/2} (\hat{\omega}_2 - \omega_{20}) \right] + o_p(1) \\ &= o_p(1). \end{aligned} \quad (34)$$

because  $T^{-1} \sum_{t=1}^T (y_t - x'_t \omega_{20}) x'_t \rightarrow_p 0$  and  $T^{1/2}(\hat{\omega}_2 - \omega_{20}) = O_p(1)$  following [Assumption R](#). In addition,

$$\tilde{\mathcal{B}} \rightarrow_p \mathcal{B} \quad (35)$$

following from the consistency of  $\hat{\omega}_1$  and [Assumption R](#). Finally, the desirable result follows from [Equation \(33\)](#)–[Equation \(35\)](#) and [Assumption R](#). The consistency of  $\hat{\Omega}$  follows from the consistency of  $\hat{\mathcal{B}}$  and  $\hat{V}$ . □

### A.3 Proof of [Theorem 3](#)

*Proof.* We obtain this result by applying Andrews and Mikusheva ([2016](#), Theorem 1). We now verify Assumptions 1-3 in Andrews and Mikusheva ([2016](#)). To show weak convergence  $\eta_T(\cdot)$  to  $\eta(\cdot)$  uniformly over  $\mathcal{P}$ , note that by a second-order Taylor expansion,

$$\begin{aligned} \eta_T(\theta) &:= T^{1/2} [\hat{g}(\theta) - g_0(\theta)] = G_0(\theta) \Omega^{1/2} \xi_T + \delta_T, \text{ where} \\ \xi_T &= \Omega^{-1/2} T^{1/2} (\hat{\omega} - \omega_0), \text{ and } \delta_T = (G(\theta, \tilde{\omega}) - G(\theta, \omega_0)) T^{1/2} (\hat{\omega} - \omega_0) \end{aligned} \quad (36)$$

and  $\tilde{\omega}$  is between  $\hat{\omega}$  and  $\omega_0$ . Because  $\|G(\theta, \tilde{\omega}) - G(\theta, \omega_0)\| \leq C \|\tilde{\omega} - \omega_0\|$ ,  $\delta_T = o_p(1)$  uniformly over  $\mathcal{P}$  following [Lemma 2](#). To show  $G_0(\theta) \Omega^{1/2} \xi_T$  weakly converges to  $\eta(\cdot)$ , it is sufficient to show (i) the pointwise convergence

$$\begin{pmatrix} G_0(\theta_1) \Omega^{1/2} \xi_T \\ G_0(\theta_2) \Omega^{1/2} \xi_T \end{pmatrix} \rightarrow_d \begin{pmatrix} \eta(\theta_1) \\ \eta(\theta_2) \end{pmatrix}, \quad (37)$$

which follows from [Lemma 2](#), and (ii) the stochastic equicontinuity condition, i.e., for every  $\varepsilon > 0$  and  $\xi > 0$ , there exists a  $\delta > 0$  such that

$$\limsup_{T \rightarrow \infty} \Pr \left( \sup_{P \in \mathcal{P}} \sup_{\|\theta_1 - \theta_2\| \leq \delta} \left\| G_0(\theta_1) \Omega^{1/2} \xi_T - G_0(\theta_2) \Omega^{1/2} \xi_T \right\| > \varepsilon \right) < \xi. \quad (38)$$

For some  $C < \infty$ , we have  $\|G_0(\theta_1) - G_0(\theta_2)\| \leq C \|\theta_1 - \theta_2\|$  under a uniform bound for the derivative in [Assumption S](#), and we have  $\|\Omega^{1/2}\| \leq C$  under [Assumption R](#) because  $F$  and  $V$  both have bounded largest eigenvalue. Thus,

$$\begin{aligned} & \limsup_{T \rightarrow \infty} \Pr \left( \sup_{P \in \mathcal{P}} \sup_{\|\theta_1 - \theta_2\| \leq \delta} \left\| G_0(\theta_1) \Omega^{1/2} \xi_T - G_0(\theta_2) \Omega^{1/2} \xi_T \right\| > \varepsilon \right) \\ & \leq \limsup_{T \rightarrow \infty} \Pr \left( C^2 \sup_{P \in \mathcal{P}} \|\xi_T\| > \frac{\varepsilon}{\delta} \right). \end{aligned} \quad (39)$$

Because  $\xi_T = O_p(1)$  uniformly over  $P \in \mathcal{P}$ , there exists  $\delta$  such that  $\varepsilon/\delta$  is large enough to make the right hand side of the inequality in [Equation \(39\)](#) smaller than  $\xi$ .

Assumptions 2 and 3 of Andrews and Mikusheva ([2016](#), Theorem 1) follow from [Assumption R](#). □