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R* in East Asia: Business, Financial Cycles, and Spillovers[1]

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ABSTRACT

This paper provides new estimates of the neutral interest rate, or r*, with a frequency domain approach using quarterly data from China, Japan, Korea, and the US. Utilizing band spectrum regressions, we estimate two types of neutral rates, which hold over the business cycle and the financial cycle respectively. To account for uncertainty around estimates of r*, we derive confidence bands via a thick modelling approach. Our estimates share a few common features with existing published estimates. Consistent with prior research, a downward trend in r* is observed, although the trend becomes less obvious when uncertainty bands are factored in. Meanwhile, our findings offer novel perspectives on the neutral rate in the four countries examined. For individual countries, our estimates for the two types of r* do not always track each other, suggesting that central banks face trade-off between business versus financial cycle considerations when setting the policy rate. Across countries, we identify significant positive spillovers from the US to the three East Asia countries, as well as spillovers from China to Kora and Japan.

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1. Introduction

Evaluating the stance of monetary policy continues to be a heavily debated question, especially given the extraordinary global rise and fall in inflation since 2021. In recent years, the discussion has centered around the concept of the neutral real interest rate, or r*. A commonly used definition of r* is the interest rate that prevails when economic slack is zero and inflation is stable. Observers then ask, by comparing r* with existing policy rates in real terms, whether current monetary conditions are overly accommodative or excessively restrictive.

The neutral rate of interest, or r*, is inherently unobservable as it is a hypothetical concept. Consequently, it must be inferred through estimation. Many estimates of r* are derived from semi-structural models, such as those pioneered by Laubach and Williams (2003) and later extended by Holston et al. (2016, 2023). The Laubach-Williams model is a linearized New Keynesian framework that incorporates an IS curve, which links the output gap to real interest rates, and a Phillips curve, which connects the output gap to inflation. The estimation of r* relies on the assumption that rising/falling inflation typically indicates that current interest rates are below/above the natural rate. Alternative estimates of r* are obtained from structural DSGE models, where r* represents the interest rate that would prevail in a hypothetical economy with flexible prices (e.g., Del Negro et al., 2017, 2019). Some other approaches use reduced-form time series models, where r* is estimated as a forecast of the future real short-term interest rate (e.g., Lubik and Matthes, 2015 and Morley et al., 2024) or as a long-run average of real short-term interest rate (e.g. Hamilton et al., 2016 and Fiorentini et al., 2018).[2,3]

Existing estimates of r* predominantly focus on the US. This emphasis is driven by several factors: the availability of a long span of data, the well-established role of the inflation and output gaps as adequate explanations of the stance of US monetary policy as in a Taylor-type rule, and the pivotal role of U.S. monetary policy in shaping the global economy and financial

[2] Additional methodologies involve surveys of financial market participants and the use of term structure models. For comprehensive reviews on the definition, estimation, and implications of r*, see Gust (2015), Obstfeld (2023), and Benigno et al. (2024).

[3] Estimates of r* can be sensitive to the approach and the sample period that are used to derive it. Also, the estimates are often associated with considerable uncertainty (e.g., see Del Negro et al. (2017, 2019), Fiorentini et al. (2018) and Brand and Mazelis (2019)). Therefore, some papers are skeptical of the usefulness of r* as a lodestar for monetary policy (inter alia, see Borio (2024, 2021), Borio et al. (2022), and Stein (2024)).

markets. For r* estimates in other advanced economies or the global economy, methodologies are often heavily influenced by the U.S. experience. Examples include Wynne and Zhang (2018), Grossman et al. (2021), and Adjalala et al. (2024).

Fewer estimates of r* for other regions of the world, including East Asia, are available. Chen and Siklos (2024) provide an overview of efforts to estimate r* for China, which largely follow the framework established by Laubach and Williams (2003). Similarly, studies on r* for Japan and Korea tend to adopt approaches developed for the U.S. Examples include Ikeida and Saito (2014), Sudo et al. (2018), and Han (2019) for Japan, as well as Lee et. al (2020), Rafiq (2021), and Kang and Do (2024) for Korea.

Nevertheless, the U.S. approach may not be directly applicable to East Asian countries. As open economies, the interaction between monetary policy and economic developments in these countries is not adequately captured by the semi-structural models designed for the US. Specifically, factors such as exchange rates and capital flows—both highly sensitive to external developments—often play a significant role in shaping monetary policy decisions and macroeconomic dynamics, which are ordinarily treated as less critical for the US. In addition, other economies must also contend with spillover effects from the US whereas spillovers from other countries to the US are likely to be relatively less important.

Moreover, the existing methodology primarily focuses on the interest rate that stabilizes the economy at the business cycle frequency, based on the definition of r* as the rate that aligns output and inflation with their potential levels. However, recent recessions triggered by financial crises have prompted policymakers to consider not only the stabilization of the business cycle but also that of the financial cycle (see, for example, Danthine (2012)). Over the past few years, a growing body of literature has highlighted the existence of a financial cycle, which is believed to be significantly longer than the standard business cycle (e.g., Drehmann et al. (2015), Yan and Huang (2020)). While business cycles typically last two to five years, financial cycles often span 10 to 20 years (e.g., Borio (2014)). It is natural, therefore, to think that there are two types of r*, namely one that evolves with the business cycle and another that evolves with the financial cycle. If the characteristics of financial cycles differ significantly from those of regular business cycles, the natural rate of interest that stabilizes the business cycle may diverge from the natural rate that stabilizes the financial cycle. While the responsibility for stabilizing financial cycles often falls to macroprudential policies rather than monetary policies, it remains important to use r* at the financial cycle frequency as a benchmark to understand the role monetary policy plays in shaping the development of financial cycles. There has been research examining r* from a similar perspective. For instance, De Fiore and Tristani (2011) analyze r* within a structural model for the US incorporating financial frictions and demonstrate that r* responds differently to shocks in economies with and without such frictions. Similarly, Akinci et al. (2020) estimate r* for financial stability, which they term r**, using a structural model for the US. Their framework models banks with leverage constraints that tighten as portfolio risk increases, with r** as the threshold where the constraint binds. It is intended to serve as a benchmark for financial stability, much like r* serves as a benchmark for monetary policy aimed at macroeconomic stability. However, neither study explicitly considers the financial cycle, which has been shown to exhibit very different properties compared to the business cycle. Our study aims to provide new estimates of r*, focusing on developments in China, Korea, and Japan. Rather than adopting the common approach of estimating a semi-structural model, we take a time series perspective rooted in the frequency domain. The key advantage of our methodology lies in its ability to differentiate r* over business and financial cycles. This allows us to capture the role of monetary policy in stabilizing the economy separately over business and financial cycles respectively. In doing so, we incorporate the consensus that factors impacting r* over time differ over the horizon monetary policy makers consider while financial factors take a considerably longer time to work their way through the economy. Having said this, r* at business cycle frequencies are also interconnected with r* at financial cycle frequencies and vice versa, as financial conditions and real activity are closely linked. Our empirical results clearly show this. However, business cycles and financial cycles operate at different frequency. Stabilizing the economy at the business cycle frequency is distinct from stabilizing it at the financial cycle frequence. In time of financial crises or deep recessions, the two can diverge quite a bit, therefore policy makers might face a trade-off on whether the priority is to maintain financial stability or to pull the economy out of deep recessions.

Specifically, we apply Engle's (1974) band spectrum regression technique, which facilitates the conversion of frequency domain estimation into the more interpretable time domain. We run band spectrum regression of real policy rate on four factors, separately over business cycle and financial cycle frequencies. These factors include an economic factor, a financial factor, a monetary factor, and a structural factor. The structural factor is an attempt to capture determinants of r* including productivity and demographic factors. The factors are constructed with a factor model that synthesizes information from a large set of variables.[4] Our sample begins in the late 1990s or early 2000s, depending on data availability for each country. To assess the uncertainty around our point estimates, we exploit the ability to generate several estimates of r* estimates and utilize the thick modelling approach (Granger and Jeon (2004)) to quantify coefficient and model uncertainty surrounding of r*.

In addition, we examine possible spillovers of r* across different countries.[5] Given the well-established role of the US in influencing monetary policy in other countries, we first ask whether changes in the neutral rate in the US also influence r* in the three East Asia economies considered. We also examine whether China's monetary policy generates global spillover effects considering China's rising role in the global economy. For example, Miranda-Agrippino et al. (2020) found that monetary policy in both the US and China has a significant global impact.

To briefly summarize our findings, our estimates for the US and Japan are comparable, but distinct from existing published estimates. The downward trend in r* reported in previous literature is replicated in our study. Nevertheless, the decline in r* in China is a relatively recent phenomenon that has shown signs of a reversal after the pandemic era. A similar notable reversal in r* is also observed in the remaining three economies considered. That said, when uncertainty of point estimates is factored in, the secular decline in r* is less evident. The assessment of whether monetary policy is too loose or too tight is sensitive to the cycle considered. For example, in China and Korea the stance of monetary policy is closer to neutral when the financial cycle is considered than when equilibrium is sought over the business cycle. In Japan and the US deviations from neutrality are also most visible at the financial cycle when there are large shocks (e.g., dot-com bubble, global financial crisis, Abenomics introduction and the pandemic). Significant positive spillovers from the US and China on Korea's and Japan's r*, and

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[4] Due to the lack of consensus in the literature about the most important determinants of r* (IMF (2023c)), the data-driven factor model approach is preferred to hand-picking a few variables.

[5] There is a large body of literature examining sources and consequences of spillovers from major advanced economies, especially the US. Space limitations prevent a comprehensive review, but a selection of recent contributions includes Cotter et al. (2023), Kearns et al. (2023), Ferreira and Shousha (2023), and Ilori et al. (2022). China's role in generating spillover effects has gained attraction recently (e.g., see Chen and Siklos (2023), Sznajderska and Kapuscinski (2020)).

similar spillovers from US r* to China are also found. Moreover, r* and spillovers are sensitive to the cycles considered in all four countries. Our results suggest that central banks trade-off business versus financial cycle factors when setting the policy rate.

What do we know about r* individually in the countries that are the focus of the present study? A selective review suggests some contradictory findings. The focus of most studies is on the role of trend growth (or potential output) and inflation performance in dictating the value of r*, and little else. An explicit role for the financial cycle is generally downplayed as is the economy's openness. The broader r* that allocates a relatively larger role to longer run factors places great emphasis on demographic drivers (e.g., Goodhart and Pradhan (2017, 2020), Carvalho, Ferrero and Nechio (2016)). Demographic factors are believed by some to be an important driver of r* in East Asia though, perhaps in part because of the theoretical literature and model constructs, productivity is often thought to be the dominant driver. This is found to be the case not only in advanced economies such as the US and Japan (e.g., see Ikeda and Saito (2014)) but also for China, and Korea (e.g., see Li and Sun (2020), Sun and Rees (2021)), Kang and Do (2024)). Indeed, Korea's experience is thought to be the mirror image of Japan's experience albeit with a long lag (see Lee, Hahm, Park, and Young Park (2020)).

That said, the demographic factor must compete with several others put forward to explain the decline in the neutral rate over the past few decades. A non-exhaustive list of determinants includes openness to foreign shocks, the degree of capital markets integration, exchange rate and monetary policy strategies, and even geopolitical risks. Relevant studies that explore how these determinants impact r* include Cesa-Bianchi, Harrison, and Sajedi (2023), Grossman, Martinez-Garcia, and Wynne (2018), and Kiley (2020a, 2020b).[6]

The remainder of the paper is organized as follows. Section 2 outlines the methodology while section 3 describes the data. Section 4 presents our main findings while section 5 concludes.

[6] The list omits how unconventional monetary policies (UMP) influenced r* (e.g., see Filardo and Nakajima (2018)) as well as the role of climate change and digitalization (Schnabel (2024)).

2. Methodology

2.1 Estimating r* based on band spectrum approach

As is true elsewhere, r* in East Asia is likely to be determined by a variety of factors. Most applications focus on short-run (i.e., monetary conditions) determinants of the neutral rate although medium to longer-term factors (e.g., financial assets, productivity, demographics) are also known to play a role. Generally, broader real economic factors (e.g., aggregate demand and supply) and monetary policy are most likely to influence r* at business cycle frequencies. Not to be forgotten are global economic factors, that is, spillovers from abroad. Global determinants suggest the possibility that changes in r* in major economies (USA and China, in the present context) may be transmitted to the other two East Asian economies considered here, namely Korea and Japan. Indeed, we must additionally acknowledge the likelihood that r* in the US impacts r* in China.

Because of the connotation that the determinants of r* are sensitive to the length of a cycle it seems natural to think in terms of the frequency domain approach to estimating r*. A limitation of frequency domain estimation may be that conditional relationships are more difficult to interpret in familiar economic terms. Fortunately, the band spectrum approach (BSR), due to Engle (1974), is well-suited to separately identify cyclical influences on the neutral rate in a time domain setting.

A useful way of thinking about the BSR technique is as a method for detecting sensitivity in parameter estimates by recognizing that time series evolve with different periodicities. Thus, for example, it is common to argue that monetary policy acts with long and variable lags, while financial shocks can act quickly but have long lasting economic effects. Similarly, productivity and demographic factors are believed to be slow moving factors. Engle (1978) uses BSRs to investigate price equations across various cycle length. Sims (1972) and Engle (1974) also apply the BSR technique to investigate economic issues.[7] In Siklos (1988), higher frequencies (i.e., cycles of short duration) are assumed to be akin to periods when policies are unanticipated while business cycle frequencies are defined to be consistent with periods when policies become fully

[7] Other more econometrically focused studies (e.g., small sample properties) of the BSR technique include Engle and Gardner (1976). Engle (1980a) introduces maximum likelihood estimators while still other econometric issues are addressed in Engle (1980b).

anticipated. This distinction serves as the basis for estimating Phillips curves using US data. A discrete Fourier transform[8] of the series is employed to filter out the undesired frequencies of interest before inverting the complex series back into the time domain where a time domain regression is estimated. Hannan (1963a, 1963b, 1965) originally developed efficient estimates (i.e., a form of Generalized Least Squares) reported below although the inspiration for the technique goes back much earlier.[9]

At a more intuitive level, BSR analysis amounts to estimating a regression where the effects are restricted to cycles defined by the researcher. As noted previously, estimates at different cycles are estimated via the frequency domain and the BSR technique inverts the results into the more familiar time domain regression form.

Consider the simple regression written as follows

$$y_t = \beta x_t + \varepsilon_t \qquad (1)$$

Conversion into the frequency domain is done via the so-called Fast Fourier transformation (FFT) that converts a time domain signal into components at different frequencies (i.e., short to long-run).[10] In the present context this effectively means we can re-write (1) in the frequency domain as

$$y_\omega = \beta x_\omega + \varepsilon_\omega \qquad (2)$$

where $\omega = [0,0.5]$ designates the frequency which ranges, in the case of the quarterly data used in the present study, from 2 quarters (i.e., $\omega = 0.5$, or 1/0.5=2) to infinity. However, as the name implies, BSR uses a range of ω set by the researcher relying on the definitions above to characterize the length of the business and financial cycles. . A slightly different way of thinking about the problem is that (2) is the multivariate equivalent of the so-called band pass filter

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[8] Essentially a function that maps a series from the time domain into the frequency domain where variability is measures across frequencies. Thus, the 'long-run' is the zero frequency while the highest frequency, at least for quarterly data is 2 quarters. This is what gave rise to Granger's (1966) seminal piece that describes a typical time series as having greatest power at low frequencies in part because many time series are dominated by trend like movement. Medel (2014) confirms Granger's original findings.

[9] Whittle (1951) introduced the idea of time series decomposition in the frequency domain to estimate of economic time series relationships. See also Amemiya and Fuller (1967).

[10] A classic account of frequency domain analysis is Granger and Hatanaka (1964). Also see Warner (1998). For the details about the FFT (and IFT; see below) see, for example, Nussbaumer (1982)

(Christiano and Fitzgerald (2003)) which unconditionally filters out variation in time series at chosen frequencies.[11] Once (2) is obtained the BSR procedure inverts the results, using an inverse Fourier transformation (IFT), back into the time domain such that regression (2) becomes

$$y_t^{BC} = \beta^{BC} F_t + \varepsilon_t^{BC} \quad (3)$$

$$y_t^{FC} = \beta^{FC} F_t + \varepsilon_t^{FC} \quad (4)$$

Equations (3) and (4) now make clear that the relationship between F_t and y_t is conditional is restricted to the business and financial cycles, as defined above. However, instead of representing the results in terms of spectra or cross-spectra, the BSR expresses the results in the more familiar time domain but constrained to permit variation at the previously defined periodicities.

Since r* is determined by a variety of factors above we first proceed by defining them more precisely. The concept of factors also captures the notion that many time series simultaneously contribute to creating business and financial cycles in time series.[12] We define four factors, namely an economic factor, a financial factor, a monetary factor, and a structural factor. These factors are defined in the introduction and are assumed to explain movements in r*. Adding the impact of the USA on the remaining three countries, and the impact of China on Korea and Japan, which can be thought of as global factors, is also taken into account. We use F_{ijt} to denote factor *i* (*i* = ECONOMIC, FINANCIAL, MONETARY, STRUCTURAL) for country *j* (*j* = China, Korea, Japan, USA) at time *t*. If **Y** denotes the vector of variables used to define each one of the factors considered, we write

$$\mathbf{Y}_t = \boldsymbol{\beta}\mathbf{F}_t + \boldsymbol{\zeta}_t \quad (5)$$

where **Y** is a vector of observed time series from which factors **F** are estimated, $\boldsymbol{\beta}$ are the factor loadings. Estimates of the factors, namely $\widehat{\boldsymbol{F}}_t$, otherwise known as the factor scores, serve as proxies for the four factors listed above. The ECONOMIC factor can, in principle, be further

[11] It is partly the unconditional nature of the filter is what gave rise to criticisms that band pass type filters can generate spurious cycles. See, for example, Benati (2001) and Schuler (2018).
[12] For example, this helps explains the creation of the FRED-MD and FRED-QD data sets (McCracken and Ng (2016)) for the US.

decomposed into aggregate demand or supply factors, depending on whether the data (i.e., the size and sign of the factor loadings) suggest they are separately identifiable.

Common components are extracted via principal components analysis (PCA). They are estimated by relying on Bai and Ng's (2002) IC_{p2} criterion, a choice that also receives support from the analysis of Stock and Watson (2016, p. 436). In addition, Hamilton and Xi (2024) demonstrate that the common component obtained from PC can be estimated without worrying about whether the series are stationary in levels or after differencing. Nevertheless, as part of a robustness exercise, we also investigate how our estimates of r* are affected according to whether the series under investigation are stationary or not in levels or log levels.[13]

In estimating (5) we collect series that are representative of each one of the factors listed. Details of the series that make up each of the factors considered are relegated to the appendix. We adopt this strategy because it is more likely to produce factor loadings consistent with economic theory. Moreover, we follow others who have used it previously and found the results are more likely compatible with economic priors (e.g., see UNCTAD (2023)), Chen and Siklos (2022), and references therein).[14]

Our methodological approach is also consistent with existing economic analysis that explore the determinants of r* over time. It is worth noting that existing studies generally focus on the experience of economies such as the US which are modelled using the New Keynesian approach. It is far from clear that this structure is appropriate, for example, in China's case. The PBoC's monetary policy, restrictions on the movement of capital, and China's role as a large open economy argues for a more data-driven approach in estimating r*. China is also large and relatively open economy, but r* estimates are typically driven by models calibrated to a much more closed economy such as the USA (e.g., see Chen and Siklos (2022)).[15] Obstfeld (2023) argues that r* cannot be understood without reference to an economy's openness. Korea is the archetypical example of a small open economy. Japan, of course, is also a large open economy

[13] We also perform panel unit root test, with and without cross-sectional dependence to substantiate the unit root properties of the data (not shown).

[14] Factor models often rely on many time series. The total number of series used in our study is comparable to ones used in other studies such as, for example, Stock and Watson (2016), or Hatzius, Hooper, Mishkin, Schoenholtz, and Watson (2010).

[15] Chortareas and Kaykhusraw (2023) and Platzer and Peruffo (2022) are exceptions.

and, while there is a resemblance to the US experience, the decades of low inflation and even mild deflation set it apart from other advanced economies.

The US Federal Reserve is also the central bank most likely to dictate monetary and financial conditions around the globe. Shocks in the US monetary policy are usually transmitted to other economies and induce the global monetary and financial cycle (Miranda-Agrippino and Rey (2020)). Therefore, we also consider econometrically potential links between r* from the USA to China. The extant literature suggests that r* spillover effects are potentially large, and persistent, and they can be amplified by crisis conditions (financial, real or monetary; e.g., see Cotter, Hallan, and Yilmaz (2023) and Kohlscheen, Mojon, and Rees (2020)). Indeed, the financial factor defined above is often omitted in standard estimates of r*. Yet, this factor has been shown empirically to play an important role in the global economy (e.g., see Obstfeld (2023)). We account for the spillover effects from the USA and China by adding r* estimates for these countries to the model used to estimate the neutral rate in Korea and Japan. We also estimate a model for China that admits spillover effects from r* in the US.

We can now write the specifications used to obtain r*. We proceed in two steps. First, we estimate the nominal equivalent of r*, namely R*.[16] In a second step we estimate r* as the difference between R* and a proxy for expected inflation. For the business cycle version, we use a short-term indicator of expected inflation (i.e., 1 to 3 years horizon). For the financial cycle we subtract estimates of long-run inflation expectations (i.e., 5 years and 6 to 10 years).

In the time domain our specification would be written

$$R_t^j = \alpha_{0j} + \alpha_{1j} ECONOMIC_t^j + \alpha_{2j} FINANCIAL_t^j + \alpha_{3j} MONETARY_t^j + \alpha_{4j} STRUCTURAL_t^j + \boldsymbol{\alpha_{5j} R_{t-1}^{*,l}} + \varepsilon_t \qquad (6)$$

$$\alpha_i > 0, i = 1,2,3,5, \qquad \alpha_4 \lesseqgtr 0$$

where R is the central bank policy rate, ECONOMIC and FINANCIAL, MONETARY, and STRUCTURAL summarize the factors described previously, while ε is the usual error term, and the time subscript, *t*, indexes time. R* is then the predicted value from the regression (i.e., $R_t^{*,j} =$

[16] Mann (2025) also prefers the nominal version of the neutral rate as a device to assess the stance of monetary policy.

$\hat{R}_t^j$). Finally, the superscript j indicates that separate estimates are to be generated for China (CHN), Korea (KOR), Japan (JPN), and the United States (USA). Equation (6), for simplicity, omits lags, or additional exogenous variables, but these can be added without affecting any of the arguments below. The vector $\boldsymbol{R}_{t-1}^{*,l}$ is meant to highlight that, in one variant of equation (6), lagged R* estimates for the USA and China are added as an exogenous variable when the neutral rate for KOR and JPN are estimated, while ***l***= USA when R* for China is estimated. The factors are estimated such that the signs for the estimated coefficients are expected to be as shown in (6). Hence, an improvement in economic conditions raises the neutral rate[17], as does a tightening of financial or monetary conditions, while a rise in STRUCTURAL implies an improvement in longer-run economic factors (e.g., productivity, or a rise in the savings rate) which also raises R*. That said, other elements of this factor can lead to a fall in R* according to the literature previously reviewed. Thus, for example, an ageing population and a rise in income inequality would put downward pressure on R*. Hence, the sign on the STRUCTURAL factor can be ambiguous. Finally, we expect a rise in the neutral rate in the US or China to also raise R* in the remaining economies in the data set. The interpretation of the remaining factors is relatively standard. A tightening of MONETARY or FINANCIAL conditions raises R as does an improvement in ECONOMIC conditions. Similarly, a loosening or deterioration in these same factors results in a lower R. Finally, we do not take a stand on the sign of the influence of USA and CHN R*. There are equally plausible arguments suggesting that the neutral rate in KOR, JPN, and CHN when USA R* is added, will rise in response to developments in these economies. This can be the case because of a significant global component driving inflation (e.g., Cascaldi-Garcia et. al. (2024), and Ciccarelli and Mojon (2010)). Similarly, foreign exchange adjustments and intervention or other domestically sourced interventions may be implemented to counter policies originating in the USA or CHN.

Now, as before, define ω as the cycle or periodicity of interest. Since the data are quarterly cycles range from as short as 2 quarters to infinity (i.e., the zero frequency). The two cycles of interest in determining the neutral rate are the business and financial cycles. We can re-write equation (6) as follows:

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[17] If the ECONOMIC factor is broken down into aggregate demand and supply components then the aggregate demand factor has a positive sign while the aggregate supply factor is expected to have a negative sign.

$$R_{\omega}^{j} = \alpha_0 + \alpha_1 ECONOMIC_{\omega}^{j} + \alpha_2 FINANCIAL_{\omega}^{j} + \alpha_3 MONETARY_{\omega}^{j} + \alpha_4 STRUCTURAL_{\omega}^{j} + \boldsymbol{\alpha_5 R_{\omega}^{*,l}} + \varepsilon_{\omega}^{j} \quad (7)$$

where ω and all the remaining variables were previously defined and replaces the time subscript. For simplicity we define $\omega = [BUS, FIN]$ where BUS and FIN are the financial cycles previously defined. In the case of the business, $\omega = [0.196, 0.785]$, that is, cycles ranging from 8 to 32 quarters; in the case of the financial cycle $\omega = [0.079, 0.157]$ or a range from 40 to 80 quarters. Notice that the cycles are distinct (i.e., they do not overlap). It is usually assumed that the processes are stationary, but this assumption can be relaxed (see Hannan (1970), chapter 8). Initial estimates discussed in the next section set $\boldsymbol{\alpha_5 = 0}$. We then consider estimates that relax this assumption. Finally, as explained above, BSR display estimates of (7) in their time domain equivalent. Hence, the coefficient signs are the same as ones in (6).[18]

A disadvantage of the BSR technique is that more parsimonious specifications are preferred since the span of data matters. This is, at least partially, offset by evidence that many time series contain a long memory component (e.g., see Robinson (2003)). In any case, the decomposition of R* into the four factors listed in equations (7) renders our specifications parsimonious. To repeat, while there is no strict definition for the business and financial cycles a vast literature generally defined business cycles as ranging from 2 to 8 years in duration (i.e., 8 to 32 quarters; see Cerra, Fatás, and Saxena (2020)). In the case of financial cycles considerable evidence suggest that they display durations of 10 to 20 years (i.e., 40 to 80 quarters; see Drehmann, Borio, and Tsatsaronis (2015)).[19]

Once $\hat{R}_t^k$ is estimated then r* is calculated as

$$\hat{r}_{jt}^{*} = \hat{R}_t^k - \pi_{jt}^{e,\omega} \qquad (8)$$

[18] As noted above, versions of (6) and (7) were also estimated using observable time series. Signs for the output and inflation gaps, the credit gap, TFP growth and fertility rate are expected to be positive. Since we use the BIS's real exchange rate data the sign on the real exchange gap is expected to be negative. Similarly, if we add a proxy for supply chain constraints the resulting coefficient is also expected to be negative.

[19] We also estimated equation (3) by generating band pass estimates of the four factors, using the Christiano and Fitzgerald (2003) filter and then estimating a regression of the form (2). Broadly speaking, the conclusions discussed in the next section hold but Hannan efficient estimation avoids the generated regressor problem as well as the possibility of including spurious cycles.

where all the variables have been previously defined except $\pi_{jt}^{e,\omega}$ which is expected inflation. We rely on professional forecasts (see the next section) to proxy these expectations and ω signals that short to medium-term inflation expectations are used for the business cycle frequency estimates of the neutral rate while long-run inflation expectations serve as the proxy for the financial cycle estimates of r*.

2.2 Constructing confidence band with thick modeling

A frequent criticism of existing r* estimates include their lack of precision. Thus, for example, the LW and HLW versions of r* for the USA do not include confidence intervals. In contrast, US estimates of r* by LM include standard errors, and these appear to be wide enough to encompass both the LW and HLW point estimates most of the time. Since there is a debate not only about the main drivers of r* but also about the relative importance of, say, business versus financial cycle effects on the neutral rate this implies that relying on a single model to obtain confidence interval estimates may not be ideal. Instead, we adopt Granger and Jeon's (2004) thick modelling approach wherein estimates from different models are retained to provide a range of estimates.

We consider three types of variations of the baseline model to construct the confidence band.

The first type is models with time-varying factor loadings. While our baseline model is estimated over the full available sample in which case β_{ij} is estimated and, therefore, does not vary over time it is possible that the factors loadings are time-varying. The time-varying common factors are estimated by the following steps. Next, using samples that span a 5-year period, each sample is rolled ahead two years at a time.[20] We extract only the first principal component for each sub-sample considered.[21] This generates overlapping samples. The estimated factor scores are then the time series consisting of either the first principal component or the arithmetic mean of the principal components when the samples overlap.

[20] The choice of a two-year rolling window is partly justified by the view that changes in the stance of monetary policy take around two years to have full effect. Longer windows reduce the amount of time variation permitted. We did some estimation with longer windows, and the results were very nearly the same as for the full sample estimates. Also see Chen and Siklos (2022).

[21] Examination of the explanatory power of the first principal component reveals that one PC is sufficient to capture the common variation of the series in $\boldsymbol{Y}$.

The second type is models with observed factors. Typically, in studies that report estimates of r*, the determinants are observed time series (e.g., inflation, credit, etc.). Accordingly, a separate set of r* were also generated by more conventional means. Thus, the output gap and expected inflation substitute for the ECONOMIC factor, the credit gap (observed less trend credit) replaces the FINANCIAL indicator, the real exchange rate gap is used to represent the MONETARY factor while, either productivity (i.e., TFP) growth or the age dependency ratio capture the STRUCTURAL determinant. We also consider the case where a constructed proxy for supply chain constraints (see next section) is added.

The third type is models estimated in first differences when series that have a unit root rather than levels that we discussed previously.

3. Data

Over 60 different series were collected for each of the four countries. Below we provide a listing of the sources used with fuller details about the series and data related issues relegated to Table A1 in the appendix. We only provide some of the main sources here. Both the Atlanta ("China's Macroeconomy: Time Series Data") and St. Louis Federal Reserve (FRED: Federal Reserve Economic Data) banks compile and update relevant data as well as productivity data for all four countries. The Bank for International Settlements (BIS) also maintains key macro-financial data (e.g., policy rates, housing prices, debt-service ratios, credit, consumer prices, nominal and real exchange rates) for the countries used in our study. Demographic data are available from the UN's database as well as the World Bank's Development Indicators. Data on global debt, as well as other macro-financial data, and commodity price data, are also available from the IMF. Expectations data (inflation and real GDP growth) were obtained from some of the sources listed above (e.g., FRED) for China from the People's Bank of China, as well as Consensus Economics. The latter data source is not publicly available. The same goes for the CEIC data source also used to collect a few additional series for China, Korea and Japan. Other data sources include the Bank of Japan, and the Bank of Korea.

With a few exceptions the raw data for the ECONOMIC, FINANCIAL and MONETARY factors are available either at the monthly or quarterly frequencies. Monthly data were converted to quarterly via arithmetic averaging. Data for the STRUCTURAL factor are generally only available at the annual frequency. The final complete sample used in the empirical work covering the relevant data were converted to the quarterly frequency via the Catmull-Rom spline.[22] Finally, in a few cases, some observations for the second half of 2023 for a few series were missing and were filled by generating 'prophet' forecasts. The technique relies on a Bayesian approach to forecast out of sample (see Taylor and Letham (2018)). The technique decomposes a time series into a trend and a residual into either a multiplicative or additive forms. The additive form in generating missing data.

The time series used in estimating factor models are either in levels or log levels. First differencing is the alternative format of the time series in our study. A few series (e.g., output, exchange rates, prices) enter in gap form. Gaps are estimated by averaging data obtained from applying four different filters, namely the Hodrick-Prescott filter, Hamilton's (2018) filter, Christiano and Fitzgerald's (2003) band pass filter.[23] Since the raw data come in various forms (i.e., level, percent, percent of GDP, index, etc.) all series are standardized before the factor models are estimated.

Data for the (nominal) policy rate series published by the BIS. The results section focuses on a discussion of the real neutral rate, or r*, since this is the measure typically mentioned in central bank policy discussions. That said, as is clear from (6) and (7) above, our original estimates are for R* and we then subtract proxies for expected inflation (see (8)). However, as Yardeni (2024) points out: "It is extremely doubtful that anyone based their economic decisions on an overnight bank lending rate that is adjusted for inflation measured on a year-to-year basis." Indeed, the widely followed US Federal Reserve's dot-plot (Summary of Economic Projections) provides economic projections for the nominal policy rate of the central bank.[24] Although we emphasize estimates of r*, the behavior of R* cannot be ignored (see equation (8)).

4. Empirical Results

Table 1 provides the number of estimated models considered and used in the application of the thick modelling approach. Sample periods for each country are also shown in the same Table.

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[22] Essentially, the technique fills missing data based on the previous two and following two non-missing values and fits the missing values to a non-linear, or curved, function.
[23] We also consider first differencing since this is also a frequently used filter to eliminate non-stationarity in the data. Our conclusions are unaffected by whether this form is included or not in calculating gaps. The same goes when the band pass filter estimates are omitted.
[24] The forward-looking fed funds rate series are only available since 2012.

Samples are primarily driven either by data availability or the desire to provide estimates covering a period when there is a largely unchanged policy regime. Hence, the data for Korea begin in 2000 since this is shortly after that country adopts inflation targeting (e.g., see Kim and Park (2006), Sanchez (2009)). In the case of China, both data availability, combined with a change in policy strategy that includes membership in the World Trade Organization in 2001, contribute to influencing the choice of the full sample period (also see Burdekin and Siklos (2025)).

Figure 1, and Tables 2 and 3 presents the main results with more results provided in Figures 2 to 4. The shaded areas of Figure 1 indicate recession dates. China is the only country that has avoided this condition throughout the sample. A separate appendix is available with additional evidence that space limitations prevented us from discussing below. For example, we relegate to the appendix band spectrum regressions using observed series since the conclusions presented below are unaffected. The same conclusion applies to the case when time-varying estimates are obtained although they are included in the calculation of the mean estimates of R* and r*.

Considerable discussion has been devoted to the implications of changing neutral real interest rates and this turn on the question whether the stance of monetary policy is 'too tight' or 'too loose' over time. Therefore, Figure 1 plots the differential between our estimate of $\hat{r}_{jt}^{*}$ (see (8)) and the real policy rate which is the observed policy rate less the same proxy for inflation expectations.[25] More formally, we can write

$$Monetary\ Policy\ Stance = \hat{r}_{jt}^{*} - r_{jt}, \frac{>0\ policy\ is\ too\ loose}{<0\ policy\ is\ too\ tight} \quad (9)$$

By construction, if the monetary policy stance indicator is positive this implies that monetary policy is too loose. Of course, the reverse result holds when the monetary policy stance indicator is negative in which case policy is too tight.

The point estimates shown in Figure 1 suggest two interesting dichotomies across the cycles and countries considered. First, other than for Japan and the US, the monetary policy stance measured over the financial cycle is relatively tighter than at the business cycle periodicities.

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[25] Hence, stated differently, equation (9) can be rewritten as the difference between R* and the observed nominal policy rate.

One interpretation is that in China and Korea, over a longer period, financial stability considerations led policy makers to perhaps devote more attention to the risks arising thereof relative to shorter-run considerations (i.e., at business cycle frequencies) where inflation control receives greater attention. That said, two other related points are worth noting. First, monetary policy gradually becomes closer to neutral after 2017 in both China and Korea. Second, the loosening impact of the GFC is especially visible for China and Korea at both the business and financial cycles. Also noticeable is the brief but sharp loosening in China in 2015 around the time of the stock market collapse and surprise devaluation of the renminbi (Burdekin and Siklos (2025)).

The second dichotomy one observes from Figure 1 is the persistence of relatively loose monetary policies in all four countries although this is less true in the case of the US. In the case of the US, the move to implement policies that are too loose, at least as measured by (9), is primarily a feature of crisis like conditions, including the 2001 dot-com bubble, the GFC and, more recently, the 2020-2021 pandemic.

The experiences of Japan and the US are of particular interest given the relative size and importance of these economies globally. The persistently loose monetary policy stance of the Bank of Japan is clearly visible, regardless of the cycle considered, until 2012 when, immediately after a recession, then Prime Minister Shinzo Abe introduces a series of monetary and fiscal reforms since called Abenomics. The impact on the stance of monetary policy suggests that the Bank of Japan first tightened and then would loosen its stance relative to the signal provided by the real policy rate, until the end of the pandemic and the global surge inflation where a sharp rightening is clearly visible. The same sharp tightening at the end of the sample is also visible in the US case.

The US stands out from the other three economies not only for the reasons already mentioned but because the Fed's monetary policy stance appears to display a form of mean reversion where the mean is dictated by the real federal funds rate. Overall, this suggests that monetary policy has been neutral more often and longer than in the other three countries. However, one cannot help but wonder whether this is also a reflection of the greater discretion practiced by the Federal Reserve over the sample considered. After all, the Fed's decisions tend to be influenced foremost by domestic considerations the role of international factors has waxed and waned over the decades (e.g., see Eichengreen (2013), and the Fed is not immune to the spillbacks from the rest of the world on the US economy (e..g, See Obstfeld (2020)). As noted previously, our samples are dominated by a succession of large crises, and this must also be factored in our interpretation of the conduct of US monetary policy.

Tables 2 and 3 present band spectrum regression estimates of equations (6) and (7). To gain some insights into the sensitivity of the results Table 3 omits a role for R* in the US and China, or China alone, from the specified regressions for China, Japan, and Korea. Hence, US results are not impacted relative to the case shown in Table 2. Table 2 shows the case when both the US and China impact changes in the policy rate for Japan and Korea while China's policy rate is allowed to be affected by the nominal neutral rate in the US.

Several results are worth highlighting. First, all estimated signs are consistent with theoretical expectations. Next, the impact of neutral rates in the US and China is significant everywhere with the smallest impact registered for China and only at the financial cycle periodicity while China's R* does not impact Japan's monetary policy at the financial cycle. The financial factor, often ignored in studies of the kind conducted here, is also seen to significantly impact policy rate changes in all four economies with, by far, the largest impact estimated for the US where the effect at the business cycle level is almost twice the size than the coefficient found for the financial cycle. The positive impact of the financial factor is also significant for China, and at both periodicities. However, the size of the impact on policy rate changes is only a fraction (20% or less) of the ones obtained for the US. While changes in financial conditions also matter for Japan and Korea, they are much larger than for China but still only just above half the size as the one estimated for the US, and they only influence monetary policy at the business cycle in Japan or at the financial cycle alone in the case of Korea.

Aggregate demand factors are important drivers of changes in the policy rate in three of the four countries examined with, by far, the largest coefficient estimated for Korea at both periodicities considered. Only in the case of Japan is no significant effect obtained which may not be surprising given the so-called 'lost decades' that the country has experienced (inter alia, see Callen and Ostry (2003), Nomura (2023)). The estimated aggregate supply factor is seen to influence monetary policy in the US alone and only at the business cycle frequencies. Whether this reflects the lately frequently discussed supply side factors in dictating inflation and monetary policy is unclear as the issue came to the attention of policy makers in the aftermath of the pandemic although supply side factors, at least in principle, were known to have the potential to influence inflation.[26] Similarly, the monetary factor is only seen as significantly contributing to policy rate changes in the US case and the size of the impact is essentially the same at the business and financial cycles. Finally, and with one exception, the structural factor tends to reduce the policy rate but only for Korea. This result may reflect the fact that this factor is a slow moving one, as noted earlier, and perhaps only potentially affects monetary policy at cycles longer than we are able to reliably estimate. In the case of Korea, it appears that the forces pushing up neutral rates, including productivity growth and savings rates are more than offset by the combination of an ageing population and possibly deterioration in income inequality. Of course, these same forces ought to play roles in the remaining three countries and while generally positive, the estimate coefficients are economically small and statistically insignificant.

Does the addition of global factors in the form of US and China's neutral rates matter for the remaining economies? Table 3 suggests that almost every conclusion obtained from Table 2 remains unchanged. The only differences are that ignoring the impact of China in the specification for Japan turns the aggregate supply factor statistically significant but economically small. Similarly, the monetary factor is significant for Japan at the financial cycle but is also economically small. Hence, it is safe to say that leaving out the effect of China's monetary policy results omits an important spillover effect. Moreover, inclusion of both US and China's influence on monetary policy in Japan and Korea especially suggests that spillover effects are quantitatively important in explaining policy rate changes.

We conclude the description of our results by exploring two other issues. First, we ask how our estimates of r* compare with others published in the literature. We restrict our analysis to the cases of Japan and the US since, to our knowledge, there are no easily obtained or updated estimates for China and Korea. The evidence is presented in Figure 2. Finally, while the literature has focused on point estimates of r* it is also well known, as discussed in section 2 above, that estimates are subject to considerable uncertainty. Accordingly, applying the principle

[26] The precise sources of the impact of supply side factors on monetary policy setting is outside the scope of this paper. See, however, Bai, Fernández-Villaverde, Li, and Zanetti (2024), Harding, Lindé, and Trabandt (2023)). China's influence on global inflation has also been widely studied. See, for example, Francis (2007), Kamin, Marazzi, and Schindler (2004), and references therein.

of thick modelling previously described, we plot the range of estimates of the monetary policy stance (i.e., equation (9)) in Figures 3 and 4. The top portion of the Figure displays data for Japan while the bottom portion considers the US experience. A priori, we don't know whether the estimates of the business or the financial cycle best describe the evolution of the monetary policy stance over time. Therefore, both estimates are plotted.

Overall, for Japan our estimates are largely comparable with those of the Bank of Japan (2025). That said, there are some differences. The change in the monetary policy stance towards loosening is much more abrupt for both our estimates than the ones the Bank has published. The same can be said when a new round of policy loosening takes place beginning in 2021 until the end of our sample (2023).[27] Furthermore, our estimates at the business cycle frequencies are much less volatile than any of the three other estimates shown.

If we now examine the US case the differences between our estimates and ones published by Holston, Laubach and Williams (2017, 2023) and Lubik and Matthes (2015) are noticeable. First, our estimates at both the business and financial cycle frequencies suggest that Fed policy was looser than if one relies on a semi-structural model of the US economy or a VAR-based approach to estimating r*. Nevertheless, the turning points in all series, including ours, are broadly similar over time. Indeed, the sharp loosening of policy in 2022 is observed and nearly identical for all four estimates shown. The same is also true at the beginning of the sample until 2002. The differences in the measure of policy looseness are most visible from the post-GFC until the COVID-19 pandemic begins.

A missing element in the results so far, including ones shown in Figure 2, is whether there are statistically meaningful differences not only in the various estimates shown but the level of uncertainty around our estimates. Accordingly, and relying on the thick modelling approach, Figure 3 plots the range of estimates obtained for all the estimated models at the business cycle frequency while Figure 4 repeats the exercise for the financial cycle estimates (also see Table 1). Visual comparisons suggest that estimates for the monetary policy stance at the business cycle frequencies for China are the most precise while the range of estimates we obtain can be wide at times for the other three economies. We also observe that the degree of uncertainty around

[27] The Bank of Japan began reversing course in July 2024, that is, is after our sample ends.

estimates of $\hat{r}_{jt}^{*} - r_{jt}$ is quite variable over time. Nevertheless, it is also observed that the monetary policy stance has been relatively tightest in China and Japan and loosest in Korea and the US. The picture is not very different when our attention turns to the estimates of equation (9) at the financial cycle. However, estimates of the monetary policy stance become much less precise for China and Korea while the range of estimates for Japan and the US suggest monetary policy stances that are tighter in Japan and the US at the financial cycle than when a business cycle perspective is adopted. In addition, estimates for Japan and the US appear more precise at the financial cycle than when $\hat{r}_{jt}^{*} - r_{jt}$ is estimated at business cycle periodicities.

The bottom line is that, as the literature surveyed earlier has also reported, there can be considerable uncertainty around point estimates of r*. As a result, the thick estimates shown in Figures 3 and 4 give less of an impression of the secular decline in neutral rates than suggested by point estimates, both ours and ones reported in the literature.

5. Conclusions

An unobservable, yet critical, variable that policy makers and academics rely on to assess the stance of monetary policy is the neutral interest rate. While there exist several published and updated estimates for the US, there is considerably less evidence for East Asia. Moreover, existing estimates tend to rely on semi-structural models for a closed economy which is not entirely appropriate for smaller or more open economies, or even emerging markets, that define many other economies. We revisit the experience of China, Japan, and Korea, alongside the US, to assess the evolution of the monetary policy stance over time in these countries. This implies producing estimates of the neutral rate and then comparing these with the observed policy rates of the central banks in question. We investigate two issues, namely whether there are differences over time in the conduct of monetary policy in the four countries considered in terms of the degree of looseness or tightness in their monetary policy. Next, we also consider the main drivers of changes in the neutral rates over time. However, unlike the extant literature, our strategy differs in two important ways. First, we estimate principal factors to summarize the roles of aggregate demand, supply, financial, monetary, and structural factors in influencing the neutral rate over time. Second, we use a time series approach that leverages the notion that the evolution of the neutral rates can be sensitive according to whether the horizon on question is the business cycle, defined as having a duration of 2 to 8 years, or the financial cycle with a cycle that is between 10 and 20 years long. We obtain our estimates using band spectrum regressions in part because we can estimate relationships in the frequency domain but present the results in the more traditional time domain.

We find that estimates, covering a sample ranging from 1996 to 2023, can be sensitive according to whether neutral rates are considered at the business or financial cycle perspectives except for the US where the two sets of estimates are more or less moving together over time. Estimates of the monetary policy stance suggests looseness in monetary policy is a persistent feature of policy since the late 1990s. However, the US is once again an exception with far more evidence of fluctuations around conditions of monetary policy neutrality. Japan is another example of this kind but only since Abenomics is introduced in 2012. Equally important is our finding that spillovers from the US and China are economically and statistically significant. As a result, rises or falls in policy rates in Japan and Korea are partly the result of similar changes in the US and China. While the US neutral rate also influences China's the impact is considerably smaller than in the two other Asian economies considered. Comparisons with other published and updated estimates for Japan and the US reveal that our estimates are broadly comparable with ones for Japan, with a few interesting differences, while our estimates for the US suggest monetary policy was tighter, from the end of the GFC until shortly before the pandemic, than the interpretation obtained from the published estimates of HLW (2017, 2023) and LM (2015). Nevertheless, our results have in common with others the potential for considerable uncertainty around point estimates.

Because our estimates may well be strongly influenced by the presence of large crises of the financial and non-financial varieties, a longer time span would be a desirable goal. Assuming our methodology is appropriate because we are able to determine the stance of policy depending on the policy horizon considered (i.e., business versus financial cycles) this would also enable a greater variety of definitions of cycles than we were able to implement here. Moreover, even though we rely on a fairly large number of time series to estimate the various factors considered, there is the potential for more time series to be added and this would also help generate more confidence in our estimates of the neutral rate. Finally, given that we successfully estimate neutral rates for a rather heterogeneous set of economies, a logical extension would be to apply our strategy to estimate neutral rates in other countries. We leave these extensions for future research.

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Tables and graphs

Sample and model information

Table 1

	CHN	KOR	JPN	USA
# estimated models[1]	16	24	10	14
Factor model-based[2] final samples	2002Q1-23Q4	2001Q1-23Q4	1995Q2-23Q4	1998Q1-23Q4
Observables[3] final samples	2002Q1-23Q4	2000Q1-23Q4	1996Q1-23Q4	1996Q1-23Q4

[1] Refers to the total number of models at both the business and financial cycle frequencies that are estimated. [2] The factor model estimates are ones defines in (6) and (7). [3] Observables refer to the output gap, inflation gap, credit gap, real exchange rate gap, TFP growth, fertility rate, and Global Supply Chain Index time series. The sampling frequency is quarterly.

Selected band spectrum regression estimates: equation (7)[1] Table 2

	USA		CHN		JPN		KOR	
Cycle	BC	FC	BC	FC	BC	FC	BC	FC
Sample	1998Q1-23Q4		2002Q1-23Q4		2004Q2-23Q4		2004Q1-23Q4	
Constant			-0.06** (.029)	-0.02 (.02)	0.42* (.14)	0.11 (.10)	-0.79* (.16)	-0.32* (.10)
Real (demand)	0.03 (0.10)	0.20** (.10)	0.42* (.11)	0.44* (.10)	-0.29 (.29)	0.34 (.37)	3.00* (.22)	2.83* (.32)
Real (Supply)	-0.11*** (.08)	-0.01 (.08)	-0.06 (.08)	0.02 (.08)	-0.25 (.32)	0.14 (.41)	0.10 (.15)	0.31 (.21)
Financial	2.87* (.50)	1.59* (.46)	0.27* (.11)	0.34* (.10)	0.92* (.32)	0.52 (.40)	0.22 (.28)	0.92* (.31)
Monetary	0.36* (.08)	0.33* (.07)	0.21 (.15)	-0.04 (.14)	-0.39 (.25)	0.31 (.36)	-0.13 (.24)	0.29 (.34)
Structural	0.03 (.17)	0.04 (.18)	0.01 (.01)	0.02 (.02)	0.10 (.09)	-0.01 (.14)	-0.53* (.07)	-0.48* (.10)
R* - USA	NA	NA	0.01 (.06)	0.12** (.06)	0.74* (.16)	0.44*** (.25)	0.42** (.17)	0.69* (.25)
R* - CHN	NA	NA	NA	NA	0.36*** (.21)	0.20 (.28)	0.58** (.26)	0.25 (.31)
$\bar{R}^2$	0.61	0.60	0.69	0.56	0.46	0.25	0.84	0.72
F-stat (p)	33.31 (.00)	32.28 (.00)	32.25 (.00)	19.42 (.00)	10.64 (.00)	4.78 (.00)	60.88 (.00)	29.91 (.00)

[1] The methodology is explained in the text but consists of Hannan Efficient estimates. BC represent the business cycle frequencies; FC represent the financial cycle frequencies. USA is the United States, CHN is China, JPN is Japan, and KOR is Korea. Standard errors are in parenthesis. F-stat (p) is the test of the joint statistical significance of the right-hand side variables and the associated p-value. The construction of the factors is also explained in the main body of the paper. */**/*** indicates significance at 1/5/10% confidence interval.

Selected additional band spectrum regression estimates: equation (6)[1] Table 3

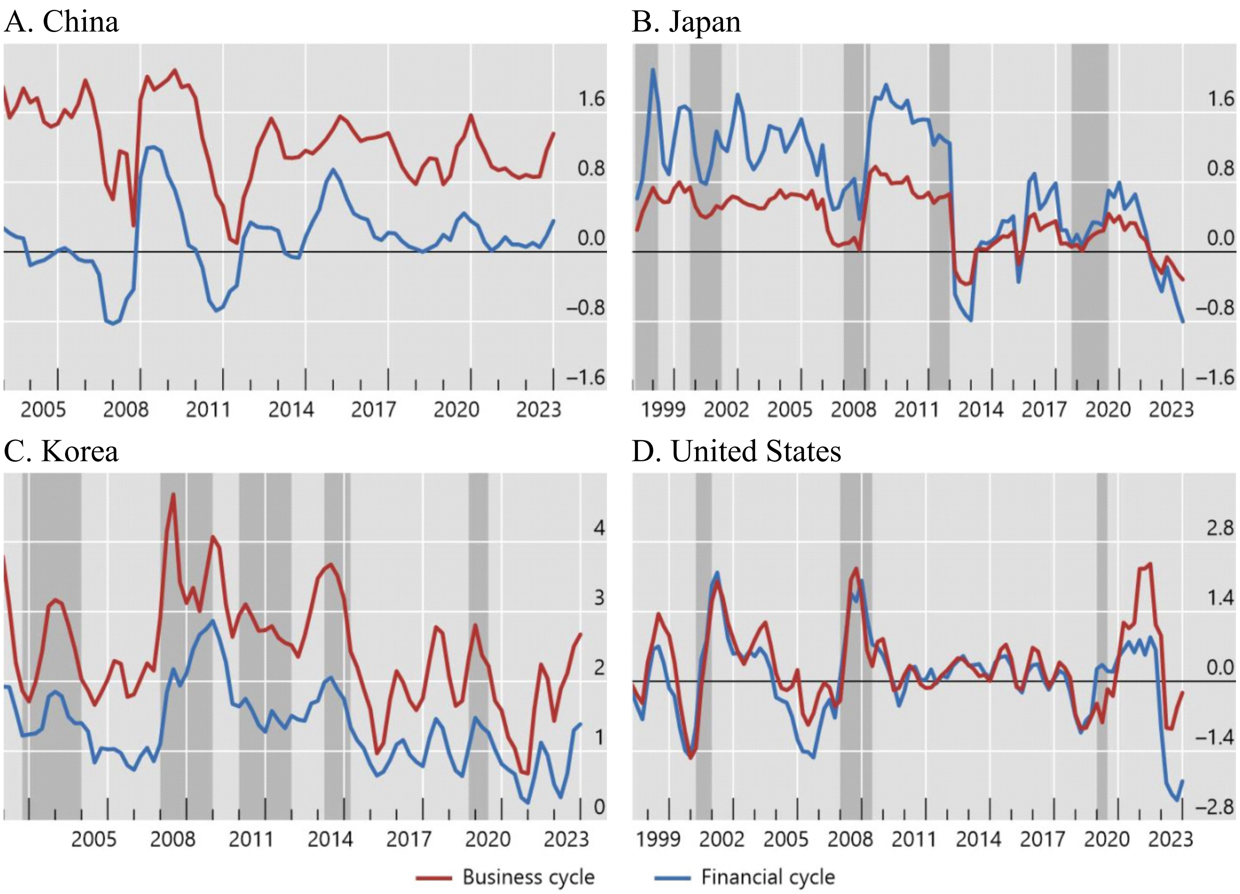
	JPN				KOR			
Cycle	BC	FC	BC	FC	BC	FC	BC	FC
Sample	1998Q1-23Q4		1995Q2-23Q4		2001Q1-23Q4		2001Q1-23Q4	
Constant	0.01 (0.01)	0.01 (.01)	-0.03** (.01)	0.22* (.03)	-0.31* (.05)	-0.13 (.03)	-0.35* (.05)	-0.15* (.03)
Real (demand)	0.03 (.03)	0.04 (.04)	0.03 (.07)	-0.09 (.11)	0.86* (.08)	0.84* (.10)	2.96* (.25)	2.86* (.33)
Real (Supply)	-0.09* (.03)	0.03 (.03)	-0.14** (.07)	-0.15 (.10)	0.06 (.05)	0.10 (.07)	0.28 (.16)	0.37 (.23)
Financial	0.12* (.03)	0.01 (.01)	0.77* (.07)	0.58* (.08)	0.09 (.09)	0.28* (.10)	0.68* (.27)	1.17* (.32)
Monetary	-0.005 (.03)	0.02*** (.01)	0.14 (.09)	0.43* (.12)	-0.11 (.08)	0.03 (.10)	-0.30 (.26)	-0.11 (.35)
Structural	0.02* (.01)	0.02 (.01)	0.14 (.09)	0.23 (.15)	-0.17* (.02)	-0.16* (.03)	-0.62* (.08)	-0.59* (.11)
R* - USA	0.01 (.01)	0.05** (.02)	NA	NA	0.14** (.06)	0.20* (.07)	NA	NA
R* - CHN	NA	NA	NA	NA	NA	NA	NA	NA
$\bar{R}^2$	0.34	0.18	0.58	0.43	0.79	0.67	0.77	0.64
F-stat (p)	9.95 (.00)	4.85 (.00)	32.32 (.00)	18.54 (.00)	53.97 (.00)	31.69 (.00)	60.76 (.00)	32.30 (.00)

[1] The methodology is explained in the text but consists of Hannan Efficient estimates. BC represent the business cycle frequencies; FC represent the financial cycle frequencies. USA is the United States, CHN is China, JPN is Japan, and KOR is Korea. Standard errors are in parenthesis. F-stat (p) is the test of the joint statistical significance of the right-hand side variables and the associated p-value. The construction of the factors is also explained in the main body of the paper. */**/*** indicates significance at 1/5/10% confidence interval.

Monetary policy stances: business cycle and financial cycle perspectives[1]

r* less real policy rate

Figure 1



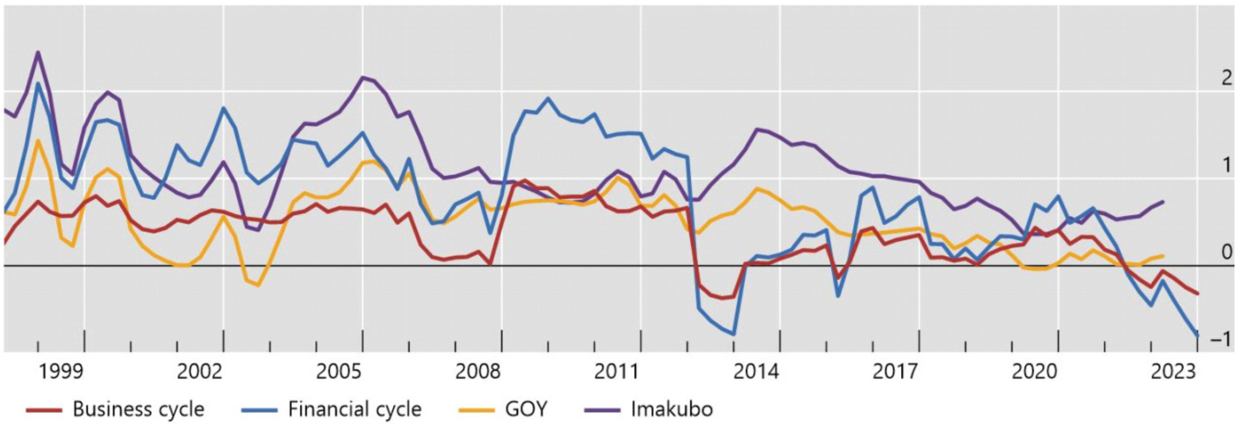
[1] Shown are point estimates of the evolution of the monetary policy stance as derived for the business cycle (BC) and financial cycle (FC) periodicities based on band spectrum regression estimates for the number of models shown in Table 1. The mean of the estimated values is plotted above for the four countries in our data set. Both full sample, time-varying and BSRs based on observables are included in the set of models used to calculate the means. The shaded areas are the recession periods. For the US the dates are those of the NBER; for Japan from the Japanese government cabinet office; for Korea from the OECD

Varieties of estimates of the monetary policy stance: Japan and the USA[1]

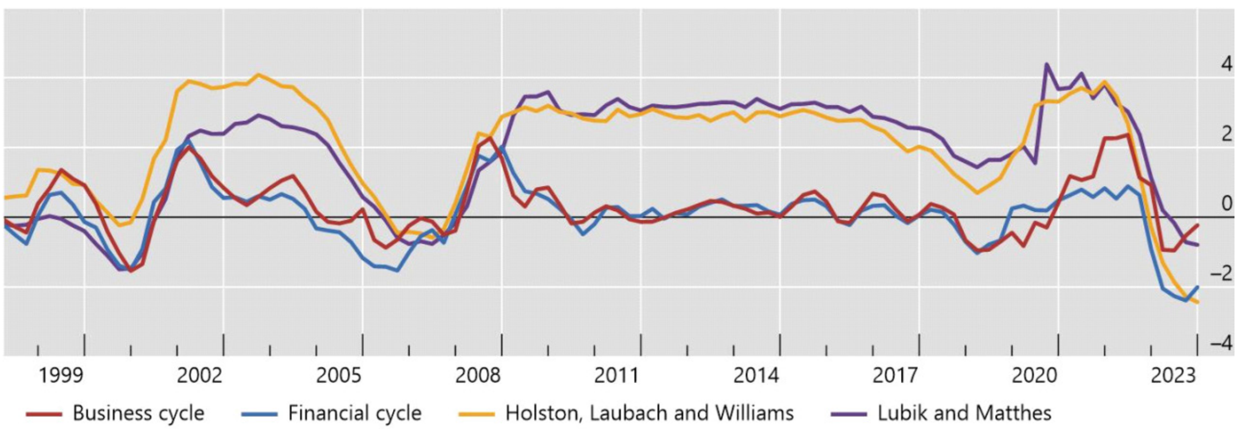
r* less real policy rate

Figure 2

A. Japan



B. United States

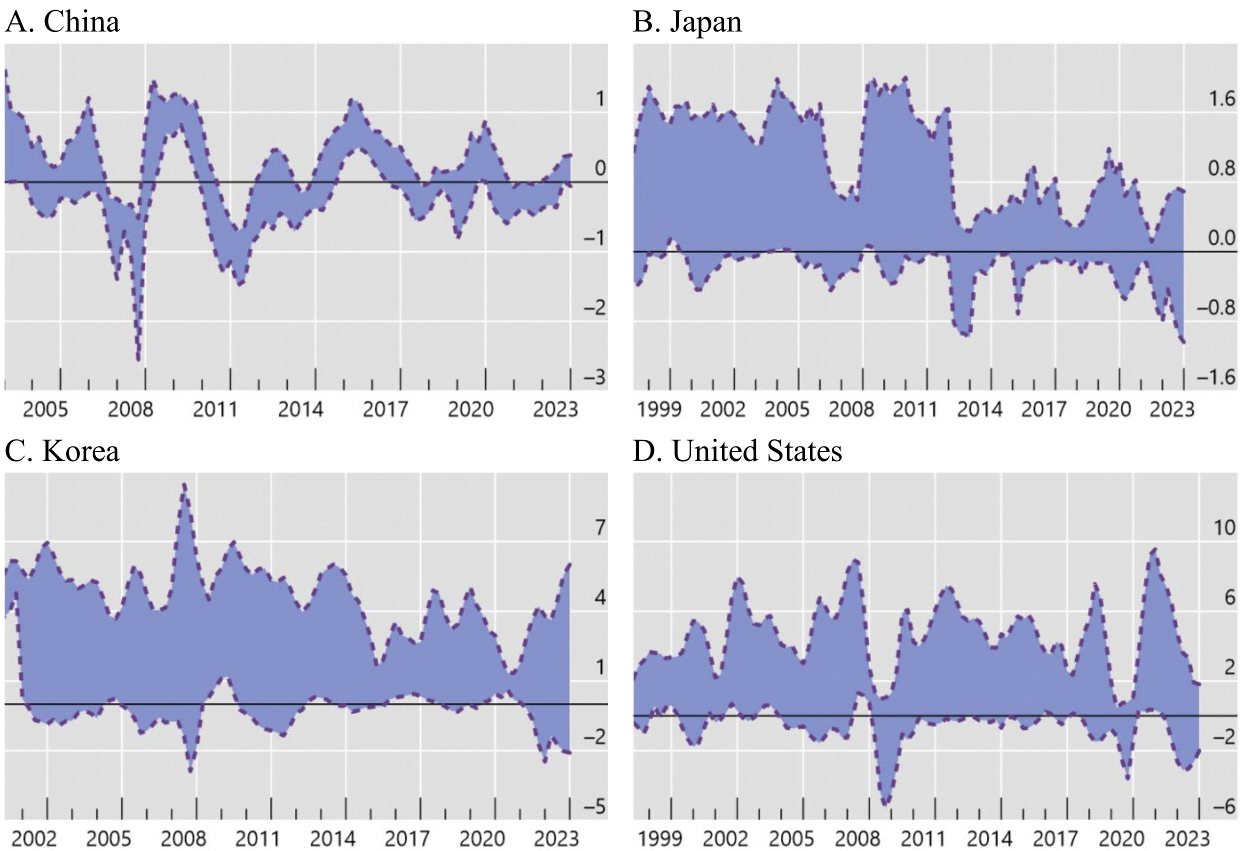


[1] Business cycle and financial cycle are from Figure 1. The remaining estimates are, for the USA: Holston, Laubach and Williams, and Lubik and Matthes from references cited in the main body of the paper and are available from the New York and Richmond Federal Reserve banks; for Japan the estimates of Goy and Imakubo are also from references cited in the text and can be obtained from the Bank of Japan. See equation (9) for the definition of the monetary policy stance. Positive values signal a policy that is too loose while a negative value suggests a policy that is too tight relative to the neutral rate.

Range of estimates of the monetary policy stance business cycle perspective[1]

r* less real policy rate

Figure 3

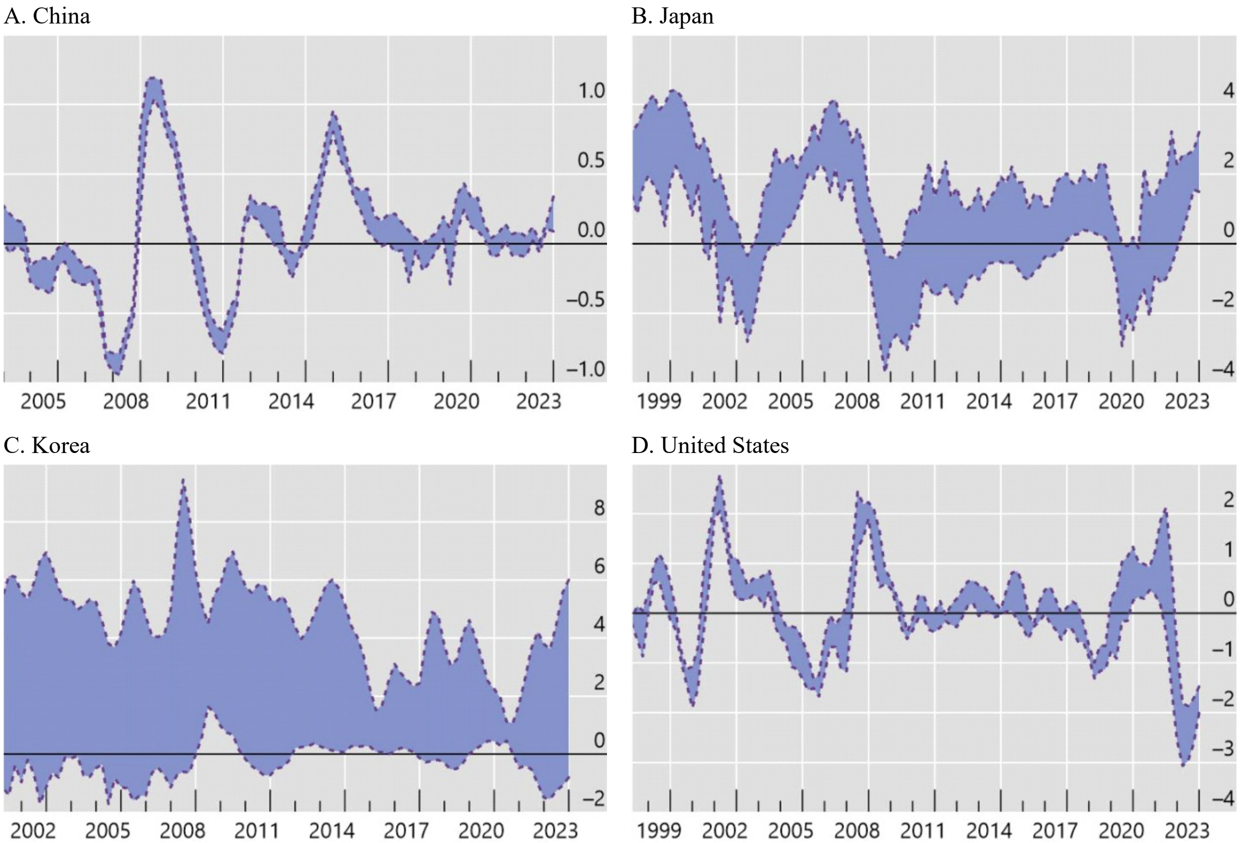


[1] The maximum and minimum estimates represent the range of estimates from all the estimated models (see Table 1) at the business cycle frequencies. Also, see note to Figure 1. . See equation (9) for the definition of the monetary policy stance. Positive values signal a policy that is too loose while a negative value suggests a policy that is too tight relative to the neutral rate.

Range of estimates of the monetary policy stance: financial cycle perspective[1]

r* less real policy rate

Figure 4



[1] See note to Figure 3. The range of estimates of for the financial cycle frequencies. . See equation (9) for the definition of the monetary policy stance. Positive values signal a policy that is too loose while a negative value suggests a policy that is too tight relative to the neutral rate.

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