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Keywords

China, silver standard, exchange rate, parity deviation, Bayesian structural VAR

JEL Classification

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This paper assesses the debate about the demise of the Chinese silver standard in the mid 1930s. One side argues the U.S. Silver Purchase Act of June 1934 drained China of silver, which caused deflation and economic crises. A related claim is the Chinese silver standard was intrinsically unstable. These hypotheses are evaluated by estimating Bayesian structural VARs with drifting parameters on China-U.K. and China-U.S. samples from April 1912 to September 1934. We find instability in the Chinese silver standard peaked during the recession of the early 1920s and the Great Depression. Hence, neither the U.S. Silver Purchase Act of June 1934 nor a design flaw led to the end of the Chinese silver standard.

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1. Introduction

The Chinese economy ran on a commodity monetary standard linked to silver for hundreds of years before its end in the mid 1930s. By 1900, payments for regional trade within China were settled in local money markets each using its own unit of account denominated in *tael* (*liang* or Chinese ounce). Embedded in a local *tael* was a price for silver, but China took the world price of silver as given. As a result, it defined parity for the Chinese silver standard.

The Nanjing government replaced the Chinese silver standard with the *fabi* (fiat currency) in November 1935 (1935m11).[1] However, Leavens (1939, pp. 300–302) describes the Nanjing government as breaking the Chinese silver standard a year earlier. A customs duty on exports of silver was levied by the Nanjing government to reduce these flows in 1934m10. At the same time, it imposed an equalization charge (step-up duty) on silver exports and began a regime of managed exchange rates. The intent of the last two actions was to restore the Chinese silver standard to parity by moving the domestic price of silver closer to the world price. These policy tools were needed because beginning in 1934m10 the customs duty on exports of silver decoupled the Chinese silver standard from the world price of silver.

There are theories to explain the Nanjing government's motives for ending the Chinese silver standard. Friedman and Schwartz (1963), Chang (1988), Friedman (1992), Burdekin (2008), Silber (2019), and Dean (2020) posit the Chinese silver standard was damaged by the U.S. Silver Purchase Act passed by Congress in 1934m06. The Act instructed the Treasury to buy silver at a price greater than in the spot market. Accordingly, silver left China causing deflation, financial and economic crises, and the Nanjing government's policy responses starting in 1934m10.

[1] The Nanjing government substituted the *yuan* (Chinese silver dollar) for the *tael* in 1933m04. This was only a change in numeraire. As Bratter (1933) notes, the rate to change *yuan* into *tael* was fixed. Banks still paid silver for their notes and deposits on demand. When the *fabi* became legal tender in China, convertibility ceased.

Shiroyama (2008), Ho and Lai (2013), and Ho (2014) offer a related hypothesis that instability was built into the Chinese silver standard. They argue its fragility was accentuated by the U.S. Silver Purchase Act because parity for the Chinese silver standard was defined by the world price of silver. Hence, domestic fiscal and monetary policy could not smooth shocks that produced deviations from parity of the Chinese silver standard.

This paper frames its evaluation of the two theories with the exchange rate-risk premium model of Engel (2016). He separated the log of the exchange rate, e_t, into trend and transitory components. The latter is the negative of the expected future paths of the cross-country interest rate spread, i_t, and deviations from parity, ρ_t, which is the risk premium. Starting from this decomposition, we construct a regression that runs currency returns, Δe_t, on its own lags, i_t, the cross-country inflation gap, π_t, ρ_t, and their lags. The regression employs the Beveridge and Nelson (1981) decomposition to impose a random walk on the trend of e_t, recognizes that Δe_t imposes differencing on its trend and transitory components, applies a result in Nason and Rogers (2008), and assumes i_t, π_t, ρ_t, and Δe_t are generated by a reduced-form VAR.

We identify several structural VARs (SVARs) that are grounded in this regression. Our baseline SVAR, SVAR-BL, consists of it and regressions of i_t, π_t, and ρ_t that place zeros on the nine off-diagonal impact coefficients. These restrictions globally identify international finance, cross-country nominal demand, risk premium, and trend exchange rate shocks. These shocks are also globally identified in 10 alternative SVARs that depart from SVAR-BL with subsets of the off-diagonal impact coefficients in the π_t and ρ_t regressions acting as free parameters.

Assessing the merits of the two hypotheses is an empirical exercise. With this in mind, we compile China-U.K. and China-U.S. samples from 1912M04 to 1934M09. The Chinese silver standard has not been studied before using these monthly samples. The samples consist of i_t,

π_t, ρ_t, and Δe_t, given China is the home country and the U.K. or U.S. is the foreign economy.

There is research disputing the two hypotheses of the end of the Chinese silver standard. Rawski (1993) shows China had inflation after passage of the Silver Purchase Act in 1934M06. Chen, Li, and Xie (2022) report evidence on pre-1935 annual data that the lack of deflation stemmed from Chinese banks issuing notes backed by silver that replaced the missing commodity money. Ho, Lai, and Gau (2013) find the U.S. dollar-*tael* exchange rate moved with the world price of silver on a pre-1934M10 sample. Finally, Brandt and Sargent (1989) focus on two goals they argue animated the Nanjing government's actions from 1934M10 to creating the *fabi* in 1935M11. The Nanjing government coveted the income generated in domestic and international money markets in China and sought to undo the constraints the Chinese silver standard imposed on domestic fiscal and monetary policy. This argues that neither the U.S. Silver Purchase Act nor its inherent fragility led the Chinese silver standard to fail.

The empirical literature studying the Chinese silver standard often uses time series econometric methods. Leading examples are Lai and Gau (2003), Burdekin (2008), Ho and Lai (2013, 2016), Ho, Lai, and Gau (2013), Ho (2014), Jacks, Yan, and Zhao (2017), Zhao and Zhao (2018), Ma and Zhao (2020), Palma and Zhao (2021), and Chen, Li, and Xie (2022). Our paper extends this literature in three ways. As already mentioned, we identify SVARs starting from the exchange rate-risk premium model of Engel (2016) and amass new China-U.K. and China-U.S. samples from 1912M04 to 1934M09. Third, the SVARs are estimated with time-varying parameters (TVPs) and stochastic volatility (SV).

The TVPs and SVs are included in the SVARs to handle the turmoil affecting China, the U.K., and U.S. from 1912 to 1935. China suffered political upheaval during these years. Important episodes were the Warlord Era, which began in 1916M06 and ended with the Northern

Expedition of 1926м07–1928м12 that unified China under the Nanjing government, the Civil War starting in 1927м08, and the invasion of Manchuria by Japan in 1931м09. For the U.K. and U.S, the First World War, its aftermath, and the Great Depression defined the era. These events drove changes in the underlying economic environment that are reflected in TVPs and SV.

We draft the Metropolis in Gibbs Markov chain Monte Carlo (MCMC) sampler of Canova and Pérez Forero (2015) to estimate the TVP-SV-SVARs. Their MCMC sampler is capable of generating posterior distributions of a TVP-SV-SVAR having non-recursive restrictions, which describes several of the alternative identifications. The posterior distributions yield TVPs, SVs, and additional moments to assess the sources and causes of the failure of the Chinese silver standard. The additional moments are month by month tests of uncovered interest parity (UIP), predictability and instability statistics of Cogley, Primiceri, and Sargent (2010) and Cogley and Sargent (2015), and impulse response functions (IRFs).

The estimated TVP-SV-SVARs yield little evidence the U.S. Silver Purchase Act of 1934м06 or intrinsic instability caused the Chinese silver standard to collapse. There are few rejections of UIP, predictability starts to rise in the Great Depression if not earlier, SV and instability peak in 1920 and 1921 and the Great Depression, and the IRFs are nearly unchanged from 1934м01 to 1934м09. This leaves the actions the Nanjing government took beginning in 1934м10 to explain the collapse of the Chinese silver standard.

The outline of the paper follows. Section 2 discusses the Chinese silver standard and data. The SVARs are constructed in section 3. Section 4 describes the TVP-SV-SVARs and Bayesian estimation methods. Results appear in sections 5 and 6. Section 7 concludes.

2. The Chinese Silver Standard and Its Data

This section reviews the Chinese silver standard and the China-U.K. and China-U.S. samples.

2.a. *The Chinese silver standard*

The only role a *tael* had in the Chinese silver standard was as a unit of account. Mediums of exchange in retail trade were copper, brass, and silver coins that had fractional claims on a *tael*; see Leavens (1939, pp. 87). Commercial and financial transactions were settled in local *tael* of which more than 170 existed by the early 1900s; see Dean (2020, p.12).

Of the unit of accounts that existed under the Chinese silver standard, Young (1931), Bratter (1933), Leavens (1939), Brandt, Ma, and Rawski (2014), Jacks, Yan, and Zhao (2017), and Ma (2019) contend the Shanghai *tael*, which dated to 1857, was preeminent by the 1910s. At the time the Republic of China was declared and Qing dynasty collapsed in early 1912, the bulk of China's international trade already flowed through the port of Shanghai. This helped Shanghai to become the hub of the most dynamic economy in China. As a result, the Shanghai *tael* came to dominate domestic and international economic and financial activity in China.

Domestic and international economic activity were supported by banks in Shanghai. They backed their notes and cleared domestic accounts with reserves held in silver ingots of about 50 Shanghai *taels* that were known as shoes of *sycee* (*i.e.*, fine silk); see Leavens (1939, p. 92).[2] Shanghai banks also exported shoes of *sycee* and imported silver denominated in Shanghai *tael* to settle international claims. Nevertheless, settling payments often prompted Shanghai banks, given the state of their balance sheets, to trade reserves in an interbank market operated by the *Shanghai Qian Ye Gong Hui* (*i.e.*, Shanghai Banking Association).[3] Since the Shanghai interbank rate, $i_{S,t}$, cleared this market, it linked $i_{S,t}$ with foreign currency-Shanghai *tael* exchange rates.

The Chinese silver standard operating mechanism differed from the gold standard. One

[2]Converting a *tael* into a domestic price of silver was possible because, as Leavens (1939, pp. 91–95) discusses, weight and fineness (*i.e.*, grains of silver and fractional content of silver) defined a *tael*. In addition, Leavens notes the long time custom of the "Shanghai convention" converting the ideal *tael* into the Shanghai *tael*.

[3]The Nanjing government nationalized the large Shanghai banks in 1935M04, which ended their interbank market.

reason was the world price of silver was set in a spot market in London or New York City (NYC).[4] Another was British pound (*GBP*)- and U.S. dollar (*USD*)-Shanghai *tael* exchange rates floated from 1912M04 to 1934M09.[5] Since the world price of silver defined parity for the Chinese silver standard, deviations from parity were observable, $\rho_t = e_{j/S,t} - SP_t$, where $e_{j/S,t}$ and SP_t denote logs of the *GBP*- or *USD*-Shanghai *tael* exchange rate and world price of silver. Hence, $\rho_t \gtreqless 0$ represented overvaluation, parity, or undervaluation of the Shanghai *tael*.[6] This equilibrium relationship also restricted the dynamic mechanism that restored the Chinese silver standard to parity.[7] Since a deviation from parity of the Chinese silver standard was its risk premium, the U.S. Silver Purchase Act of 1934M06 was a one-off risk premium shock internal to the China-U.S. sample, but was external to the China-U.K. sample.

2.b. The U.S. Silver Purchase Act of 1934

The U.S. Congress passed the Silver Purchase Act of 1934 on June 11. Silver was in effect nationalized in the U.S. by the Silver Purchase Act of 1934. Leuchtenburg (1963, p. 82) discusses that the Act laid out two ways in which this could happen. The U.S Treasury could accumulate silver until it reached 25% of the U.S. monetary base or go to the open market and buy silver until its price was $1.29 per ounce. Kennedy (2005, p. 198) and Silber (2019, p. 47) note advocates of the Act claimed in either case would reinstate or even better the pre-1873 ratio that 16 ounces of silver was worth an ounce of gold. This reflected the policy goals of the Act, which were to reinflate the U.S. economy to undo the deflation of 1930-1933 by returning the

[4]The world silver market was in New York City from 1915M01 to 1934M08. London housed the world silver market from 1912M04 to 1914M12 and in 1934M09.

[5]The U.K. and U.S. operated under different monetary systems (*i.e.*, the gold standard) than China on our sample.

[6]Leavens (1939, p. 102) estimates that just the cost of shipping silver between Shanghai and NYC drove $e_{USD/S,t}$ to vary on average by ±2.5% around SP_t. Shipping costs between Shanghai and London raised this to ±5.0% for the $e_{GBP/S,t}$ after 1920, but Leavens argues it was triple this during the First World War. Jacks, Yan, and Zhao (2017) present econometric evidence that confirm Leaven's estimates.

[7]This differs from the modern floating exchange rate regime in which the real exchange rate serves this purpose as, for example, in the vector error correction model estimated by Engel (2016).

U.S. to a bimetallic standard grounded in silver and gold.[8] Nevertheless, congressional approval to buy silver at above market prices was not clear in the winter of 1933–1934 as recounted by Leuchtenburg (1963, pp. 82–83).

Attempts were made to add silver purchases to the Gold Reserve Act of January 1934. However, Congress had already given President Roosevelt the power to buy silver at above market prices in the Thomas amendment to the Agricultural Adjustment Act of 1933. Leuchtenburg (1963, p. 82) describes President Roosevelt as ordering the Treasury to buy all silver produced in the U.S. at the above market price of \$0.645 $(= \$1.29/2)$ beginning in December 1933. Hence, Roosevelt opposed amendments to the Gold Reserve Act of January 1934 that enlarged on the Thomas amendment. Only when there was majority in the U.S. Senate for the Silver Purchase Act did President Roosevelt agree to back it, according to Leuchtenburg (1963, p. 83).[9] This occurred on May 22, 1934 as told by Silber (2019, p. 62) and Dean (2020, p. 161). Since the Silver Purchase Act of June 1934 passed three weeks later, it leaves little room for news about it to matter for identifying shocks in a SVAR. Moreover, almost two months elapsed until the Treasury nationalized the U.S. silver market on August 9; see Silber (2019, p. 64).

2.c. The China-U.K. and China-U.S. samples: 1912M04–1934M09

The founding of the Republic of China in 1912M01 and fall of the Qing dynasty the next month motivate us to begin the China-U.K. and China-U.S. samples in 1912M04. The samples end in 1934M09. The following month, the Nanjing government severed the link between e_t and SP_t; see Leavens (1939, pp. 300–302) and Ho, Lai, and Gau (2013).

[8]Silver was demonetized in the U.S. by the Coinage Act of 1873. It began the process of moving the U.S. to the gold standard. As a result, supporters of a bimetallic monetary standard came to call this Act the 'Crime of 1873'. Friedman (1990) argues the U.S. would have been better served by a bimetallic standard during the late 1800s. Sargent (2019) shows the gold standard offers a higher level of welfare compared to bimetallism.

[9]Leuchtenburg (1963, p. 83) and Kennedy (2005, p. 198) are clear the Silver Purchase Act of 1934 never achieved its goals of a bimetallic monetary standard and reinflating the U.S. economy. Rather, the Act transferred resources from the federal government to the silver mining industry during the 1930s and 1940s.

Figure 1 plots the China-U.K. and China-U.S. samples, $y_t = [i_t \;\; \pi_t \;\; \rho_t \;\; \Delta e_t]'$, from 1912M04 to 1934M09, $T = 270$. From left to right and top to bottom, the panels contain plots of i_t, π_t, ρ_t, and Δe_t. China-U.K. variables are dot-dashed (red) lines, dotted (blue) lines are China-U.S. variables, Burns and Mitchell (1946) recession dates for the U.K. are vertical tan bars, and NBER recession dates for the U.S. are vertical silver bars. Appendix A1 reviews the ways we compile the data and business cycle dates in figure 1, but summaries of i_t, π_t, ρ_t, and Δe_t follow.

Interest rate spread, i_t**:** Cross-country interest rate spreads are $i_t = i_{S,t} - i_{j,t}$, $j = UK$, US. *Zhongguo ren min yin hang Shanghai Shi fen hang* (1960) has monthly observations for the Shanghai interbank rate, $i_{S,t}$.[10] Monthly money market rates for the U.K., $i_{UK,t}$, and U.S., $i_{US,t}$, are available in the NBER Macrohistory database.

Inflation differential, π_t**:** Shanghai, U.K., and U.S. wholesale price indexes (WPIs) define inflation, $\pi_{m,t} = 100\left(p_{m,t} - p_{m,t-1}\right)$, where $p_{m,t} = \ln WPI_{m,t}$ and $m = S, UK, US$. The Shanghai Research Institute of Economics (1958) supplies monthly $WPI_{S,t}$ starting in 1922. Before 1922, Kong (1988) has an annual WPI for China. We compile a new $WPI_{S,t}$ from 1912M04 to 1934M09 by placing these WPIs on the same 1921 base year, interpolating the former into months, and splicing together these monthly WPIs at 1921M12–1922M01. The NBER Macrohistory database has monthly $WPI_{UK,t}$ and $WPI_{US,t}$. The WPIs yield $\pi_t = \pi_{S,t} - \pi_{j,t}$ for $j = UK, US$.

Deviations from parity, ρ_t**:** Deviations from parity of the Chinese silver standard are calculated as $\rho_t = 100\left(e_{j/S,t} - SP_t\right)$, $j =$ *GBP*, *USD*. Wu (1935) has observations on ρ_t from 1912M04 to 1933M12. Ho and Lai (2016) are tapped for the last nine data points of the samples.

Nominal currency returns, Δe_t**:** We obtain $e_{GBP/S,t}$ and $e_{USD/S,t}$ from Kong (1988). First differencing $e_{j/S,t}$ yields the currency return, $\Delta e_{j/S,t} = 100\left(e_{j/S,t} - e_{j/S,t-1}\right)$, $j =$ *GBP*, *USD*.

[10] Ho and Li (2014) use $i_{S,t}$ to study instability in the Shanghai government bond market of the 1920s and 1930s.

3. An Exchange Rate-Risk Premium SVAR

This section presents the exchange rate-risk premium model of Engel (2016), our baseline SVAR, and broadens it to 10 alternatives.

3.a. An exchange rate model with deviations from Chinese silver standard parity

The exchange rate-risk premium model of Engel (2016) is grounded in a concept of excess currency returns linking deviations from parity for the Chinese silver standard to a risk premium denominated in *tael* earned for holding deposits in *GBP* or *USD*. The risk premium appears in a first-order approximation of currency returns, $\Delta e_{t+1} = i_t + \rho_{t+1}$. The related law of motion of the exchange rate is $e_t = \mathbf{E}_t e_{t+1} - \left(i_t + \mathbf{E}_t\rho_{t+1}\right)$, where $\mathbf{E}_t\rho_{t+1} \neq 0$ violates UIP and $\mathbf{E}_t\{\cdot\}$ is the mathematical expectations operator conditional on date t information. Push the law of motion ahead a period, pass $\mathbf{E}_t\{\cdot\}$ through, replace $\mathbf{E}_t e_{t+1}$, and repeat $\mathcal{J}$ times to find

$$e_t \;=\; \mathbf{E}_t e_{t+J+1} - \sum_{j=0}^{\mathcal{J}} \mathbf{E}_t\left\{i_{t+j} + \rho_{t+j+1}\right\}. \tag{1}$$

Equation (1) shows exchange rate fluctuations are driven by its expectation $\mathcal{J}{+}1$-periods ahead net of the sum of $\mathcal{J}{+}1$ expected returns that are excess, $\mathbf{E}_t\rho_{t+j}$, and otherwise, $\mathbf{E}_t i_{t+j}$.

3.b. Permanent and transitory components of the nominal exchange rate

Engel (2016) decomposes e_t into trend and transitory elements, $\tau_{e,t}$ and $\varepsilon_{e,t}$, using the Beveridge and Nelson (1981) decomposition and equation (1). The Beveridge-Nelson (BN) decomposition requires $\tau_{e,t}$ to be a random walk with drift, $\tau_{e,t} = \tau_e^* + \tau_{e,t-1} + \gamma_{e,\eta}\eta_{e,t}$, where τ_e^* is the deterministic growth rate of e_t and $\eta_{e,t} \sim \mathcal{N}(0, 1)$. Taking $\mathcal{J} \to \infty$ in equation (1) yields

$$e_t \;=\; \tau_{e,t} - \sum_{j=0}^{\infty} \mathbf{E}_t\left\{i_{t+j} + \rho_{t+1+j}\right\}, \tag{2}$$

where the BN trend is $\tau_{e,t} = \lim_{\mathcal{J}\to\infty} \mathbf{E}_t\left\{e_{t+\mathcal{J}+1} - \mathcal{J}\tau_e^*\right\}$. Equation (2) decomposes e_t into $\tau_{e,t}$ and its transitory component, $\varepsilon_{e,t} = -\sum_{j=0}^{\infty}\mathbf{E}_t\left\{i_{t+j} + \rho_{t+1+j}\right\}$, that restricts i_t and $\rho_t \sim I(0)$. We difference equation (2), $\Delta e_t = \gamma_{e,\eta}\eta_{e,t} - \sum_{j=0}^{\infty}\mathbf{E}_t\left\{i_{t+j} + \rho_{t+j+1}\right\} + \sum_{j=0}^{\infty}\mathbf{E}_{t-1}\left\{i_{t+j-1} + \rho_{t+j}\right\}$. Following Nason and Rogers (2008), add and subtract $\mathbf{E}_{t-1}\left\{i_{t+j} + \rho_{t+j+1}\right\}$ inside the brackets of the second infinite sum of the previous expression to obtain

$$\Delta e_t = -\left(i_{t-1} + \mathbf{E}_{t-1}\rho_t\right) + \left[\mathbf{E}_t - \mathbf{E}_{t-1}\right]\varepsilon_{e,t} + \gamma_{e,\eta}\eta_{e,t}, \tag{3}$$

where $\left[\mathbf{E}_t - \mathbf{E}_{t-1}\right]\varepsilon_{e,t} = -\sum_{j=0}^{\infty}\left[\mathbf{E}_t - \mathbf{E}_{t-1}\right]\left\{i_{t+j} + \rho_{t+1+j}\right\}$. Equation (3) links excess currency returns to $\mathbf{E}_{t-1}\rho_t \neq 0$, the forecast innovation of $\varepsilon_{e,t}$, $\left[\mathbf{E}_t - \mathbf{E}_{t-1}\right]\varepsilon_{e,t}$, and the innovation of $\tau_{e,t}$, $\eta_{e,t}$. Hence, moving from the trend-cycle exchange rate decomposition (2) to equation (3) ties Δe_t to observables, forecasts of observables, and a forecast error.

3.c A structural currency return generating regression

We eliminate $\mathbf{E}_{t-1}\rho_t$ and $\left[\mathbf{E}_t - \mathbf{E}_{t-1}\right]\varepsilon_{e,t}$ from equation (3) by assuming the joint probability distribution of $\mathcal{Y}_t$ is a reduced-form VAR, $\mathcal{Y}_t = \sum_{\ell=1}^{k}\mathbf{B}_\ell\mathcal{Y}_{t-\ell} + \lambda_t$, where intercepts are ignored, $\mathbf{B}_\ell$ is a $n\times n$ matrix of lag coefficients, $n = 4$, and $\lambda_t \sim \mathcal{N}\left(\mathbf{0}_{n\times 1}, \mathbf{\Omega}_\lambda\right)$. A singularity is ruled out in the VAR(k) by assuming p_t and the real exchange rate, $q_t \equiv e_t - p_t$, are $I(1)$ implying e_t and p_t do not cointegrate. Invertibility is imposed on the VAR(k) by assuming $\left[\mathbf{I}_n - \mathbf{B}(\mathbf{L})\right]^{-1}$ is square summable, where $\mathbf{B}(\mathbf{L}) = \sum_{\ell=1}^{k}\mathbf{B}_\ell\mathbf{L}^\ell$. Stack k lags of $\mathcal{Y}_t$ in $\mathcal{U}_t = \left[\mathcal{Y}_t' \;\; \mathcal{Y}_{t-1}' \; \ldots \; \mathcal{Y}_{t-k+1}'\right]'$ to find the companion form of the VAR(k), $\mathcal{U}_t = \boldsymbol{\mathcal{B}}\mathcal{U}_{t-1} + \Lambda_t$, where $\boldsymbol{\mathcal{B}}$ is the $nk\times nk$ companion matrix, $\Lambda_t = \left[\lambda_t' \;\; \mathbf{0}_{1\times n(n-1)}\right]'$ for $k > 1$, and $\Lambda_t = \lambda_t$ for $k = 1$. The VAR(1) produces the j-month ahead forecast $\mathbf{E}_t\mathcal{U}_{t+j} = \boldsymbol{\mathcal{B}}^j\mathcal{U}_t$ that yields $\mathbf{E}_{t-1}\rho_t = s_\rho\boldsymbol{\mathcal{B}}\mathcal{U}_t$, $\left(\mathbf{E}_t - \mathbf{E}_{t-1}\right)\mathcal{U}_{t+j} = \boldsymbol{\mathcal{B}}^j\Lambda_t$, and $\left[\mathbf{E}_t - \mathbf{E}_{t-1}\right]\varepsilon_{e,t} = \left[s_i + s_\rho\boldsymbol{\mathcal{B}}\right]\left[\mathbf{I}_{n^2} - \boldsymbol{\mathcal{B}}\right]^{-1}\Lambda_t$ that substituted into equation (3) sets $\Delta e_t = -\left(i_{t-1} + s_\rho\boldsymbol{\mathcal{B}}\mathcal{U}_t\right) + \gamma_{e,\eta}\eta_{e,t} - \left[s_i + s_\rho\boldsymbol{\mathcal{B}}\right]\left[\mathbf{I}_{n^2} - \boldsymbol{\mathcal{B}}\right]^{-1}\Lambda_t$, where s_i (s_ρ) is a $1\times nk$ vector full of zeros except for a one in the first (third) position. Using the companion form of the VAR(k) to eliminate Λ_t in the last formula produces the regression

$$\Delta e_t = a_{\Delta e,i} i_t + a_{\Delta e,\pi} \pi_t + a_{\Delta e,\rho} \rho_t + \sum_{\ell=1}^{k} \mathcal{b}_\ell \mathcal{Y}_{t-\ell} + \gamma_{e,\eta} \eta_{e,t}, \tag{4}$$

where the impact and lag coefficients, $a_{\Delta e,i}$, $a_{\Delta e,\pi}$, $a_{\Delta e,\rho}$, and $\mathcal{b}_1, \ldots, \mathcal{b}_k$, are nonlinear in $\boldsymbol{\mathcal{B}}$. Regression (4) restricts Δe_t. It responds to i_t, π_t, and ρ_t at impact, $\mathcal{b}_1, \ldots, \mathcal{b}_k$ drive persistence, and the innovation to $\tau_{e,t}$, $\eta_{e,t}$, scaled by $\gamma_{e,\eta}$ creates unsystematic variation in Δe_t.

3.d Identifying Exchange Rate-Risk Premium SVARs

The impact coefficients of regression (4) are short-run restrictions that help to identify the SVAR $\mathbf{A}\mathcal{Y}_t = \mathbf{A}\sum_{\ell=1}^{k}\mathbf{B}_\ell \mathcal{Y}_{t-\ell} + \mathbf{\Gamma}\,\eta_t$, where structural shocks have unit variances, $\eta_t \sim \mathcal{N}\left(\mathbf{0}_{n\times 1}, \mathbf{I}_n\right)$, the mapping from the structural to reduced-forms shocks is $\eta_t = \mathbf{\Gamma}^{-1}\mathbf{A}\,\lambda_t$, and $\mathbf{\Gamma}$ is a diagonal matrix containing scale volatilities. We combine $a_{\Delta e,i}$, $a_{\Delta e,\pi}$, and $a_{\Delta e,\rho}$ with zeros imposed on the impact coefficients in the i_t, π_t, and ρ_t regressions to construct the impact matrix

$$\mathbf{A}_{\mathsf{BL}} = \begin{bmatrix} 1 & 0 & 0 & 0 \\ 0 & 1 & 0 & 0 \\ 0 & 0 & 1 & 0 \\ -a_{\Delta e,i} & -a_{\Delta e,\pi} & -a_{\Delta e,\rho} & 1 \end{bmatrix}, \tag{5}$$

of our baseline SVAR, SVAR-BL. Appendix A2 shows $\mathbf{A}_{\mathsf{BL}}$ of equation (5) meets the necessary and sufficiency conditions for global identification of Rubio-Ramírez, Waggoner, and Zha (2010).

There are 510 possible **A**s if up to three of the nine impact coefficients in the i_t, π_t, and ρ_t regressions are non-zero. We limit the model space to SVAR-BL and 10 alternative SVARs that order i_t first followed by π_t, ρ_t, and Δe_t, include $a_{\Delta e,i}$, $a_{\Delta e,\pi}$, and $a_{\Delta e,\rho}$, and are globally identified. The 10 alternative globally identified SVARs have impact matrices $\mathbf{A}_{\mathsf{M1}}, \ldots, \mathbf{A}_{\mathsf{M9}}$, and $\mathbf{A}_{\mathsf{RC}}$, that are listed in table 1. Since $\mathbf{A}_{\mathsf{M1}}$ to $\mathbf{A}_{\mathsf{M4}}$ have at least one impact coefficient above the diagonal, these SVARs are non-recursive. Restrictions on $\mathbf{A}_{\mathsf{M5}}$ to $\mathbf{A}_{\mathsf{M9}}$ have at least one zero below the diagonal and only zeros above the diagonal. No zeros are below the diagonal of $\mathbf{A}_{\mathrm{RC}}$ giving SVAR-RC a recursive identification. The SVARs globally identify international financial, cross-country nominal demand, risk premium, and trend exchange rate shocks.

4. The TVP-SV-SVAR and an MCMC Sampler

We introduce the TVP-SV-SVAR and outline the Metropolis in Gibbs MCMC sampler of Canova and Pérez Forero (2015) in this section.

4.a A TVP-SV-SVAR for the Chinese silver standard

Canova and Pérez Forero (2015) create a TVP-SV-SVAR(k) by endowing $\mathbf{A}$ and $\mathbf{B}_1 \ldots, \mathbf{B}_k$ of the fixed coefficient SVAR with TVPs and the scale volatilities of its diagonal matrix $\boldsymbol{\Gamma}$ with SVs

$$\mathbf{A}_t \mathcal{Y}_t \;=\; \mathbf{A}_t \mathbf{c}_t \,+\, \mathbf{A}_t \sum_{\ell=1}^{k} \mathbf{B}_{\ell,t} \mathcal{Y}_{t-\ell} \,+\, \boldsymbol{\Gamma}_t \eta_t, \quad \eta_t \sim \mathcal{N}\Big(\mathbf{0}_{n\times 1}, \mathbf{I}_n\Big), \tag{6}$$

where $\mathbf{c}_t$ is a $n\times 1$ vector of reduced-form TV intercepts.[11] The TVPs and SVs evolve as multivariate random walks with Gaussian innovations, $a_{t+1} = a_t + \psi_{t+1}$, $\psi_{t+1} \sim \mathcal{N}\Big(\mathbf{0}, \boldsymbol{\Omega}_{\psi}\Big)$, $\mathbb{B}_{t+1} = \mathbb{B}_t + \vartheta_{t+1}$, $\vartheta_{t+1} \sim \mathcal{N}\Big(\mathbf{0}, \boldsymbol{\Omega}_{\vartheta}\Big)$, and $\ln \gamma_{t+1}^2 = \ln \gamma_t^2 + \xi_{t+1}$, $\xi_{t+1} \sim \mathcal{N}\Big(\mathbf{0}, \boldsymbol{\Omega}_{\xi}\Big)$, where a_t is a vector of the off-diagonal TVPs of $\mathbf{A}_t$, $\mathbb{B}_t = \mathrm{vec}\Big(\big[\mathbf{B}_{1,t} \ \ldots \ \mathbf{B}_{k,t} \ \ \mathbf{c}_t\big]\Big)$, and $\mathrm{diag}\Big(\boldsymbol{\Gamma}_t\Big) \equiv \gamma_t = \Big[\gamma_{1,t} \ \ldots \ \gamma_{n,t}\Big]'$. Canova and Pérez Forero (CPF) assume a block diagonal covariance matrix

$$\mathcal{V} \;=\; \begin{bmatrix} \mathbf{I}_n & 0 & 0 & 0 \\ 0 & \boldsymbol{\Omega}_{\vartheta} & 0 & 0 \\ 0 & 0 & \boldsymbol{\Omega}_{\psi} & 0 \\ 0 & 0 & 0 & \boldsymbol{\Omega}_{\xi} \end{bmatrix}, \tag{7}$$

for the structural shocks, η_t, and random walk innovations ϑ_t, ψ_t, and ξ_t.

[11] Implicit is the assumption that the elements of $\mathcal{Y}_t$ are stationary. Appendix A1.6 reports Dickey-Fuller (DF) tests that reject a unit root in i_t, π_t, ρ_t, and Δe_t on the China-U.K. and China-U.S. samples from 1912M04 to 1934M09. On the same samples, DF tests do not reject a unit root in p_t, e_t, and the real exchange rate, q_t.

4.b A Metropolis in Gibbs MCMC sampler for TVP-SV-SVARs

Posterior distributions of the TVP-SV-SVAR(k)s of equation (6), the multivariate random walks of the TVPs and SVs, and covariance matrices of equation (7) are assembled using the Metropolis in Gibbs MCMC sampler of CPF (2015) and our priors. Their MCMC sampler solves the problem of sampling from the posterior of a TVP-SV-SVAR that has a nonlinear likelihood induced by a non-recursive identification.

Implementing the CPF-MCMC sampler is an algorithm organized around the correction to Primiceri (2005) devised by Del Negro and Primiceri (2015). The corrected sampler runs for $\mathcal{D}$ steps. Its dth iteration begins with a Gibbs step to draw $\mathbb{B}_{1:T}$ using the Carter and Kohn (1994) multi-move scheme constrained by the Koop and Potter (2011) rule to toss out $\mathbb{B}_{1:T}$ if any $\mathbb{B}_t$ is explosive. Next, draw $a_{1:T}$ in a Metropolis step. This is followed with a Gibbs step that draws $\gamma_{1:T}$ employing the 10-point mixture normal distribution of Omori, Chib, Shephard, and Nakajima (2007). The last step of iteration d draws $\mathbf{\Omega}_{\vartheta}$, $\mathbf{\Omega}_{\psi}$, and $\mathbf{\Omega}_{\xi}$. Our priors are set to attain acceptance rates of 50 to 60% for non-explosive draws from the posterior distribution of $\mathbb{B}_{1:T}$ and about 32% of the draws from the posterior distribution of $a_{1:T}$. Appendix A3 describes our priors and the CPF-Metropolis in Gibbs MCMC sampler.

We take $\mathcal{D}$ = 500,000 draws from posterior distributions of the TVP-SV-SVAR(k)s after $0.5\mathcal{D}$ burn-in steps, given $\mathcal{Y}_{CUK,1:T}$ $\left(\text{or } \mathcal{Y}_{CUS,1:T}\right)$, our priors, and $k = 2$. Posterior distributions are thinned to $0.005\mathcal{D}$ by random sampling without replacement.

5. Posterior Moments of the TVP-SV-SVARs

This section assesses which of the 11 TVP-SV-SVAR(2)s are favored by the China-U.K. and China-U.S. samples and reports the posterior distributions of these TVP-SV-SVAR(2)s.

5.a. Evaluating the TVP-SV-SVARs on the China-U.K. and China-U.S. samples

We evaluate the TVP-SV-SVAR(2)s using marginal data densities (MDDs) and the widely applicable information criterion (WAIC) of Watanabe (2010). The MDDs are informative about which TVP-SV-SVAR(2) is preferred by $y_{CUK,1:T}$ or $y_{CUS,1:T}$. Geweke (2005) is our guide to computing the MDDs. The WAIC compliments the MDD by finding the TVP-SV-SVAR(2) that has the smallest posterior 1-month ahead predictive loss, given $y_{CUK,1:T}$ or $y_{CUS,1:T}$ and our priors. Gelman, et al (2014) advise calculating the WAIC as the negative of twice the posterior mean of the log predictive likelihood minus its variance. The variance of the posterior mean of the log predictive likelihood acts as a penalty term in the WAIC.

The bold entries in the ln MDD and WAIC columns of table 2 give decisive evidence that TVP-SV-SVAR(2)-BL offers the best fit and forecast (*i.e.*, largest ln MDD and smallest WAIC) on $y_{CUK,1:T}$ and $y_{CUS,1:T}$. The evidence is the samples prefer $\mathbf{A}_{\mathsf{BL}}$ of equation (5) that restricts i_t, π_t, and ρ_t to be causally prior to one another and to Δe_t. Hence, the international financial shock affects currency returns through i_t at impact. The same holds for the impact responses of Δe_t to the π_t-cross-country nominal demand shock and ρ_t-risk premium shock pairs. Since $y_{CUK,1:T}$ and $y_{CUS,1:T}$ prefer TVP-SV-SVAR(2)-BL, the rest of the paper relies on estimates of this SVAR to study the Chinese silver standard.

5.b. Time-varying impact coefficients

Figure 2 displays moments of $a_{\Delta e,j,1:T,CUK}$ $\left(a_{\Delta e,j,1:T,CUS}\right)$ in the top (bottom) row taken from posterior distributions of TVP-SV-SVAR(2)-BL, $j = i$, π and ρ. Solid lines are posterior medians of $a_{\Delta e,j,1:T,CUK}$ and $a_{\Delta e,j,1:T,CUS}$ that are shaded by 68% Bayesian credible sets (*i.e.*, 16% and 84% quantiles). Burns and Mitchell (1946) monthly U.K. recessions are vertical (tan) bars in the top row of figure 2. Its bottom row has vertical (gray) bars that are NBER recession dates.

The bottom row of figure 2 has posterior medians of $a_{\Delta e,i,1:T,CUS}$, $a_{\Delta e,\pi,1:T,CUS}$, and $a_{\Delta e,\rho,1:T,CUS}$ that exhibit co-movement with every NBER recession from 1912M04 to 1934M09. However, only the posterior medians of $a_{\Delta e,i,1:T,CUK}$ take a path that resembles the plots in the top row of panels of figure 2. Still, the top right panel shows only two troughs and one peak that occur during U.K. recessions dated by Burns and Mitchell (BM). These are at the start of the 1912M12–1914M09, end of the 1918M10–1919M04, and middle of the 1929M07–1932M08 (*i.e.*, the U.K.'s Great Depression) recessions. Posterior medians of $a_{\Delta e,\pi,1:T,CUK}$ and $a_{\Delta e,\rho,1:T,CUK}$ peak during the U.K.'s Great Depression and 1920M03–1921M06 recession.

Mean reversion is evident in the posterior medians of $a_{\Delta e,i,1:T,CUS}$, $a_{\Delta e,\pi,1:T,CUS}$, and $a_{\Delta e,\rho 1:T,CUS}$ in the bottom row of figure 2. Posterior medians of $a_{\Delta e,\pi,1:T,CUS}$ fluctuate around one in the lower middle panel of the figure. The 68% Bayesian credible sets in the panel cover one in 190 of the 270 months of the sample suggesting relative purchasing power parity (PPP) is often not rejected for $\Delta e_{USD/S,t}$ and $\pi_{S,t} - \pi_{US,t}$ under the Chinese silver standard. The lower right panel of figure 2 shows posterior medians of $a_{\Delta e,\rho 1:T,CUS}$ moving around zero indicating the Chinese silver standard was restored to parity as frequently as the six NBER recessions occur in the sample. There is also mean reversion in the posterior medians of $a_{\Delta e,i,1:T,CUS}$ in the bottom left panel of figure 2, but $a_{\Delta e,i,1:T,CUS} < 0$. The implication is $i_{S,t} > i_{US,t}$ led the Shanghai *tael* to appreciate suggesting it carried less risk than the *USD* during the sample.

Posterior distributions of $a_{\Delta e,i,1:T,CUK}$, $a_{\Delta e,\pi,1:T,CUK}$, and $a_{\Delta e,\rho,1:T,CUK}$ yield different inferences. First, the top left panel of figure 2 shows posterior medians of $a_{\Delta e,i,1:T,CUK} > 0$. This implies the Shanghai *tael* was riskier than the *GBP* when $e_{GBP/S,t}$ depreciated in response to $i_{S,t} > i_{UK,t}$. Next, there is less evidence supporting relative PPP for $\Delta e_{GBP/S,t}$ and $\pi_{S,t} - \pi_{UK,t}$ because of the greater variability in the posterior medians of $a_{\Delta e,\pi,1:T,CUK}$ around one and the width of the 68% Bayesian credible sets compared with the top middle panel. Lastly, the top right panel shows posterior medians of $a_{\Delta e,\rho,1:T,CUK}$ predict a one percent change in ρ_t generated a 40 to 60 basis point increase in $\Delta e_{GBP/\$,t}$. These responses contrast with the mean reversion to parity observed in the bottom right panel of figure 2.

5.c. The Chinese silver standard: Volatility

Figure 3 depicts posterior medians of the SVs with solid lines that are covered by 68% Bayesian credible sets. These moments of the posterior distributions of TVP-SV-SVAR-BL are conditional on $y_{CUK,1:T}$ $\left(y_{CUS,1:T}\right)$ in the top (bottom) row of figure 3. From left to right, its columns contain posterior moments of $\gamma_{i,1:T,j}$, $\gamma_{\pi,1:T,j}$, $\gamma_{\rho,1:T,j}$, and $\gamma_{\Delta e,1:T,}$, $j = CUK, CUS$.

Three of four posterior median SVs peak during BM dated U.K. recessions in the top row of figure 3. The 1920M03–1921M06 recession has peaks in $\gamma_{\pi,t,CUK}$, $\gamma_{\rho,t,CUK}$, and $\gamma_{\Delta e,t,CUK}$, but $\gamma_{i,t,CUK}$ peaks in 1921M10. Secondary peaks are observed in $\gamma_{i,t,CUK}$ between the 1920M03–1921M06 and 1924M11–1926M07 recessions, during the 1927M03–1928M09 recession for $\gamma_{\rho,t,CUK}$, and for $\gamma_{\pi,t,CUK}$ and $\gamma_{\Delta e,t,CUK}$ in the Great Depression.

The bottom row of figure 3 displays peaks in the posterior median SVs during NBER recessions. Two peak in the 1920M01–1921M07 recession $\left(\gamma_{i,t,CUS} \text{ and } \gamma_{\pi,t,CUS}\right)$. Another peak coincides with the 1918M08–1919M03 recession $\left(\gamma_{\rho,t,CUS}\right)$ and one in the Great Depression $\left(\gamma_{\Delta e,t,CUS}\right)$. There are secondary peaks in the First World War $\left(\gamma_{\pi,t,CUS} \text{ and } \gamma_{\Delta e,t,CUS}\right)$ and the 1923M05–1924M07 $\left(\gamma_{i,t,CUS}\right)$ and 1926M10–1927M11 $\left(\gamma_{\rho,t,CUS}\right)$ recessions.

Figure 3 also shows posterior medians of the SVs are falling at the end of the sample except for $\gamma_{\rho,1:T,CUK}$ and $\gamma_{\rho,1:T,CUS}$. These medians take a W-shaped path from a peak in 1920M06 and 1918M09 to 1934M09, but substantial disparities exist in the height of these peak SVs compared with observations from 1934M06 to 1934M09. Posterior medians during these months are 50 to 60% less than the peaks of $\gamma_{\rho,1:T,CUK}$ at 1920M06 and $\gamma_{\rho,1:T,CUS}$ at 1918M09.

6. Time-varying UIP, Predictability, and (In)Stability

This section employs monthly tests of UIP, predictability and instability statistics, and IRFs to study the efficiency and (in)stability of the Chinese silver standard.

6.a The Chinese silver standard; Time-varying tests of UIP

Rejections of UIP on the Chinese silver standard only appear in the China-U.K. sample from spring 1918 into the 1920s. The evidence is obtained by adapting methods of Hodrick (1992) to test the Chinese silver standard for violations of UIP. His formulas and posterior distribution of a TVP-SV-SVAR are engaged to compute the TV-slope coefficient of the Fama regression, $\Delta e_{t+1} = \delta_{0,t} + \delta_{1,t}\left(i_{S,t} - i_{j,t}\right) + \zeta_{\Delta e,t+1}$. We also report estimates of the TV-slope coefficient of the real exchange rate cousin of the Fama regression, $\Delta q_{t+1} = \varrho_{0,t} + \varrho_{1,t}\left(r_{j,t} - r_{S,t}\right) + \zeta_{\Delta q,t+1}$, that is constructed by substituting ρ_{t+1} for the dependent variable of the Fama regression to turn its slope coefficient into $\delta_{1,t}-1$ and using $\Delta e_{t+1} - \pi_{t+1} = r_t + \rho_{t+1} - \left(\pi_{t+1} - \mathbf{E}_t\pi_{t+1}\right)$, where the ex ante real rate spread $r_t = i_t - \mathbf{E}_t\pi_{t+1}$, $r_t = r_{S,t} - r_{j,t}$, and $j = UK, US$. Our interest is in the null hypotheses of UIP, $\delta_{1,t} = 1$ and $\varrho_{1,t} = 1$, at $t = 1, \ldots, T$.

Engel (2016) stresses the signs of $\text{cov}\left(\mathbf{E}_t\rho_{t+1}, r_{j,t} - r_{S,t}\right)$ and $\text{cov}\left(\mathbf{E}_t\sum_{j=0}^{\infty}\rho_{t+j+1}, r_{j,t} - r_{S,t}\right)$ are useful for studying UIP. Rejection rests on either covariance not equaling zero, but the $\text{cov}\left(\mathbf{E}_t\rho_{t+1}, r_{j,t} - r_{S,t}\right) > 0$ and $\text{cov}\left(\mathbf{E}_t\sum_{j=0}^{\infty}\rho_{t+j+1}, r_{j,t} - r_{S,t}\right) < 0$ summarize Engel's exchange rate-risk premium paradox. The first inequality signals excess sensitivity of Δe_{t+1} to a change in i_t. This points to greater risk on the Shanghai *tael* compared with the *GBP* or *USD*. The latter inequality indicates excess volatility in e_t implying a stronger Shanghai *tael* carried less risk. We approximate the sign of $\text{cov}_t\left(\mathbf{E}_t\sum_{j=0}^{\infty}\rho_{t+j+1}, r_{j,t} - r_{S,t}\right)$ with the TV-slope coefficient,

$\phi_{H,t}$, of the long-horizon regression $\rho_{t+1} = \phi_{0,t} + \phi_{H,t} \sum_{h=0}^{H-1} \left(r_{j,t-h} - r_{S,t-h} \right) + \zeta_{\rho,t+1}$ and replicate it as $H \to \infty$. When $H = 1$, the sign of $\text{cov}_t \left(\mathbf{E}_t \rho_{t+1},\, r_{j,t} - r_{S,t} \right)$ is recovered from $\phi_{H,t}$. Formulas in Hodrick (1992) are used to compute $\phi_{h,t}$, $h = 1, \ldots, H$, given a TVP-SV-SVAR. Appendix A4 has details about these constructing these regressions.

Figure 4 collects posterior moments of $\delta_{1:T,\,j}$, $\varrho_{1:T,\,j}$, $\phi_{1,1:T,\,j}$, and $\phi_{3,1:T,\,j}$ for $j = CUK$ in the top and $j = CUS$ in the bottom row in panels from left to right. The posterior distributions of TVP-SV-SVAR(2)-BL conditional on $\mathcal{Y}_{1:T,CUK}$ $\left(\mathcal{Y}_{1:T,CUS}\right)$ and our priors yield medians (solid lines) shaded by 68% Bayesian credible sets in the top (lower) row of figure 4. Since posterior medians of $\phi_{H,1:T} \approx 0$ at $H > 3$, we report long-horizon regressions at $H = 3$.

Rejections of UIP appear in the upper middle two panels of figure 4, but its lower middle panels show this is not true for plots of $\varrho_{1:T,\,CUS}$ and $\phi_{3,1:T,\,CUS}$. The 68% Bayesian credible sets of $\varrho_{1918\text{M}12:1920\text{M}06,\,CUK}$ exclude one, which rejects UIP. These dates run from the month after the Armistice that ended the First World War to two months after the Bank of England rate hit 7%, which Howson (1974) argues ended the U.K.'s post-war boom. More evidence against UIP are 68% Bayesian credible sets of $\phi_{1,1918\text{M}07:1927\text{M}11,\,CUK}$ that are bounded above zero. At the 1-month horizon, investors anticipated excess returns on deposits in the *GBP* relative to the Shanghai *tael* from two months after a new Governor began at the Bank of England to the middle of the BM dated recession of 1927M03–1928M09.

The other four panels of figure 4 do not reject UIP. The 68% Bayesian credible sets of $\delta_{1:T,\,CUK}$ and $\delta_{1:T,\,CUS}$ $\left(\phi_{3,1:T,\,CUK}\right.$ and $\left.\phi_{3,1:T,\,CUS}\right)$ cover one (zero) in the first (last) column of figure 4. However, the exchange rate-risk premium paradox of Engel (2016) is satisfied by the posterior medians of $\phi_{1,1:T,\,CUK} > 0$ and $\phi_{3,1917\text{M}03:T,\,CUK} < 0$ in the top row of figure 4. The row below shows the posterior medians of $\phi_{1,1:T,\,CUS}$ and $\phi_{3,1:T,\,CUS}$ reverse the signs.

6.b. *The Chinese Silver Standard: Predictability of deviations from trend of p_t and e_t*

Cogley, Primiceri, and Sargent (2010) propose a h-step ahead TV-predictability statistic, $\mathcal{R}^2_{z,h,t}$, for any $z_t \sim I(1)$ that is one minus the ratio of its conditional to its unconditional variance. As they do, we invoke the anticipated utility model (AUM) of Kreps (1998) to address the problem of forecasting with TVPs and SVs. This makes $\mathcal{R}^2_{z,h,t}$ an approximation. Nevertheless, $\mathcal{R}^2_{z,h,t}$ measures the h-step ahead predictability of deviations from the trend of z_t. If the deviations are unpredictable h-steps ahead at date t, $\mathcal{R}^2_{z,h,t} = 0$. Also, $\lim_{h\to\infty} \mathcal{R}^2_{z,h,t} = 0$. Predictability places $\mathcal{R}^2_{z,h,t} \in (0, 1)$. When $\mathcal{R}^2_{z,h,t}$ is rising, it is evidence of decreasing stability in z_t at the h-step ahead horizon. Appendix A5 reviews our approach to computing $\mathcal{R}^2_{z,h,t}$.

Figure 5 plots posterior medians (solid lines) and 68% Bayesian credible sets of $\mathcal{R}^2_{p,1,1:T}$, $\mathcal{R}^2_{p,6,1:T}$, $\mathcal{R}^2_{e,1,1:T}$ and $\mathcal{R}^2_{e,6,1:T}$ from 1912M04 to 1934M09 in columns from left to right. The top (bottom) row displays moments taken from the posterior distribution of TVP-SV-SVAR(2)-BL conditional on $y_{1:T,CUK}$ $\left(y_{1:T,CUS}\right)$ and our priors.

Predictability of deviations from trend is greatest in the left column of panels of figure 5 for the posterior medians of $\mathcal{R}^2_{p,1,1:T}$. Comparing posterior medians of $\mathcal{R}^2_{p,1,t}$ shows the top left panel offers greater predictability pre-1920M07 than after. The top and bottom left panels also show posterior medians of $\mathcal{R}^2_{p,1,t}$ rising from 1932M09 and 1927M01 to 1934M09. Although the U.K.'s Great Depression recession ended a month after 1932M09 and the Nanjing government obtained control of the Shanghai Customs Office in 1927M01, these dates are two to seven and a half years before the U.S. Silver Purchase Act of 1934M06.

The last three columns of figure 5 display posterior moments of $\mathcal{R}^2_{p,6,1:T}$, $\mathcal{R}^2_{e,1,1:T}$, and $\mathcal{R}^2_{e,6,1:T}$. The 68% Bayesian credible sets of $\mathcal{R}^2_{p,6,t}$ are well above zero by no later than the end of the First World War in the second column of panels. The third column shows substantial predictability of deviations from trend at the 1-month horizon for $e_{GBP/S,t}$ and $e_{USD/S,t}$ from 1912M04 to 1934M09. However, the right most panels of figure 5 display 68% Bayesian credible sets with lower quantiles ranging from zero to 0.008 that indicate predictability of deviations from trend for $e_{GBP/S,1:T}$ and $e_{USD/S,1:T}$ vanishes by the 6-month horizon.

6.c. The Chinese Silver Standard: (In)Stability of p_t and e_t

We gauge instability in the Chinese silver standard with a statistic developed by Cogley and Sargent (2015). They calculate instability in the $I(1)$ variable z_t as the square root of the sum of two components. The first is the forward-looking uncertainty around $\mathbf{E}_t\Delta z_{t+h}$, which is the conditional variance of this forecast, $\mathcal{V}_t\left(z_{t+h} - \mathbf{E}_t z_{t+h}\right)$. Add to this the variance of the h-step ahead expected accumulated growth of z_t, $\mathbf{E}_t z_{t+h} - z_t$. The result is the Cogley and Sargent measure of instability $\sigma_{z,h,t} \approx \sqrt{\mathcal{V}_t\left(z_{t+h} - \mathbf{E}_t z_{t+h}\right) + \left(\mathbf{E}_t z_{t+h} - z_t\right)^2}$; see their equation (8). We appeal to the AUM, as in Cogley and Sargent (2015), to compute $\sigma_{z,h,t}$ on the posterior distributions of TVP-SV-SVAR(2)-BL. Appendix A7 has details about adapting Cogley and Sargent (2015) to calculate $\sigma_{z,h,t}$ on a TVP-SV-SVAR.

Figure 6 displays posterior distributions of $\sigma_{p,h,1:T}$ and $\sigma_{e,h,1:T}$ at 1- and 12-month horizons in columns from left to right on the 1912M04–1934M09 sample. The 68% Bayesian credible sets cover solid lines that are posterior medians. Distributions in the top (bottom) row depend on TVP-SV-SVAR(2)-BL, $y_{CUK,1:T}$ $\left(y_{CUS,1:T}\right)$ and our priors.

The Chinese silver standard saw instability peak during the recession of the early 1920s and Great Depression as shown in the top and bottom rows of figure 6. In the top row, posterior medians of $\sigma_{p,1,t}$, $\sigma_{p,12,t}$, $\sigma_{e,1,t}$, and $\sigma_{e,12,t}$ peak at 1921M02, 1920M11, 1920M12, and 1920M12, respectively. The same row has secondary peaks in the posterior medians of these instability statistics in 1931M10 remembering the U.K. left the gold standard and Japan invaded Manchuria the previous month. In contrast, the bottom row of figure 6 shows posterior medians of $\sigma_{p,1,t}$, $\sigma_{p,12,t}$, $\sigma_{e,1,t}$, and $\sigma_{e,12,t}$ peaking during the NBER recession of 1920M01–1921M07 and the Great Depression at 1920M10, 1931M05, 1931M05, and 1931M06. Secondary peaks occur at 1931M05, 1920M10, 1920M04, and 1920M04 in the same row of the figure.

Posterior distributions of $\sigma_{p,1,t}$, $\sigma_{p,12,t}$, $\sigma_{e,1,t}$, and $\sigma_{e,12,t}$ shift down after the Great Depression both rows of figure 6. The upshot is that, although 68% Bayesian credible sets are wider around posterior medians at the 12-month horizon in 1933 and 1934, figure 6 shows the posterior medians are lower after the Great Depression.

6.d. The Chinese Silver Standard: TV-IRFs in 1934

Figures 7 and 8 display posterior median IRFs at 1934M01, 1934M06, and 1934M09 created using the TVP-SV-SVAR(2)-BL, given $\mathcal{Y}_{1:T,CUK}$ or $\mathcal{Y}_{1:T,CUS}$ and our priors. The dates are six months before the U.S. Silver Purchase Act became law, its month of passage, and end of the samples. The top row depicts IRFs of i_t with respect to international financial, cross-country nominal demand, risk premium, and trend exchange rate shocks from impact to the 12-month horizon. The IRFs of p_t, ρ_t, and e_t on the same shocks appear in the next three rows in figures 7 and 8. We accumulate IRFs of π_t and Δe_t to obtain IRFs of p_t and e_t. Appendix A8 reports on the complete set of TV-IRFs.

The structural shocks produce median posterior IRFs of i_t and ρ_t at 1934M01, 1934M06, and 1934M09 that are mean reverting in figures 7 and 8. The top row of these figures have posterior median IRFs of i_t that return to zero by five months after an international financial, a cross-country nominal demand, a risk premium, or a trend exchange rate shock. Similarly, the posterior median IRFs of ρ_t revert to zero by the 5-month horizon. Hence, deviations from parity of the Chinese silver standard lacked persistence in the first nine months of 1934.

The four shocks produce posterior median IRFs of p_t and e_t at 1934M01, 1934M06, and 1934M09 in figures 7 and 8 that often predict $p_{S,t}$ increasing faster than $p_{UK,t}$ or $p_{US,t}$ and depreciation of $e_{GBP/S,t}$ and $e_{USD/S,t}$. The depreciation predicted for the Shanghai *tael* is found in the positive hump-shaped posterior median IRFs to the cross-country nominal demand, risk premium, and trend exchange rate shocks in the bottom row of the figures. The international financial shock also produces hump-shaped posterior median IRFs of e_t in the bottom left panels of figures 7 and 8. However, the latter panel displays IRFs that are negative, which shows the international financial shock generated an appreciation of $e_{USD/S,t}$.

Lastly, we garner evidence about the stability of the Chinese silver standard from 1934M01 to 1934M09 by comparing posterior median IRFs in figure 7 and figure 8. Going panel by panel finds posterior median IRFs at 1934M01, 1934M06, and 1934M09 are often identical in height, shape, and persistence. Hence, figures 7 and 8 give additional evidence the U.S. Silver Purchase Act of 1934M06 had little impact on the Chinese silver standard before 1934M10.

7. Conclusion

This paper studies whether the Chinese silver standard failed because its operating mechanism was fragile or the U.S. Silver Purchase Act of 1934M06 drained China of silver. The hypotheses are examined using the framework of the exchange rate-risk premium model of Engel (2016). Starting from his model, we build a baseline and 10 alternative structural VARs (SVARs). The SVARs include time varying parameters (TVPs) and stochastic volatility (SV). Bayesian methods are used to estimate the TVP-SV-SVARs on newly available China-U.K. and China-U.S. samples from 1912M04 to 1934M09. The estimates let us to test for uncovered interest parity (UIP) under the Chinese silver standard and its predictability and instability month by month.

Our results show the Chinese silver standard was resilient as it weathered the turmoil China, the U.K., and U.S. faced from 1912 to 1934. Rejections of UIP are only at the 1-month horizon for the *GBP*-Shanghai *tael* exchange rate from the end of the First World War into the 1920s. Predictability in deviations from the trends of the China-U.K. and China-U.S. WPI differentials begin to rise two years or more before the U.S. Silver Purchase Act of 1934M06. Episodes of instability in the WPI differentials and *GBP*- and *USD*-Shanghai *tael* exchange rates are focused on the recession of the early 1920s and Great Depression. Similarly, impulse response functions change little in the 1930s. These results are testimony the U.S. Silver Purchase Act of 1934M06 had little impact on the Chinese silver standard. We conclude the Chinese silver standard did not fail because of its own fragility or the actions of the U.S. Congress.

This leaves actions the Nanjing government took starting in 1934M10 to explain the demise of the Chinese silver standard. Brandt and Sargent (1989) argue the Chinese silver standard was disrupted by the Nanjing government pursuing goals that caused private agents to revise their expectations about the monetary regime in China. After 1934M10, changes in these expectations led to a growing exodus of silver from China that, according to Brandt and Sargent, contributed to the Nanjing government replacing the Chinese silver standard with the *fabi* in 1935M11. It is beyond the scope of this paper to assess the account Brandt and Sargent offer to explain the end of the Chinese silver standard. Nonetheless, along with stimulating research on the Chinese silver standard, we hope this paper spurs work on the consequences of policy making that leads to unforeseen revisions to private sector expectations.

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TABLE 1. IMPACT MATRICES OF THE ALTERNATIVE GLOBALLY IDENTIFIED SVARS

$$\mathbf{A}_{\mathtt{M1}} = \begin{bmatrix} 1 & 0 & 0 & 0 \\ 0 & 1 & 0 & -a_{\pi,\Delta e} \\ 0 & 0 & 1 & -a_{\rho,\Delta e} \\ -a_{\Delta e,i} & -a_{\Delta e,\pi} & -a_{\Delta e,\rho} & 1 \end{bmatrix} \quad \mathbf{A}_{\mathtt{M2}} = \begin{bmatrix} 1 & 0 & 0 & 0 \\ 0 & 1 & -a_{\pi,\rho} & 0 \\ 0 & 0 & 1 & -a_{\rho,\Delta e} \\ -a_{\Delta e,i} & -a_{\Delta e,\pi} & -a_{\Delta e,\rho} & 1 \end{bmatrix}$$

$$\mathbf{A}_{\mathtt{M3}} = \begin{bmatrix} 1 & 0 & 0 & 0 \\ 0 & 1 & 0 & -a_{\pi,\Delta e} \\ 0 & -a_{\rho,\pi} & 1 & 0 \\ -a_{\Delta e,i} & -a_{\Delta e,\pi} & -a_{\Delta e,\rho} & 1 \end{bmatrix} \quad \mathbf{A}_{\mathtt{M4}} = \begin{bmatrix} 1 & 0 & 0 & 0 \\ 0 & 1 & -a_{\pi,\rho} & 0 \\ -a_{\rho,i} & 0 & 1 & 0 \\ -a_{\Delta e,i} & -a_{\Delta e,\pi} & -a_{\Delta e,\rho} & 1 \end{bmatrix}$$

$$\mathbf{A}_{\mathtt{M5}} = \begin{bmatrix} 1 & 0 & 0 & 0 \\ 0 & 1 & 0 & 0 \\ 0 & -a_{\rho,\pi} & 1 & 0 \\ -a_{\Delta e,i} & -a_{\Delta e,\pi} & -a_{\Delta e,\rho} & 1 \end{bmatrix} \quad \mathbf{A}_{\mathtt{M6}} = \begin{bmatrix} 1 & 0 & 0 & 0 \\ 0 & 1 & 0 & 0 \\ -a_{\rho,i} & 0 & 1 & 0 \\ -a_{\Delta e,i} & -a_{\Delta e,\pi} & -a_{\Delta e,\rho} & 1 \end{bmatrix}$$

$$\mathbf{A}_{\mathtt{M7}} = \begin{bmatrix} 1 & 0 & 0 & 0 \\ -a_{\pi,i} & 1 & 0 & 0 \\ 0 & 0 & 1 & 0 \\ -a_{\Delta e,i} & -a_{\Delta e,\pi} & -a_{\Delta e,\rho} & 1 \end{bmatrix} \quad \mathbf{A}_{\mathtt{M8}} = \begin{bmatrix} 1 & 0 & 0 & 0 \\ 0 & 1 & 0 & 0 \\ -a_{\rho,i} & -a_{\rho,\pi} & 1 & 0 \\ -a_{\Delta e,i} & -a_{\Delta e,\pi} & -a_{\Delta e,\rho} & 1 \end{bmatrix}$$

$$\mathbf{A}_{\mathtt{M9}} = \begin{bmatrix} 1 & 0 & 0 & 0 \\ -a_{\pi,i} & 1 & 0 & 0 \\ 0 & -a_{\rho,\pi} & 1 & 0 \\ -a_{\Delta e,i} & -a_{\Delta e,\pi} & -a_{\Delta e,\rho} & 1 \end{bmatrix} \quad \mathbf{A}_{\mathtt{RC}} = \begin{bmatrix} 1 & 0 & 0 & 0 \\ -a_{\pi,i} & 1 & 0 & 0 \\ -a_{\rho,i} & -a_{\rho,\pi} & 1 & 0 \\ -a_{\Delta e,i} & -a_{\Delta e,\pi} & -a_{\Delta e,\rho} & 1 \end{bmatrix}$$

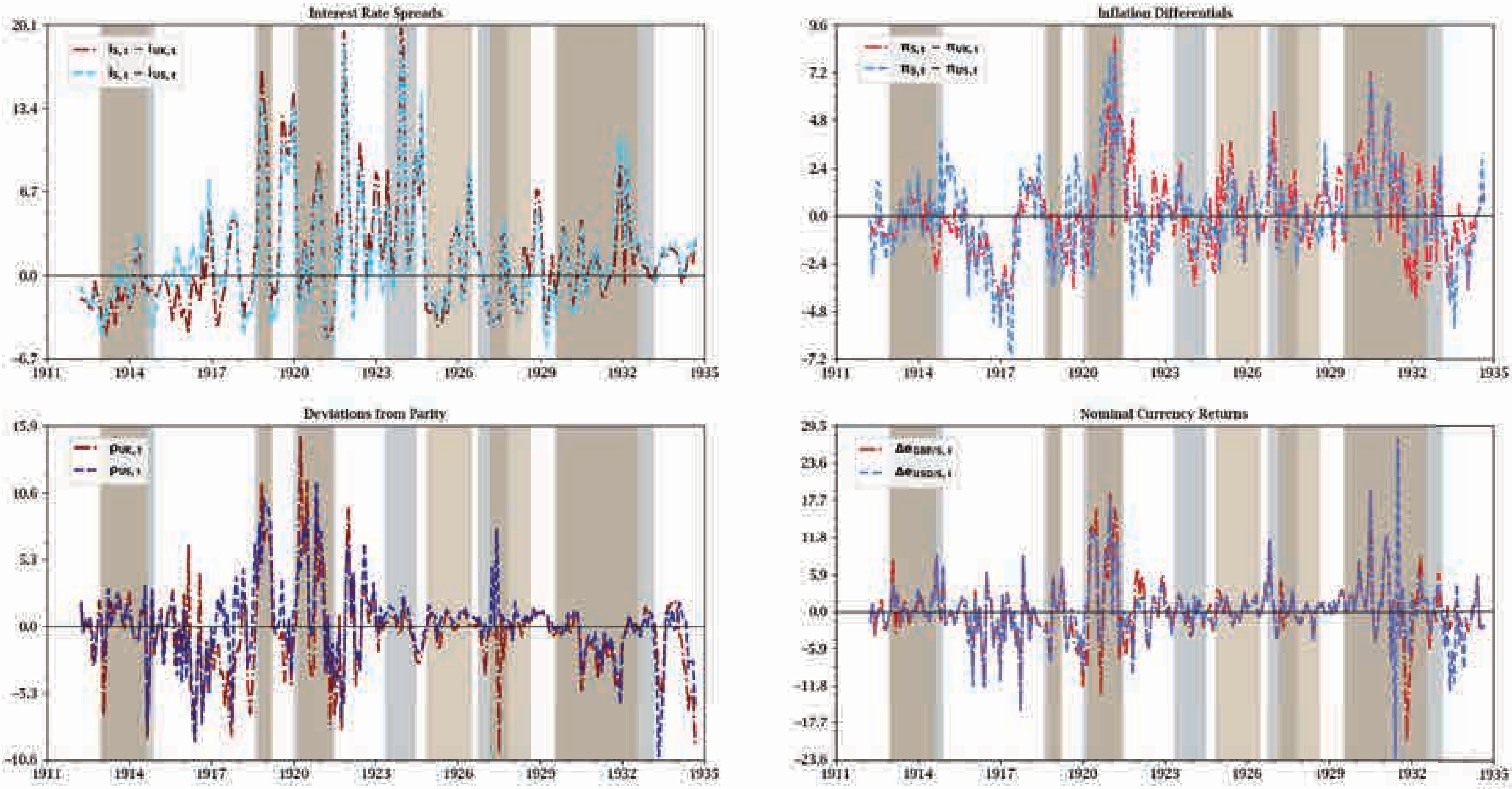
Notes: Global identification is verified using criteria developed by Rubio-Ramírez, Waggoner, and Zha (2010). The (d, j) element of $\mathbf{A}_\ell$ is the impact coefficient a_{dj}, where $d \neq j$ and $\ell = \mathtt{M1}, \ldots, \mathtt{M9}$, and RC, where Mx denotes SVAR x $= 1, \ldots, 9$, and RC is the recursive SVAR. Equation (5) has the impact matrix of our baseline SVAR, SVAR-BL.

Table 2. MDDs and WAICs of the Baseline and Alternative Globally Identified SVARs on the China-U.K. and China-U.S. Samples, 1912M04–1934M09

	$y_{CUK,1:T}$		$y_{CUS,1:T}$	
SVAR	ln MDD	WAIC	ln MDD	WAIC
BL	**−2049.45**	**5838.59**	**−2115.40**	**5853.55**
M1	−2564.19	6150.10	−2826.01	6067.70
M2	−2508.43	6088.25	−2716.09	6180.55
M3	−2151.04	5985.73	−2331.61	5941.21
M4	−2247.96	5908.49	−2426.98	5923.75
M5	−2257.41	5847.40	−2244.37	5889.24
M6	−2185.39	5905.78	−2170.32	5883.31
M7	−2173.85	5840.45	−2363.15	5940.85
M8	−2167.46	5867.85	−2221.30	5868.09
M9	−2249.30	5858.85	−2224.19	5876.26
RC	−2421.83	5967.97	−2542.05	6007.94

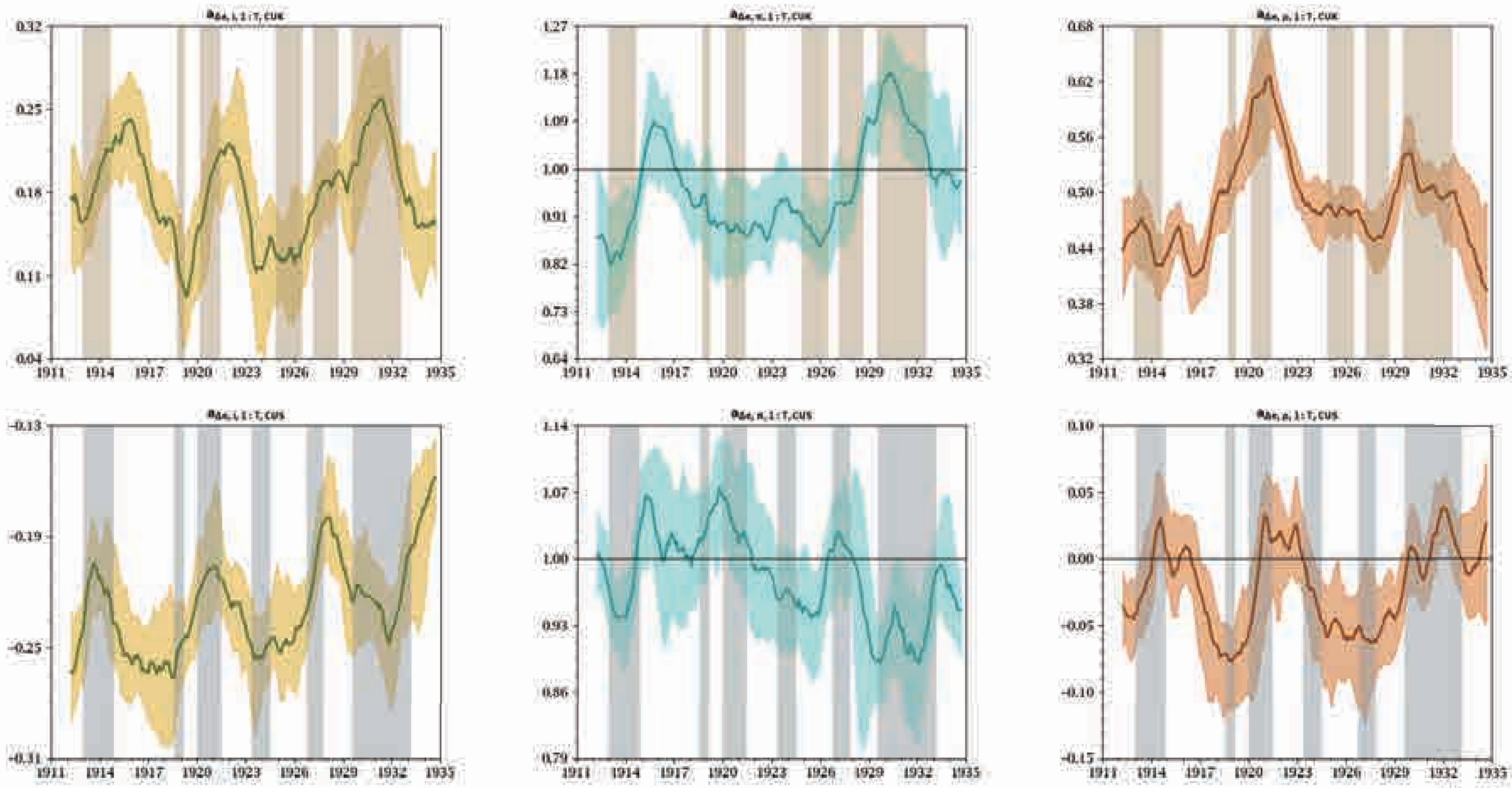
Notes: Log marginal data densities (MDDs) appear in the column headed lnMDD. The MDDs are calculated using the harmonic mean estimator of Geweke (2005), 2,500 draws from the posterior of the TVP-SV-SVAR(2)s, the China-U.K. and China-U.S samples, $y_{CUK,1:T}$ and $y_{CUS,1:T}$, and our priors. The support $y_{CUK,1:T}$ or $y_{CUS,1:T}$ give to a TVP-SV-VAR(2) is summarized by its MDD. The column headed WAIC reports the widely applicable information criterion of Watanabe (2010), which is also known as the Watanabe-Akaike-IC. The WAIC is an estimate of the 1-month ahead predictive loss of a TVP-SV-SVAR(2). Gelman et al (2014) advise computing the predictive loss as twice the difference between a penalty term and the mean of the log predictive likelihoods. Estimates of the likelihood are obtained from the predictive steps of the Kalman filter and posterior of a TVP-SV-SVAR(2). The penalty term of the WAIC measures the effective dimension of the parameter vector. It is the sum of the posterior variances of the likelihood of a TVP-SV-SVAR(2). Values in bold are the largest ln MDD and smallest WAIC on $y_{CUK,1:T}$ or $y_{CUS,1:T}$.

Figure 1: China-U.K. and China-U.S. Samples, 1912M04 to 1934M09



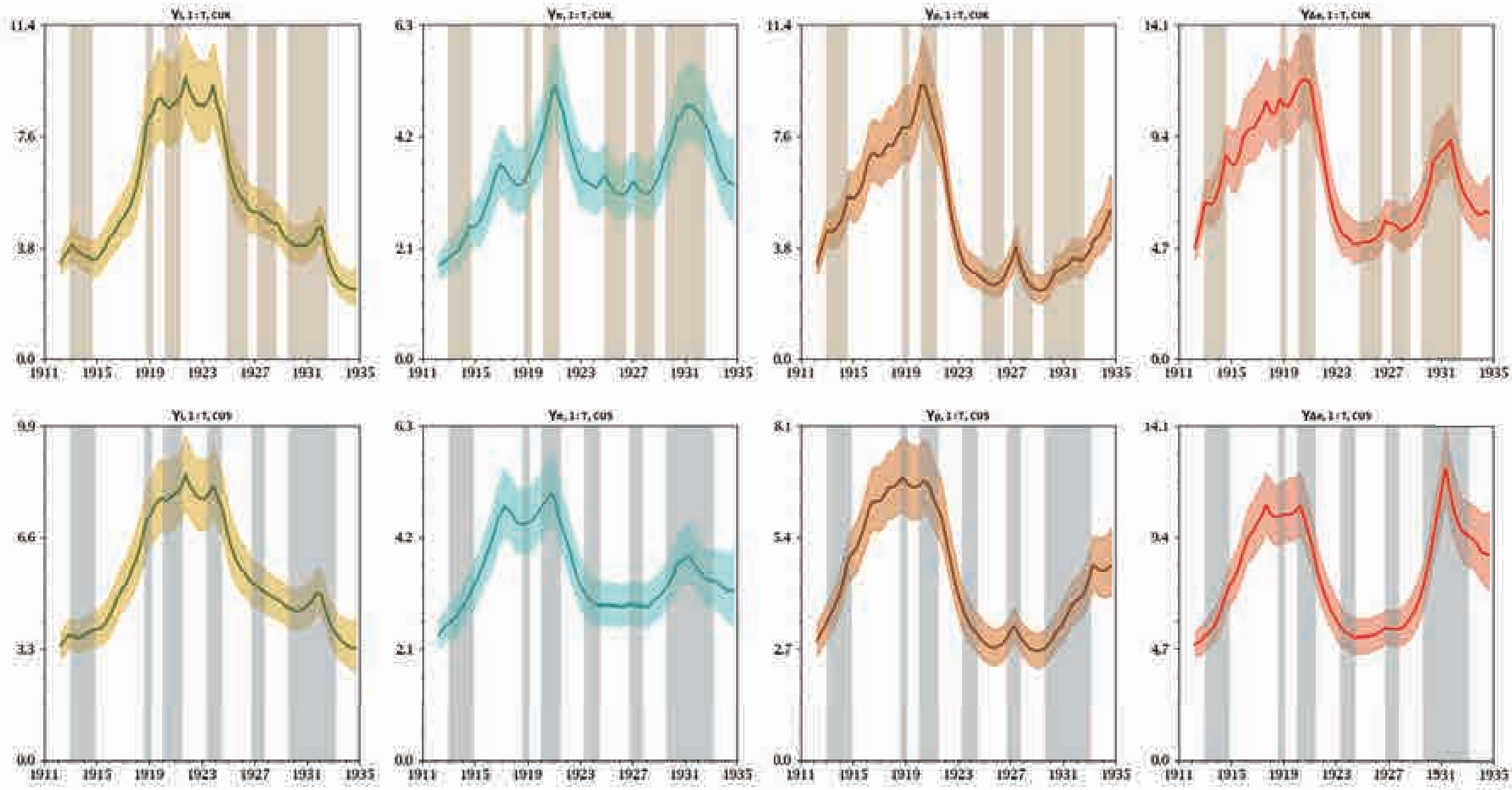
Notes: The top left panel plots the interest spreads, $i_t = i_{S,t} - i_{\ell,t}$, $\ell = UK, US$. Plots of the inflation differentials, $\pi_t = \pi_{S,t} - \pi_{\ell,t}$, appear in the top right panel. Risk premiums as deviations from parity, $\rho_{\ell,t}$, are displayed in the bottom left panel. The bottom right panel contains month over month nominal currency returns, $\Delta e_{GBP/S,t}$ and $\Delta e_{USD/S,t}$. The tan (silver) shaded vertical bars are Burns-Mitchell (NBER) recession dates for the U.K. (U.S.).

Figure 2: Posterior Moments of the $a_{1:T}$s on the China-U.K. and China-U.S. Samples, 1912M04 to 1934M09



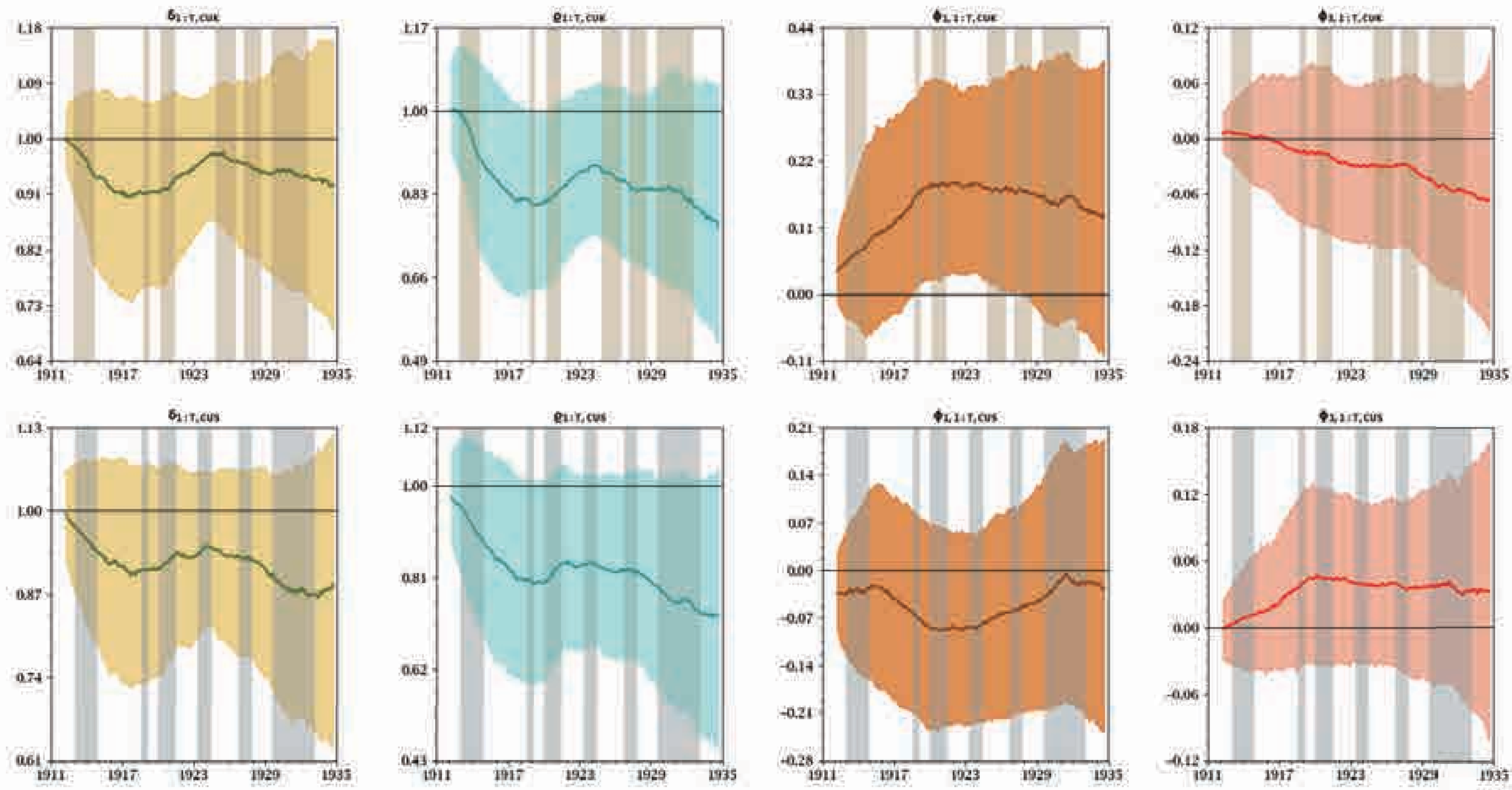
Notes: The top (bottom) row of panels contain solid lines that are the medians of the posterior distributions of $a_{\Delta e,i,1:T,CUK}$, $a_{\Delta e,\pi,1:T,CUK}$, and $a_{\Delta e,\rho,1:T,CUK}$ $\left(a_{\Delta e,i,1:T,CUK},\ a_{\Delta e,\pi,1:T,CUK},\ \text{and}\ a_{\Delta e,\rho,1:T,CUK}\right)$ conditional on TVP-SV-SVAR-BL, $y_{CUK,1:T}$ $\left(y_{CUS,1:T}\right)$ and our priors. The shadings around the medians of the posterior are 68% Bayesian credible sets (*i.e.*, 16th and 84th quantiles). The tan (silver) shaded vertical bars in the upper (lower) row are Burns-Mitchell (NBER) recession dates for the U.K. (U.S.).

Figure 3: Posterior Moments of $\gamma_{1:T}$ on the China-U.K. and China-U.S. Samples, 1912M04 to 1934M09



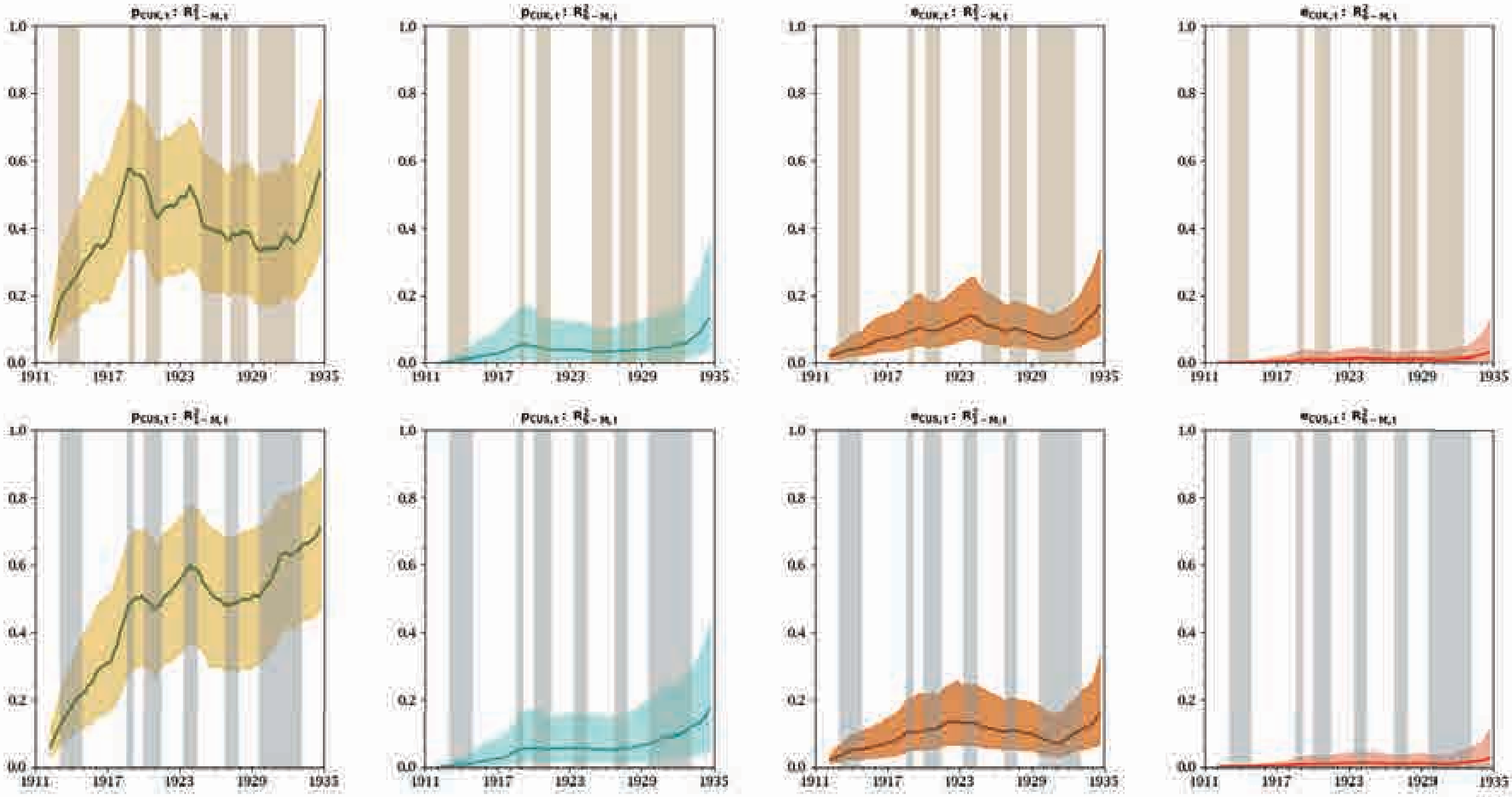
Notes: The top (bottom) row displays panels with solid lines plotting medians of posterior distributions of $\gamma_{j,1:T,CUK}$ $\left(\gamma_{j,1:T,CUS}\right)$ conditional on TVP-SV-SVAR-BL, $\mathcal{Y}_{CUK,1:T}$ $\left(\mathcal{Y}_{CUS,1:T}\right)$ and our priors, $j = i$, π, ρ, and Δe. The shadings around the posterior medians are 68% Bayesian credible sets. The tan (silver) shaded vertical bars in the upper (lower) row are Burns-Mitchell (NBER) recession dates for the U.K. (U.S.).

Figure 4: Slope Coefficients of the Fama and Engel UIP Regressions on the China-U.K. and China-U.S. Samples, 1912м04 to 1934м09



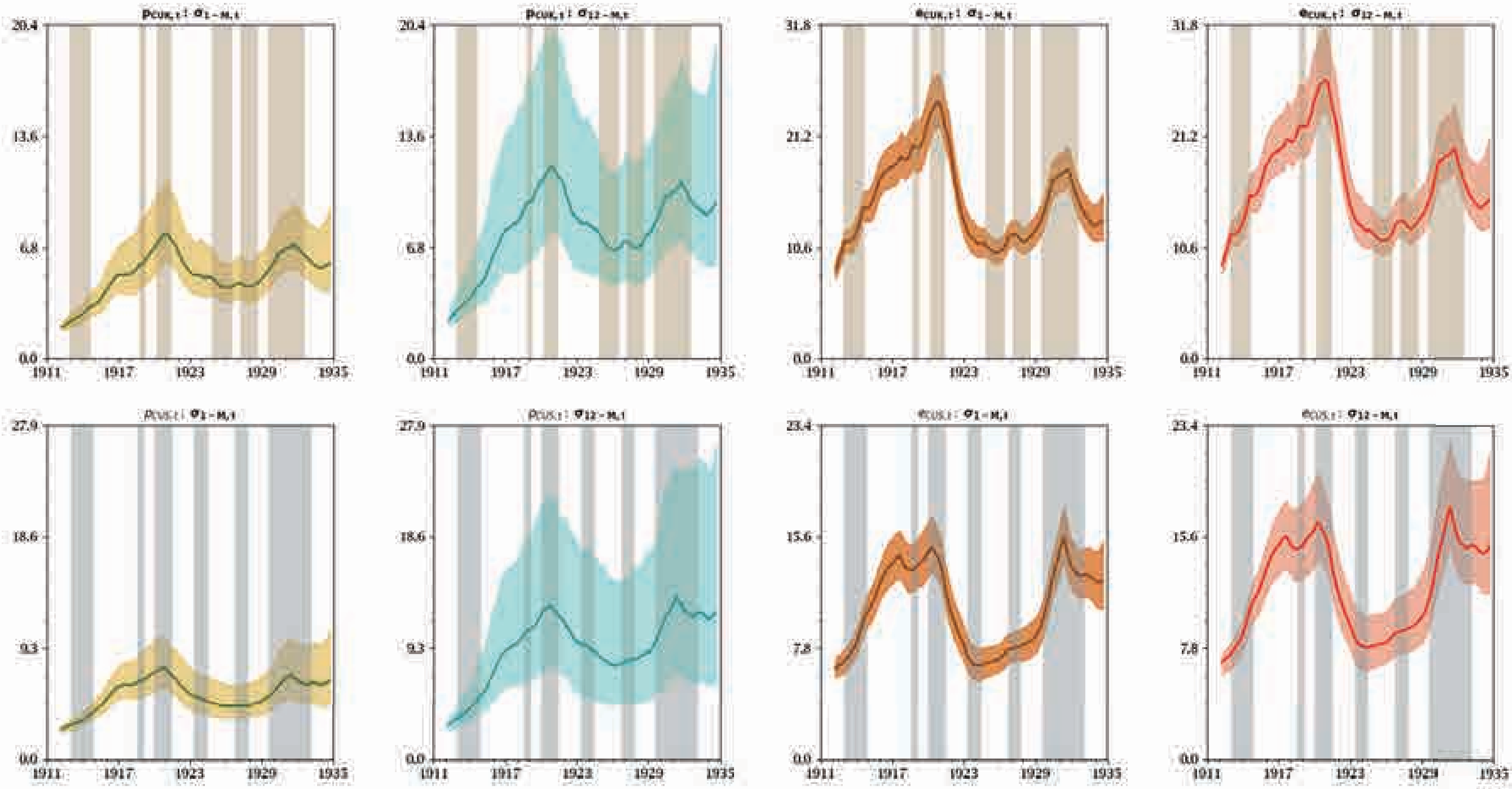
Notes: Moving from left to right, the figure presents posterior distributions of the slope coefficients $\delta_{1,t}$, $\varrho_{1,t}$, and $\phi_{H,t}$ from regressions of Δe_{t+1} on an intercept and $i_{j,t} - i_{S,t}$, Δq_{t+1} on an intercept and $r_{j,t} - r_{S,t}$, and ρ_{t+1} on an intercept and $\sum_{h=0}^{H-1}\left(r_{j,t-h} - r_{S,t-h}\right)$, where $j = UK, US$ and $H = 1, 3$. The top (bottom) row has slope coefficients computed using the posterior distributions of TVP-SV-SVAR-BL conditional on $\mathcal{Y}_{CUK,1:T}$ $\left(\mathcal{Y}_{CUS,1:T}\right)$, and our priors. The solid lines are posterior medians of $\delta_{1,t}$, $\varrho_{1,t}$, $\phi_{1,t}$, and $\phi_{3,t}$. The shadings around the posterior medians are 68% Bayesian credible sets. The tan (silver) shaded vertical bars in the upper (lower) row are Burns-Mitchell (NBER) recession dates for the U.K. (U.S.).

Figure 5: Posterior Moments of $\mathcal{R}^2_{p,h,1:T}$ and $\mathcal{R}^2_{e,h,1:T}$ on the China-U.K. and China-U.S. Samples, 1912M04 to 1934M09



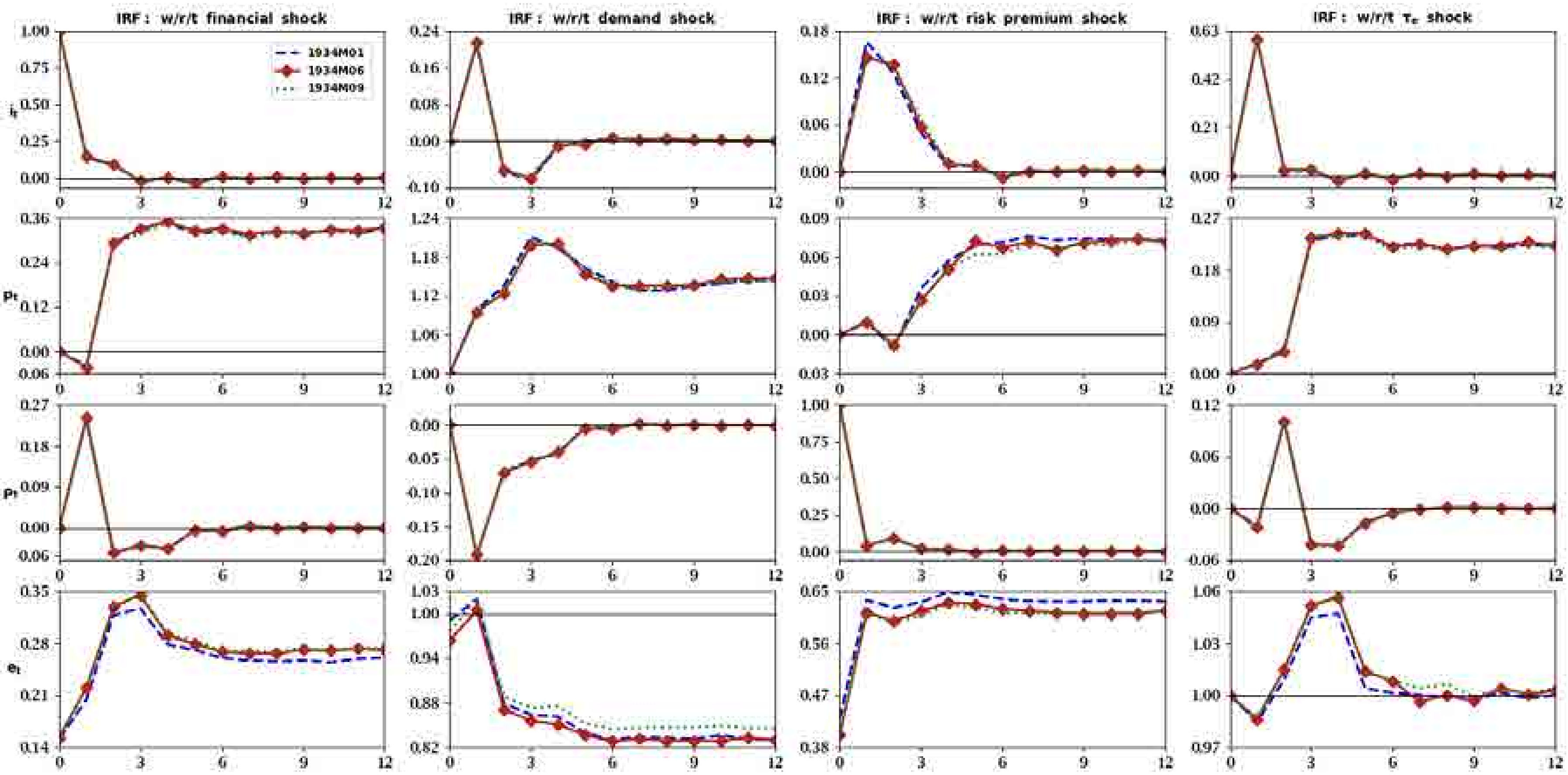
Notes: The top (bottom) row plot the medians of the posterior distributions of $\mathcal{R}^2_{z,1,1:T}$ and the $\mathcal{R}^2_{z,6,1:T}$ produced by TVP-SV-SVAR-BL, $\mathcal{Y}_{CUK,1:T}$ $\left(\mathcal{Y}_{CUS,1:T}\right)$, and our priors, where $z = p$ or e. The shadings around the posterior medians are 68% Bayesian credible sets. The tan (silver) shaded vertical bars in the upper (lower) row are Burns-Mitchell (NBER) recession dates for the U.K. (U.S.).

Figure 6: Posterior Moments of $\sigma_{p,h,1:T}$ and $\sigma_{e,h,1:T}$ on the China-U.K. and China-U.S. Samples, 1912M04 to 1934M09



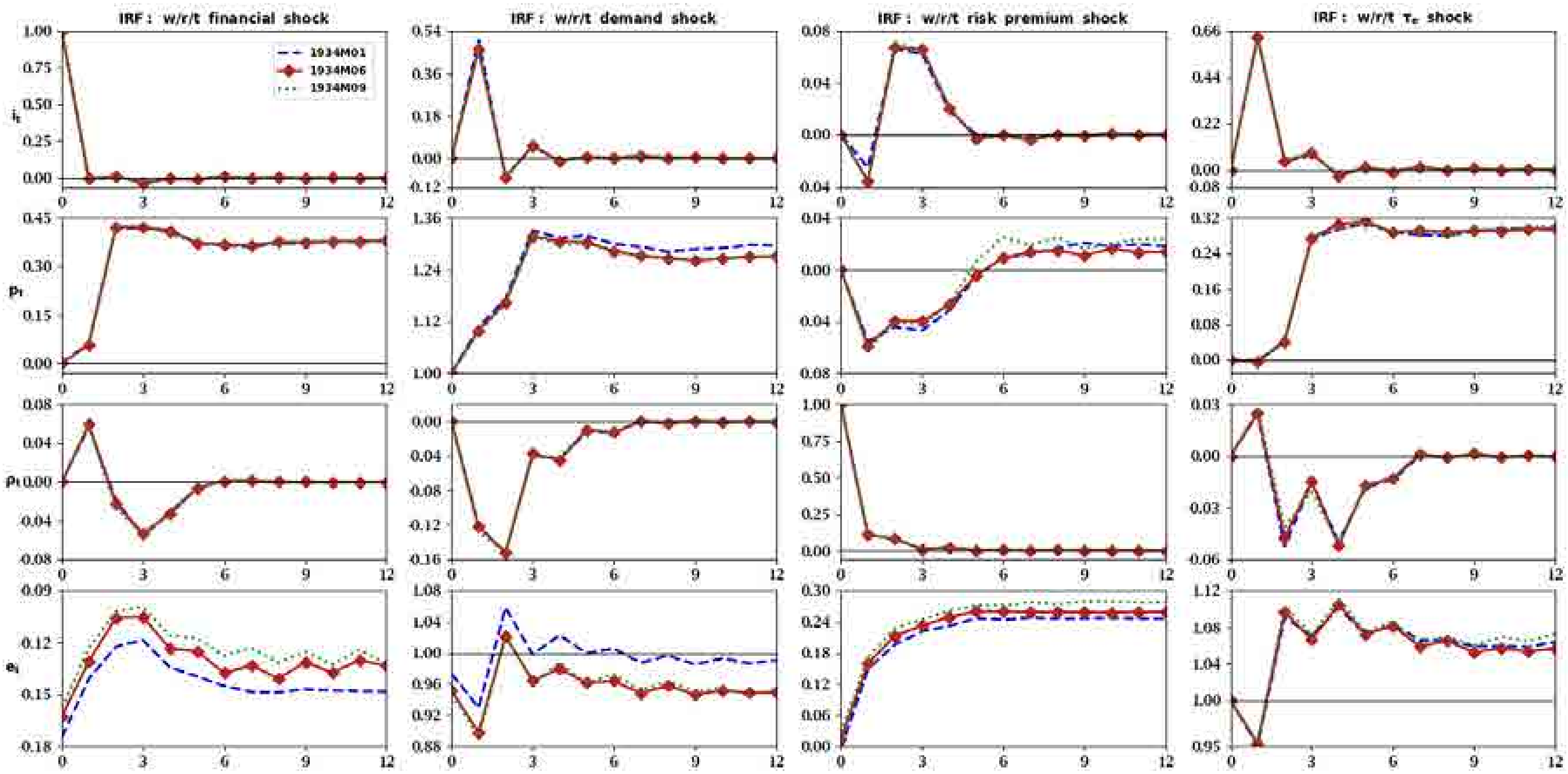
Notes: The top (bottom) row plot the medians of the posterior distributions of $\sigma_{z,h,1:T} = \sqrt{\mathcal{V}_t\left(z_{t+h} - \mathbf{E}_t z_{t+h}\right) + \left(\mathbf{E}_t z_{t+h} - z_t\right)^2}$ produced by TVP-SV-SVAR-BL, $\mathcal{Y}_{CUK,1:T}$ $\left(\mathcal{Y}_{CUS,1:T}\right)$, and our priors, where $z = p$ or e, and $h = 1$ and 12. The shadings around the posterior medians are 68% Bayesian credible sets. The tan (silver) shaded vertical bars in the upper (lower) row are Burns-Mitchell (NBER) recession dates for the U.K. (U.S.).

Figure 7: Impulse Response Functions in 1934 on the China-U.K. Sample



Notes: The top, second, third, and bottom rows display IRFs of i_t, p_t, ρ_t, and e_t at 1934M01 (dashed line), 1934M06 (diamond line), and 1934M09 (dotted line) computed using the posterior distributions of TVP-SV-SVAR-BL on $\mathcal{Y}_{CUK,1:T}$ and our priors. From right to left, the columns plot IRFs with respect to the international financial, nominal cross-country demand, risk premium, and trend exchange rate shocks. The IRFs run from impact to a 12-month horizon, h = 0, 1, . . . , 12.

Figure 8: Impulse Response Functions in 1934 on the China-U.S. Sample



Notes: The top, second, third, and bottom rows display IRFs of i_t, p_t, ρ_t, and e_t at 1934M01 (dashed line), 1934M06 (diamond line), and 1934M09 (dotted line) computed using the posterior distributions of TVP-SV-SVAR-BL on $\mathcal{Y}_{CUS,1:T}$ and our priors. From right to left, the columns plot IRFs with respect to the international financial, nominal cross-country demand, risk premium, and trend exchange rate shocks. The IRFs run from impact to a 12-month horizon, $h = 0, 1, \ldots, 12$.