Asymmetric Interest Rate Pass-through and Its Effects on Macroeconomic Variables: Evidence from Thailand

THOSAPON TONGHUI

School of Economics, Yonsei University, Seoul 03722, Korea
Email: thosaponth@gmail.com

JIN SEO CHO

School of Economics, Yonsei University, Seoul 03722, Korea
Email: jinseocho@yonsei.ac.kr

September 2023

Abstract

This study employs the two-step Nonlinear Autoregressive Distributed Lag estimation method proposed by Cho, Greenwood-Nimmo, and Shin (2019) to identify the asymmetric impact of monetary policy on economic variables using monthly data from Thailand between 2001 and 2023. The primary objective is to investigate the effects of policy rate shocks on the economy. This study examines three key aspects: (i) the asymmetric pass-through of policy rates to commercial bank deposits and loan rates; (ii) pass-through variations across bank sizes; and (iii) the asymmetric macroeconomic effects on output and inflation. The empirical findings reveal the presence of asymmetry within the relationships between the variables. First, the study identifies an incomplete interest rate pass-through with deposit rates ranging from 28.1% to 102.7% and loan rates ranging from 12.7% to 89.6%. Notably, long-term upward asymmetry is observed for loan rates, whereas the evidence for deposit rates is limited. Second, regarding bank size, large banks exhibit a greater pass-through effect on loan rates, whereas small and medium-sized banks display higher responsiveness to short-term deposits or savings rates. Finally, this study provides strong evidence of long-term asymmetric macroeconomic impacts. Quantitatively, rate hikes have a more substantial effect on output growth, being 1.4 times larger than the impact of rate cuts. Conversely, rate decreases exhibit a more pronounced effect on inflation, being 3.4 times larger than the impact of rate increases. These findings suggest the presence of downward price rigidity associated with monetary policy shocks in the context of Thailand.

Key Words: Interest rate pass-through; asymmetric impact; macroeconomic effects; Nonlinear Autoregressive Distributed Lag model.

Subject Classification: E43, E51, E52.

Acknowledgments: Cho acknowledges financial support from the Ministry of Education of the Republic of Korea and the National Research Foundation of Korea (NRF-2020S1A5A2A01040235).

1 Introduction

"*The universe is asymmetric and I am persuaded that life, as it is known to us, is a direct result of the asymmetry of the universe or of its indirect consequences. The universe is asymmetric.*"
— Louis Pasteur

Monetary policies have an asymmetric impact on the economy. For instance, Bernanke, Gertler, and Gilchrist (1996) note that financial frictions exist in the financial market, resulting in varied effects on the economy when policy interest rates change. Another example is the survey conducted by Égert and MacDonald (2009) which explores various monetary transmission mechanisms in Central and Eastern European financial markets, revealing their complexity, which cannot be comprehensively understood within a single framework.

Identifying asymmetric monetary policy effects is often the primary objective. However, when assuming the classical New Keynesian model and analyzing the impact of monetary policy, there can be bias, as this model assumes a symmetric effect of monetary policy on the economy. Nevertheless, the identification of asymmetry is constrained by the application of specific models designed for particular economies and a generic model assumption for identification has not yet been defined.

This study primarily examines the asymmetric effects of interest rate shocks on the economy. Our main focus is to analyze the differences in the effects of increases and decreases in policy rates. To accomplish this, we employ a Nonlinear Autoregressive Distributed Lag (NARDL) model to investigate the transmission of policy rates and their impact on macroeconomic variables. Specifically, we empirically apply the two-step NARDL (2SNARDL) estimation method proposed by Cho et al. (2019) to demonstrate its effectiveness in capturing asymmetric impacts. We utilize financial and macroeconomic data from Thailand for our analysis.

We achieve our goal by estimating the impact of policy rates on macroeconomic variables through pass-through. Initially, we examine the asymmetric pass-through of policy interest rates to commercial bank rates, a crucial step given the pivotal role of financial intermediaries in transmitting policy changes to real economic activities such as credit provision and savings accumulation. We operate under the assumption that interest rate changes exhibit an asymmetric pass-through, drawing on the foundational mark-up pricing model of Rousseas (1985). Subsequently, we explore the variations in pass-through across banks of different sizes. Finally, we investigate how key macroeconomic variables, including inflation and economic growth, respond to changes in interest rates. This analysis considers the economic indicators of trading partners and global energy prices to provide a comprehensive understanding of these relationships. Consistent with recent literature, we anticipate asymmetric effects of interest rate shocks on these macroeconomic variables.

Recent studies have challenged the traditional view of a complete and symmetric monetary policy pass-through by introducing the concepts of incomplete and asymmetric pass-through (e.g., see Gregor, Melecký, and Melecký, 2021). As discussed in Section 2.1, various factors influence interest rate pass-through, including market structure and competition (e.g., Hannan and Berger, 1991; Gigineishvili, 2011), economic uncertainty (e.g., Égert and MacDonald, 2009), liquidity conditions (e.g., Gigineishvili, 2011; Cho, Greenwood-Nimmo, and Shin, 2023b), financial frictions (e.g., Bernanke et al., 1996), asymmetric information, and risk perception (e.g., Stiglitz and Weiss, 1981). Moreover, the direction of policy rate adjustments also plays a significant role (e.g., Hannan and Berger, 1991; De Bondt, 2005). In addition, two distinct lines of research address the asymmetric macroeconomic impacts of interest rates. The first strand examines state-dependent asymmetry to differentiate between the effects of interest rate shocks during economic slowdowns and expansions (e.g., see Lo and Piger, 2005; Peersman and Smets, 2005). The second focuses on the direction of interest rate changes. For example, Debortoli, Forni, Gambetti, and Sala (2020) and Hayford (2006) reveal that, in the case of the U.S., increases in interest rates have more pronounced effects on output than rate cuts. The latter is the primary focus of this study.

In terms of methodology, we choose the NARDL model because of its practicality, simplicity, and effectiveness in analyzing asymmetric phenomena, as highlighted by Cho, Greenwood-Nimmo, and Shin (2023a). This choice diverges from most existing literature on interest rate pass-through, which relies primarily on error-correction models (ECMs), dynamic ordinary least squares (OLS), and autoregressive distributed lag (ARDL) models. These conventional models have limited capacity to capture the asymmetrical effects of interest rate changes. The NARDL model offers several advantages, including the use of partial sum decompositions for asymmetry analysis, accommodation of cointegration models for non-stationarity and nonlinearity, and the ability to construct interpretable visualizations using cumulative dynamic multipliers. Consequently, the NARDL model has gained popularity in the literature for analyzing both interest rate pass-through (e.g., Greenwood-Nimmo, Shin, van Treeck, and Yu, 2013; Yu, Chun, and Kim, 2013; Apergis and Cooray, 2015; Galindo and Steiner, 2022) and the macroeconomic effects of interest rate shocks (e.g., Adelakun and Yousfi, 2020; Gocer and Ongan, 2020; Claus and Nguyen, 2020).

Typically, NARDL models are estimated using a single-step method with OLS estimation as proposed by Shin, Yu, and Greenwood-Nimmo (2014). However, we employ the 2SNARDL method introduced by Cho et al. (2019). We chose this because we recognize that the single-step method may encounter issues related to the asymptotic singular matrix problem. The 2SNARDL model overcomes this challenge and offers analytical tractability. In summary, the 2SNARDL approach initially estimates the long-run relationship using the fully modified OLS (FM-OLS) estimator of Phillips and Hansen (1990). Subsequently, we estimate the short-run dynamic parameters using OLS, as detailed in Section 3.

To the best of our knowledge, research on interest rate pass-through in Thailand is limited in terms of both the number of studies and the utilization of asymmetric methodologies. Existing studies, which have predominantly relied on symmetrical assumptions, consistently reveal incomplete pass-through. For instance, during the sample period from 1997 to 2006, both the deposit and loan rates were estimated to exhibit a pass-through rate of less than 50% (e.g., Disyatat and Vongsinsirikul, 2003; Charoenseang and Manakit, 2007). More recently, Yu et al. (2013) explore the asymmetric interest rate pass-through using Thai data from 2000 to 2009, revealing a notable asymmetry in lending rates. However, it is important to note that these earlier studies were conducted during a period marked by substantial excess liquidity, which may have contributed to the observed low pass-through rates (see the Bank of Thailand, Monetary Policy Department, 2008). Therefore, it is imperative to conduct a comprehensive and up-to-date analysis.

Our contributions to the existing literature include the application of 2NARDL to Thai data spanning 2001–2023, providing up-to-date insights into the effects of interest rate dynamics, particularly in Thailand. Additionally, we uncover variations in pass-through levels across banks of different sizes and address the gap in the limited studies that use Thai data to explore the asymmetric macroeconomic effects of interest rates. Finally, we demonstrate the use of the 2SNARDL method to reveal the impact of an asymmetric pass-through on macroeconomic variables, offering a valuable approach that can be applied to other countries.

The remainder of this paper is organized as follows. Section 2 summarizes the related theoretical background and empirical studies along with the motivation for this research. Section 3 introduces the details of the NARDL model. Section 4 provides insights into the data used in the study. We present the empirical results in Section 5, and Section 6 concludes. In the Appendix, we include the preliminary testing results for the application of the 2SNARDL model to the data used in this study.

2 Literature Review and Motivation

In this section, we review the development of the prior literature and motivation for the current study. After reviewing the theoretical literature, we review the empirical literature separately.

2.1 Theoretical Literature Review

For the goal of this study, it is essential to understand the interest rate channel thoroughly within the information technology (IT) framework. It is well known that monetary policy affects the economy through four channels: exchange rates, interest rates, asset prices, and credit (e.g., see Mishkin, 1996). Before discussing incomplete and asymmetric interest rate pass-through, we briefly review the literature on interest channels.

The central bank employs operational tools such as adjusting short-term money market rates in accordance with decisions made by the Monetary Policy Committee (MPC) with the primary goal of achieving price stability. These adjustments also influence the long-term money market rates. The standard cost-of-funds model, initially proposed by Rousseas (1985) and summarized by De Bondt (2002), posits a positive correlation between retail bank rates and policy-controlled interest rates. Policy-controlled interest rates are often approximated using central bank policy and money market rates. This link is described using a markup pricing model:

$$r_t^B = \alpha_* + \beta_* r_t^M, \tag{1}$$

where r_t^B represents the interest rates set by commercial banks, such as deposit interest rates and bank lending rates; α_* is the constant markup; β_* is the degree of long-run interest rate pass-through; and r_t^M is the marginal cost approximated by the policy rate.

Next, policy rates influence the investment and savings decisions of private agents through lending and deposit rates, thereby adjusting consumption and investment and ultimately affecting aggregate demand and prices. Clarida, Galí, and Gertler (1999) provide a macroeconomic model describing the connections among interest rates, output, and inflation in the New Keynesian framework. In the basic setting, this relationship involves three primary equations that align with the Aggregate Supply and Aggregate Demand models:

$$\text{NKPC} \qquad \hat{\pi}_t = \beta E_t \hat{\pi}_{t+1} + \kappa \hat{y}_t + \epsilon_t^s \tag{2}$$

$$\text{Dynamic IS} \qquad \hat{y}_t = E_t \hat{y}_{t+1} + \sigma^{-1}(\hat{r}_t - E_t \hat{\pi}_{t+1}) + \epsilon_t^d \tag{3}$$

$$\text{MP Schedule} \qquad \hat{r}_t = \phi^\pi \hat{\pi}_t + \phi^y \hat{y}_t + \epsilon_t^R \tag{4}$$

where $\hat{\pi}_t$ is the current inflation rate, $\hat{y}_t$ is the output gap, $\hat{r}_t$ is the interest rate, $\hat{r}_t - E_t \hat{\pi}_{t+1}$ is the real interest rate, ϵ_t^s is a cost-push shock, ϵ_t^d is the demand shock, and ϵ_t^R is a monetary policy shock.

The economic intuition of this framework is as follows. First, the New Keynesian Phillips Curve (NKPC) plays the role of aggregate supply. This signifies that current inflation depends on anticipated future inflation, the output gap, and other economic shocks. The NKPC is typically derived from firms' pricing decisions, assuming that producers have monopoly power. This enables them to set prices above marginal costs, leading to price stickiness, as firms refrain from immediately adjusting their prices in response to shifts in economic conditions. Second, the dynamic investment-saving curve illustrates how private agents decide consumption and investment based on their forward-looking economic outlook on the demand side. This involves the intertemporal allocation of household resources, determined through household optimization. Specifically, it demonstrates that current output relies on expected future output, the real interest rate, and anticipated future inflation. Finally, the monetary policy schedule represents how interest rates react to fluctuations in inflation and the output gap. Monetary policy rules govern this response, typically the Taylor rule. For instance, if the inflation rate exceeds the target and output lags behind its potential, the central bank can raise interest rates to combat inflation and stabilize the economy.

This framework offers a modern perspective on the role of central banks in macroeconomics, emphasizing short-term nominal interest rates over aggregate money supply. Its applications have been widely reported in the literature. For instance, by referring to these equations and assuming sticky prices, Poutineau, Sobczak, and Vermandel (2015) demonstrate that monetary policy shocks, realized through an increase in nominal interest rates, depress aggregate demand, leading to reduced output. This outcome arises from households' delayed consumption driven by the consumption-smoothing mechanism. Consequently, declining demand can contribute to deflation. As time progresses, the economy begins to recover from these shocks, and a subsequent decrease in nominal interest rates occurs after the initial shock period, following Taylor's rule.

Despite its popularity, it has been noted that the effectiveness of the model proposed in Clarida et al. (1999) depends on the speed and extent to which changes in policy interest rates affect retail interest rates. Particularly, when the policy interest rate changes quickly and has a full impact on the retail rate, it can swiftly influence domestic demand and inflation (see Grigoli and Mota, 2017).

Recent studies have introduced the concept of incomplete and asymmetric pass-through, challenging conventional monetary theory, which assumes that pass-through is symmetrical, rapid, and complete. Comprehensive surveys conducted by Andries and Billon (2016) and Gregor et al. (2021) identify several factors that influence the magnitude of pass-through. These factors include marginal pricing costs, the bank's market power (or degree of competition), its balance sheet, liquidity constraints, the development of financial markets, and prevailing economic conditions. In addition to these factors, Grigoli and Mota (2017) provides a theoretical background to explain the stickiness and asymmetry of retail rate adjustments in response to policy rate changes. These phenomena are attributed primarily to incomplete market competition and information asymmetry.

These factors can be summarized as follows:

(a) First, market structure and the degree of competitiveness significantly influence the interest rate channel. Banks may have more power in markets characterized by limited competition. Consequently, they tend to exert greater control over interest rates, making them less responsive to policy-rate changes. As shown by Gigineishvili (2011), the level of competition among banks can enhance the degree of pass-through. Furthermore, Hannan and Berger (1991) investigates price rigidity in the banking industry and concludes that banks operating in more concentrated markets are less likely to adjust their deposit rates in response to changes in security rates.[1]

(b) Second, economic uncertainty plays a significant role in the interest rate channel. As emphasized by Égert and MacDonald (2009), macroeconomic conditions have a notable impact on the stickiness of retail rates. During periods of economic uncertainty or instability, banks often adopt a more cautious stance and are less inclined to adjust their interest rates in response to policy changes. This cautious approach can result in incomplete pass-through. Additionally, pass-through rates are likely to be higher during periods of high inflation because prices are adjusted more frequently.

(c) Third, liquidity conditions also play a significant role. The availability of funds in money markets and the overall prevailing liquidity conditions can influence the speed at which adjustments in policy rates are transmitted throughout the financial system. This phenomenon is intrinsically linked to credit markets because it depends on how monetary policy affects the broader liquidity landscape, cost of funds, and cost of credit (see Cho et al., 2023a). Empirically, Gigineishvili (2011) finds that an abundance of liquidity can impede the extent of transmission.

(d) Fourth, asymmetric information and risk perception are other factors affecting the interest rate channel. As emphasized by Stiglitz and Weiss (1981), asymmetric information and risk perception play significant roles in the stickiness of retail rates. When banks perceive an increased risk of default, they tend to maintain a spread between the lending and deposit rates. Moreover, they may resist raising their rates even when their funding costs increase to mitigate the heightened default risk within their loan portfolio. This implies an upward rigidity in loan rate adjustments. This finding supports De Bondt (2005), who also suggests that banks may be reluctant to fully transmit an increase in policy rates to mitigate credit risk.

(e) Fifth, financial frictions can have a notable impact on the interest rate channel. As pointed out by Bernanke et al. (1996), financial frictions are linked to the vicious-cycle effect of interest rate changes on the economy, primarily through their influence on firms' balance sheets and credit conditions. These factors contribute to the stickiness of retail rates.

(f) Finally, empirical studies have highlighted that the direction in which policy rates are adjusted and the type of interest rate can also impact the interest rate channel. Practically, when loan rates increase, banks may face the risk of losing customers, leading to rigidity in upward adjustments. As shown by Hannan and Berger (1991), deposit rates tend to exhibit more rigidity in response to increases than to decreases. Regarding lending rates, a cross-country analysis conducted by Borio and Fritz (1995)

[1] The study includes 398 bank data in 132 markets and measures security-rate by three-month T-bill rates.

demonstrates that the adjustment of lending rates is swifter in response to upward rate changes in certain countries.

2.2 Empirical Literature Review

Based on theoretical studies of asymmetric interest rate pass-through, numerous further studies investigate its properties both qualitatively and quantitatively. These studies aim to elucidate how long-term retail interest rates, including lending and deposit rates, respond to policy rate changes by estimating the empirical models stemming from (1). Empirical evidence on this topic may vary depending on factors such as bank location, sample period, and estimation methods used. Most studies apply cointegration analysis in the form of ECMs, dynamic OLS, and ARDL models. It is commonly observed that the pass-through process is incomplete, implying that the adjustment in retail rates in response to policy-rate changes is less than unity. According to Gregor et al.'s (2021) meta-analysis of 54 studies, the long-run pass-through coefficient is estimated at 0.803 with a median of 0.854.

Recent studies have begun to recognize the nonlinear cointegration structure in (1) and estimate asymmetric models of monetary policy pass-through. For instance, Greenwood-Nimmo et al. (2013) apply the NARDL to U.S. data to examine how banks adjust their deposit and lending rates in response to changes in policy-controlled rates.[2] Their results provide strong evidence of asymmetry. Specifically, they find that, in the long run, rate cuts are passed through more completely than rate hikes, whereas, in the short run, rate hikes are transmitted more rapidly than rate cuts. Additionally, the degree of pass-through diminishes during great moderation. Similarly, Apergis and Cooray (2015) apply NARDL to data from the U.S., the U.K., and Australia, observing positive asymmetry for lending rates and negative asymmetry regarding deposit rates. In other words, decreases in deposit interest rates exhibit a greater interest rate pass-through, whereas increases in lending rates demonstrate greater pass-through. Furthermore, Galindo and Steiner (2022) examine the case of Colombia and suggest upward rigidity in lending rates, indicating that they respond more to rate cuts than to hikes. This finding suggests that financial intermediaries are more reluctant to raise rates than to decrease them.

In terms of empirical studies on Thailand, there have been relatively few investigations of interest-rate pass-through. According to these studies, the degree of pass-through has declined since the 1997 financial crisis, ranging from 27% to 39% for lending rates and from 22% to 50% for savings rates. For instance, Disyatat and Vongsinsirikul (2003) examine the degree of the pass-through of money market rates to the minimum lending rate (MLR) and three-month deposit rates using a dynamic multiplier (DM) model and

[2] For details on NARDL, refer to Section 3.

an ECM with Thai data from 1989 to 2002. In the pre-crisis 1997 samples, deposit rates are adjusted by approximately 70% and 50% in response to RP14 rate changes by DM and ECM, respectively, while lending rates are adjusted by 56% (DM) and 40% (ECM). However, the results are considerably lower for post-crisis samples with deposit rates adjusted by 43% (DM) and 35% (ECM) and MLR rates adjusted by 39% (DM) and 36% (ECM). Similarly, Charoenseang and Manakit (2007) examine pass-through to various retail rates, such as MLR, minimum retail rate (MRR), savings rate, and 3- and 12-month deposit rates, employing ECM using observations from 2000 onwards, coinciding with the adoption of IT monetary policy by the MPC. Their findings indicate that commercial banks charged higher loan markups than deposits, resulting in an incomplete pass-through. Specifically, the average pass-through for the loan rate is 26.5%, whereas that for the savings deposit rate is 21.9%. Furthermore, the pass-through rate is higher for long-term rates, exceeding 50%.

In another empirical study on asymmetric pass-through, Yu et al. (2013) examine the asymmetric interest rate pass-through in Asian countries, including Thailand, using the NARDL model applied to data between 2000 and 2009. The results for Thailand indicate that the short-term market rate is cointegrated with the bank lending rate. The estimated coefficients suggest a pass-through degree ranging from 69% to 82%, assuming long-term symmetry. However, when assuming an asymmetric pass-through over the long run, positive increases in market rates pass through at approximately 76% to 77%, whereas negative changes pass through more than one-to-one at approximately 103%. This implies that Thailand's lending rates exhibit a negative asymmetric pass-through in the long run.

The macroeconomic effects of monetary policy shocks have also been studied extensively, focusing on both output and inflation. These empirical studies conventionally adopt the New Keynesian framework and rely on econometric models that assume symmetric pass-through. These include the vector autoregression (VAR) (e.g., Christiano, Eichenbaum, and Evans, 1998), sign restriction VAR (e.g., Ahmadi and Uhlig, 2015), and factor-augmented VAR (e.g., Bernanke, Boivin, and Eliasz, 2005).

Recent studies empirically reveal the presence of asymmetric effects resulting from monetary policy shocks by considering factors such as confidence levels, wage rigidity, prevailing economic conditions, credit conditions, and price inflexibility (see Pierre-Richard, 2001). Two distinct lines of research are noteworthy. The first examines the asymmetry across different economic states, showing that monetary policy shocks have a more pronounced influence on economic activity during economic slowdowns than during expansions. This phenomenon has been supported by Lo and Piger (2005) and Peersman and Smets (2005), who explore data from the U.S. and Europe, respectively. The second strand explores the potential asymmetric impact of interest rate changes on economic variables with a specific focus on the direction of interest rate changes. For example, Debortoli et al. (2020) apply SVAR with exogenous variables to U.S. data and find that, during easing periods, shocks have more pronounced effects on prices, whereas rate hikes exhibit more substantial effects on output. This finding aligns with the conclusions of Hayford (2006), who suggests that positive shocks to the federal funds rate have a greater impact on real GDP growth than negative shocks.

The NARDL model has gained popularity for investigating asymmetric macroeconomic effects linked to monetary policy. For example, Adelakun and Yousfi (2020) apply the NARDL model to South African data and find that positive monetary policy shocks have a more significant effect on output growth, whereas the results are reversed for inflation. Furthermore, the asymmetric impact of monetary policy shocks intensifies when government borrowing is considered. In a separate investigation, Gocer and Ongan (2020) examine U.K. data to explore the Fisher effect. They identify an asymmetric coefficient relationship between inflation and interest rates from 1995 to 2008, although the evidence weakens after 2008. In addition, Claus and Nguyen (2020) utilize the NARDL model on Australian consumer survey data to analyze consumer responses to monetary policy shocks. Their findings provide substantial evidence of asymmetric reactions in both short and long runs.

These developments in the literature motivated us to investigate the Thai economy using NARDL in terms of interest rate pass-through and the macroeconomic impact of monetary policy shocks. Specifically, we explore the asymmetric features of the interest rate channel while considering the unique aspects of the Thai economy. For this purpose, we consider the bank size effect in Thailand as detailed below.

3 Methodology

Before examining the Thai economy using the NARDL, we first briefly outline the NARDL model and its estimation procedures. The development of NARDL model estimation is comprehensively summarized by Cho et al. (2023a). The NARDL model builds upon a foundational understanding of analyzing autocorrelated trend-stationary processes. It combines autoregressive (AR) components with distributed lag components or sets of explanatory variables. Over time, various variants of the ARDL model have emerged, beginning with Pesaran and Shin (1999) and Pesaran, Shin, and Smith (2001), who assume a linear long-run relationship. This is followed by the NARDL model introduced by Shin et al. (2014), who simultaneously estimate long- and short-run parameters using the OLS method. This approach allows for the partial sum decompositions of explanatory variables, accommodating asymmetrical relationships. In addition, Cho et al. (2019) propose a two-step estimation method to address the asymptotic singularity problem that arises when an OLS estimation is applied.

There are several advantages of using the NARDL model. First, researchers can use partial sum decompositions of the explanatory variables by employing the NARDL model, which facilitates the analysis of asymmetry. In terms of econometric properties, NARDL is a cointegration model that enables researchers to jointly analyze data with nonstationary and nonlinear issues, such as mixing I(1) or I(0) data. Second, as highlighted by Shin et al. (2014) and Cho et al. (2023a), NARDL allows researchers to construct cumulative DMs, providing easily interpretable visualizations of dynamic adjustment toward an equilibrium position following a shock. Moreover, the cumulative DMs do not rely on controversial procedures to identify structural shocks.

The NARDL model captures a nonlinear cointegrating relationship by decomposing the positive and negative influences of exogenous variables. Suppose that we have $\boldsymbol{x}_t \in \mathbb{R}^k$ as exogenous independent variables ($t = 1, 2, \ldots, T$). The positive partial sum can be denoted as $\boldsymbol{x}_t^+ := \sum_{j=1}^t \Delta \boldsymbol{x}_j^+$, while the negative partial sum is denoted by $\boldsymbol{x}_t^- := \sum_{j=1}^t \Delta \boldsymbol{x}_j^-$ where $\Delta \boldsymbol{x}_j^+ := \max[\mathbf{0}, \Delta \boldsymbol{x}_j]$ and $\Delta \boldsymbol{x}_j^- := \min[\mathbf{0}, \Delta \boldsymbol{x}_j]$. That is, $\boldsymbol{x}_t$ decomposes into $\boldsymbol{x}_t^+$ and $\boldsymbol{x}_t^-$ at approximately zero. We assume that the following NARDL relationship holds between y_t and $\boldsymbol{x}_t$:

$$y_t = \gamma_* + \sum_{j=1}^{p} \phi_{j_*} y_{t-j} + \sum_{j=0}^{q} \left(\boldsymbol{\theta}_{j*}^{+\prime} \boldsymbol{x}_{t-j}^+ + \boldsymbol{\theta}_{j*}^{-\prime} \boldsymbol{x}_{t-j}^- \right) + \epsilon_t. \tag{5}$$

Here, ϕ_{j_*} is the AR parameter, $\boldsymbol{\theta}_{j*}^{+\prime}$ and $\boldsymbol{\theta}_{j*}^{-\prime}$ are the asymmetric distributed-lag parameters, and ϵ_t is an independently and identically distributed (iid) error term with zero mean and constant variance. Note that $\boldsymbol{x}_t = \boldsymbol{x}_0 + \boldsymbol{x}_t^+ + \boldsymbol{x}_t^-$ and suppose that $\Delta \boldsymbol{x}_t$ is a strictly stationary process.

The above equation captures the asymmetric effects between y_t and $\boldsymbol{x}_t$ through the different parameters $\boldsymbol{\theta}_{j*}^{+\prime}$ and $\boldsymbol{\theta}_{j*}^{-\prime}$. If y_t is cointegrated with $\boldsymbol{x}_t^+$ and $\boldsymbol{x}_t^-$, then $u_{t-1} := y_{t-1} - \boldsymbol{\beta}_*^{+\prime} \boldsymbol{x}_{t-1}^+ - \boldsymbol{\beta}_*^{-\prime} \boldsymbol{x}_{t-1}^-$ becomes the cointegration error term. We can rewrite (5) in the ECM form:

$$\Delta y_t = \rho_* u_{t-1} + \gamma_* + \sum_{j=1}^{p-1} \varphi_{j_*} y_{t-j} + \sum_{j=0}^{q-1} \left(\boldsymbol{\pi}_{j*}^{+\prime} \boldsymbol{x}_{t-j}^+ + \boldsymbol{\pi}_{j*}^{-\prime} \boldsymbol{x}_{t-j}^- \right) + \epsilon_t \tag{6}$$

where u_{t-1} is a stationary process possibly correlated with $\Delta \boldsymbol{x}_t$, while $\boldsymbol{\beta}_*^{+\prime} := -\left(\boldsymbol{\theta}_*^{+\prime}/\rho_*\right)$ and $\boldsymbol{\beta}_*^{-\prime} := -\left(\boldsymbol{\theta}_*^{-\prime}/\rho_*\right)$ are the asymmetric long-run parameters, and $\boldsymbol{\pi}_{j*}^{+\prime}$ and $\boldsymbol{\pi}_{j*}^{-\prime}$ are the parameters capturing nonlinear short-run dynamics.

Because of the singularity problem that arises when estimating unknown parameters using OLS, we employ the methodology developed by Cho et al. (2019), referred to as 2SNARDL estimation. The 2SNARDL estimation consists of two steps. First, when $k = 1$, it estimates the re-parameterized long-run relationship as

$$y_t = \alpha_* + \boldsymbol{\lambda}_*' \boldsymbol{x}_t^+ + \boldsymbol{\eta}_*' \boldsymbol{x}_t + u_t. \tag{7}$$

This estimation is performed using the FM-OLS estimator developed by Phillips and Hansen (1990). Here, $\boldsymbol{x}_t \equiv \boldsymbol{x}_t^+ + \boldsymbol{x}_t^-$, $\boldsymbol{\lambda}_* = \boldsymbol{\beta}_*^+ - \boldsymbol{\beta}_*^-$, and $\boldsymbol{\eta}* = \boldsymbol{\beta}_*^-$. For the long-run parameters, it follows that $\boldsymbol{\beta}_*^+ = \boldsymbol{\lambda}_* + \boldsymbol{\eta}_*$ and $\boldsymbol{\beta}_*^- = \boldsymbol{\eta}_*$.

This procedure aims to achieve a consistent estimation of the cointegrating equation while eliminating any singularity issues that may arise from collinearity between the positive and negative cumulative partial sums of the explanatory variable. Furthermore, it leverages the properties of FM-OLS, which is free from asymptotic bias when confronted with endogenous regressors and/or serial correlation, and follows an asymptotic mixed normal distribution.

If there are multiple exogenous variables, such as $k > 1$, Cho et al. (2019) reparameterize the long-run equation as follows:

$$y_t = \boldsymbol{\beta}_*^{+\prime}\boldsymbol{m}_t^+ + \boldsymbol{\beta}_*^{-\prime}\boldsymbol{m}_t^- + \delta_* t + \alpha_* + u_t.$$

The reparameterization is based on the following relationships:

$$\boldsymbol{x}_t^+ = \boldsymbol{\mu}_*^+ t + \boldsymbol{m}_t^+ \quad \text{and} \quad \boldsymbol{x}_t^- = \boldsymbol{\mu}_*^- t + \boldsymbol{m}_t^-, \tag{8}$$

where $\boldsymbol{\mu}_*^+ := \mathbb{E}[\Delta \boldsymbol{x}_t^+]$, $\boldsymbol{\mu}_*^- := \mathbb{E}[\Delta \boldsymbol{x}_t^-]$, $\boldsymbol{m}_t^+ := \sum_{j=1}^t (\max[\boldsymbol{0}, \Delta \boldsymbol{x}_j] - \boldsymbol{\mu}_*^+)$ and $\boldsymbol{m}_t^- := \sum_{j=1}^t (\max[\boldsymbol{0}, \Delta \boldsymbol{x}_j] - \boldsymbol{\mu}_*^-)$. so that $\delta_* := \boldsymbol{\beta}_*^{+\prime}\boldsymbol{\mu}_*^+ + \boldsymbol{\beta}_*^{-\prime}\boldsymbol{\mu}_*^-$.

To estimate the long-run parameters, Cho et al. (2019) suggest using OLS to obtain the residuals $\widehat{\boldsymbol{m}}_t^+$ and $\widehat{\boldsymbol{m}}_t^-$ from (8) that are obtained by regressing $\boldsymbol{x}_t^+$ and $\boldsymbol{x}_t^-$ against t, respectively. They then estimate the long-run parameters by regressing y_t against $\widehat{\boldsymbol{m}}_t^+$, $\widehat{\boldsymbol{m}}_t^-$, t, and a constant term using the FM-OLS method. This approach ensures that the estimates of the long-run parameters do not suffer from singularity issues. The long-run parameter estimates obtained in this manner are super-consistent.

The second step of 2SNARDL estimates the short-run parameters in (6). Specifically, denoting $\widehat{u}_{t-1} := y_{t-1} - \widehat{\boldsymbol{\beta}}_T^{+\prime}\boldsymbol{x}_{t-1}^+ - \widehat{\boldsymbol{\beta}}_T^{-\prime}\boldsymbol{x}_{t-1}^-$, where $\widehat{\boldsymbol{\beta}}_T^+$ and $\widehat{\boldsymbol{\beta}}_T^-$ are the long-run parameter estimates obtained from the first step, the unknown short-run parameters are estimated by OLS.

The 2SNARDL estimation leverages the distinct convergence rates of the two-step estimators, resulting in consistency, and long- and short-run parameters asymptotically follow mixed-normal and normal distributions, respectively. Specifically, because the long-run coefficients are estimated using FM-OLS, which has faster convergence rates than OLS for short-run parameters, we can treat the long-run parameters as known when estimating the short-run parameters in the second step.

For post-estimation analysis, Cho et al. (2019) propose the use of Wald testing to test the symmetry hypothesis of long- and short-run parameters. Because the 2SNARDL method is based on a two-step ECM form, cointegration testing can be performed by examining the stationarity of the fitted residuals obtained from the cointegrating equation. Cho et al. (2019) employ Phillips and Perron's (1988) test for this purpose, and rejection of the null hypothesis suggests the presence of cointegration.

We primarily employ the 2SNARDL method to analyze Thai economic data, which involves a three-step process. Initially, we validate the applicability of the 2SNARDL estimation to the collected data. We provide a descriptive data summary and apply Dickey and Fuller's (1981) augmented test to assess the unit-root hypothesis. This preliminary step ensures the suitability of the data for the subsequent application of the 2SNARDL method.

Second, we specify multiple bivariate NARDL models to investigate interest rate pass-through in the Thai economy. In these models, commercial bank interest rates serve as dependent variables, and the policy rate is the explanatory variable. Furthermore, we differentiate the extent of the pass-through based on bank size. To analyze the macroeconomic consequences of interest rate changes, we employ multivariate model specifications, focusing on their impacts on output and inflation. This approach enables us to examine the dynamics of the interest rate pass-through and its broader effects on the economy.

In addition to the 2SNARDL estimation mentioned earlier, we estimate the long- and short-run parameters using the single-step NARDL estimation method. Importantly, however, the single-step estimation method suffers from a singularity problem; as a result, its limit distribution remains unavailable. Nevertheless, simulations confirm the consistency of the single-step NARDL estimation (see Cho et al., 2019). Therefore, we present bootstrap standard errors as an alternative to asymptotic standard errors. We enhance the robustness of our model inference by comparing the results obtained from the single-step NARDL estimation with those obtained from the 2SNARDL estimation. During the model specification, we determine the optimal lag length using the Schwarz Information Criterion (SIC), as advocated by Cho et al. (2023a).

For each specified model, we outline the 2SNARDL estimation procedure as follows.

(a) First, for the long-run parameter estimation of interest rate pass-through, we treat commercial bank rates as dependent variables. The policy rate is decomposed into a detrended positive partial sum process and the policy rate.

(b) Second, for the long-run parameter estimation of the macroeconomic impact model, we consider output growth and inflation as dependent variables, similarly decomposing the policy rate into positive and negative components. Additionally, we include the following control variables: log trading partner GDP for output growth and global energy price index for the inflation rate equations.

(c) Third, we perform the unit-root test, as proposed by Phillips and Perron (1988), on the fitted residual cointegrating equation. Rejecting the null hypothesis indicates stationarity, implying a long-term relationship.

(d) Fourth, for the short-run parameter estimation, we treat the fitted residual from the long-run equation estimation as the ECM term. We estimate the short-run parameters in (6) by OLS. We also apply the Wald test to assess long- and short-run asymmetries.

(e) Fifth, we report the estimation results obtained by the single-step NARDL estimation along with bootstrap standard errors computed from 5,000 replications. For the cointegration test, we employ a t_{BDM} test, as proposed by Banerjee, Dolado, and Mestre (1998).

(f) Finally, we compute DMs and present them along with a 90% confidence interval obtained using the bootstrap method with 5,000 replications.

4 Data

In this section, we outline the data used in the empirical study before discussing the empirical estimation and inference results.

We obtain the monthly data required for this study from the Thai authorities from 2001 to 2022. This timeframe aligns with the Bank of Thailand's (BOT's) adoption of the IT framework where the policy interest rate serves as its primary tool[3]. Table 1 summarizes the data statistics.

We partition the economic variables into three groups and detail Thailand's policy rates (RP) in Table 1. Policy rates have averaged approximately 2.01% annually over the past two decades. The lowest recorded rate is approximately 0.5%, whereas the highest is approximately 5%. The median rate is below the mean and registers at 1.5%. When considering the incremental rate changes, the rates of increase and decrease are nearly equal with an absolute value of 0.03%. The most significant rate hike observed is 0.76%, and the largest rate cut amounts to -0.9%.

Figure 1 illustrates historical trends in policy rates. There are three distinct cycles of interest rate increases and four significant periods of interest rate cuts. They can be classified into the following five categories.

(a) First, until 2004, policy rates remained relatively low, fluctuating between approximately 1.25% and 2.5%. However, from 2004 to 2006, the MPC significantly raised policy rates by implementing three rate hikes in the second half of 2004, six in 2005, and four in 2006. This led to an increase in the policy rate from approximately 2% in 2004 to 4% by the end of 2005, peaking at 5% in June of 2006. These measures aimed to address the upward price pressures resulting from improved domestic

[3]The BOT operated under a fixed exchange rate system until 1997 and subsequently employed monetary targeting until April 2000

economic growth, persistently high oil prices, and the global tightening of monetary policies initiated by the U.S. and other major developed economies.

(b) Second, in 2007, Thailand's economic growth was expected to be influenced by a slowdown in exports, primarily due to the adverse impacts of the subprime problem on the U.S. economy. Simultaneously, the inflationary pressure eased. Consequently, the MPC reduced its policy interest rate five times in the first half of 2007 from 5% to 3.25%. As of 2008, the global economy was anticipated to experience slow growth because of the worsening subprime crisis, which had a negative effect on the Thai economy. The MPC responded by reducing the policy rate by 1% at the end of 2008 and 1.5% over three consecutive meetings in 2009, ultimately reducing it to 1.25% by 2009.

(c) Third, between 2010 and 2011, the Thai economy entered a phase of expansion. In response, the MPC raised the policy rate three times in 2010, resulting in a cumulative increase of 0.75%. In 2011, the MPC implemented six additional consecutive rate hikes, resulting in a 1.25% increase. By October 2011, the policy rate had reached 3.5

(d) Fourth, Between 2015 and 2019, the policy rate remained stable at 1.5% with the aim of fostering economic growth. However, economic progress faced challenges owing to structural issues, including declining competitiveness in certain export sectors. This period was marked by a slow growth rate and a persistently low inflation rate, occasionally dipping into negative territory. Moreover, during this timeframe, the global financial markets experienced excess liquidity and sustained low interest rates. In response, the MPC gradually reduced its policy rate from 2019 to 2020, ultimately reaching its lowest point of 0.5%. This reduction was intended to support the sluggish economic growth and stimulate inflation. Additionally, the MPC undertook these measures to prepare and bolster the overall economy, anticipating the adverse effects of the COVID-19 pandemic that emerged at the end of 2019.

(e) Finally, after two years of economic lockdown, the Thai economy gained momentum in its recovery in 2022, and this recovery became more widespread. In response to this positive trend, the MPC made a series of decisions regarding the policy rate. They raised it three times, gradually increasing it from 0.5% to 1.25% by the end of 2022. By continuing its efforts in 2023, the MPC implemented additional rate hikes, bringing the policy rate to 2% during the first half of the year. These decisions were made to address the increasing inflation rate, which was influenced by high global oil prices, as well as to support the post-COVID-19 economic recovery. These actions were also aligned with the global trend of interest rate normalization.

Next, we describe the interest rates offered by commercial banks in Thailand. Thailand's financial system has a substantial dependence on commercial banks, which comprise approximately 47%, nearly half, of the total assets of financial institutions as of Quarter 3 of 2021. They account for approximately 70% of corporate and 44% of consumer loans. Consequently, commercial banks play a pivotal role in transmitting interest rates to the real economy. The key features are summarized as follows:

(a) Table 1 demonstrates that, on average, commercial banks establish deposit rates ranging from 0.86% to 1.52%, depending on the fixed-term period. By contrast, loan rates vary from 6.57% to 7.26% based on the type of loan. Generally, deposit rates are lower than policy rates and tend to increase over longer fixed-term periods. In terms of loan rates, the MLR, offered to well-established corporate customers, is typically set lower than the minimum overdraft rate (MOR) and MRR. Notably, when considering bank size, it becomes apparent that small and medium-sized banks tend to offer higher rates than their larger counterparts.

(b) Figure 2 compares the time series of policy rates and interest rates. It becomes evident that Long-term fixed deposit rates have moved closer to policy rates. This alignment is supported by the strong correlation between the policy rate and the 3- and 12-month deposit rates, which stands at 0.8. By contrast, the savings rate exhibits a lower correlation of only 0.5. Among the loan rates, the MLR shows the highest positive correlation with the policy rate, approximately 0.7, while the MOR and MRR have correlations of approximately 0.6. This pattern remains consistent when we consider the size of commercial banks, as shown in Figure 3. However, it is noteworthy that small and medium-sized commercial banks tend to adjust their deposit rates more closely in line with the policy rate, especially the short-term deposit rate.

(c) Finally, it is worth noting the interest rate hikes that occurred in 2005. During this period, although the MPC significantly increased the policy rate, interest rates at commercial banks for both deposits and lending remained unchanged for approximately six months. This phenomenon was attributed to bank balance sheet positions, which had substantial excess liquidity, particularly among the four largest banks (see Bank of Thailand, Monetary Policy Department, 2008; Charoenseang and Manakit, 2007). During this phase, banks shifted from being net borrowers, as seen in the pre-crisis period of 1997, when they held liquid assets close to the reserve requirement, to become net lenders. Generally, when banks are net borrowers, an increase in policy rates swiftly elevates their funding costs, prompting them to seek alternative funding sources and ultimately amplifying the pass-through effect. Excess liquidity was resolved in late Quarter 2 of 2005, leading commercial banks to adjust their deposit and lending rates accordingly.

Finally, we describe the macroeconomic variables used in this study in detail. We incorporate four primary economic indicators: the coincident economic indicator (CEI), the headline consumer price index

(HCPI), log trading partner GDP (lnTPGDP), and log global energy price (lnENERGYP). The following is a summary of these variables:

(a) The CEI serves as a proxy for output given its availability in monthly data. Notably, there is a strong correlation of 0.9 between the CEI and quarterly GDP, highlighting its reliability as an indicator. As indicated in Table 1, the CEI demonstrates an average growth rate of 3.32% over the sample period, ranging from a minimum of -13.24% to a maximum of 21.24%. Figure 4 visually depicts the challenges faced by the Thai economy, including the impact of the global financial crisis in 2009, major flooding in 2011, internal political turmoil in 2014, and the global COVID-19 pandemic in 2020-2021.

(b) Inflation is quantified by the annual growth of the HCPI, which serves as the current indicator for IT. The target range is set at 1%-3%. Over the same sample period, inflation rate averaged 2.07%, with lows of -4.22% and highs of 9.18%. Figure 4 also illustrates the positive correlation between the inflation rate and output, aligned with conventional macroeconomic theory. The Thai inflation rate has remained low since 2015 because of structural issues and low oil prices. However, it experienced a significant increase owing to a sharp drop in growth during the 2020 pandemic. Following the pandemic, the inflation rate enjoyed a remarkable rebound owing to robust economic growth.

(c) The other two variables serve as control variables in the NARDL model. Given the Thai economy's heavy reliance on exports, we exercise control over the output equation by incorporating the trading partners' GDP index. This index is constructed from leading indicators of major Asian economies, including China, Japan, Korea, India, and Indonesia, as well as the U.S., the U.K., and Australia, collectively accounting for approximately 50% of Thailand's export value. Additionally, this study accounts for the influence of supply shocks, represented by global energy prices, in assessing the impact of monetary policy on the inflation rate.

5 Empirical Results

In this section, we present the findings derived from NARDL model estimations. We primarily focus on the results obtained from the 2SNARDL model. First, we present the estimated outcomes of the pass-through of monetary policy interest rates to commercial banks, encompassing both deposit and lending rates. We then compare these findings across different bank sizes. Finally, we report the estimated macroeconomic impacts of monetary policy shocks on output and inflation. The results of the unit-root tests for the variables are provided in the Appendix.

Our presentation plans remain consistent across the various models. We initiate a cointegration test by employing Phillips and Perron's (1988) test for the fitted error-correction term. The rejection of the null hypothesis indicates the presence of a long-term relationship. In the single-step method, we apply a residual bootstrap to compute standard errors. Additionally, we apply the Wald test to assess the long- and short-run symmetries. Finally, we provide the dynamic impulse response functions of the variables of interest using the estimated parameters.

5.1 Results of Interest Rate Pass-through

In this section, we discuss the NARDL estimation and inference results for interest rate pass-through on deposit and loan rates. Tables 2 and 3 present the findings concerning interest rate pass-through on deposit and loan rates, respectively. We estimate the NARDL(2,2) and NARDL(3,2) models depending on the type of bank interest rates. The estimation results are summarized as follows.

(a) The results offer empirical evidence of a long-run relationship between retail rates and policy rates, as demonstrated by Phillips and Perron's (1988) test of the cointegrating equation residuals and t_{BDM} exceeding the critical value at a significance level of 0.05. In summary, the findings suggest limited evidence for the presence of a long-run asymmetric pass-through for deposit rates while strongly supporting the existence of long-run asymmetry for loan rates.

(b) In terms of the degree of pass-through on deposit rates, as indicated in Table 2, the analysis reveals a range of 90.3% to 102.7% for the pass-through from policy rates to term-fixed deposit rates in the long run. Specifically, the pass-through degree for three-month term-fixed deposit rates is approximately 90.3% for rate reductions, slightly lower than that for rate increases at approximately 91%. However, this difference was not statistically significant. In contrast, the pass-through to 12-month term-fixed deposit rates is more pronounced with rates of approximately 91.4% for rate cuts and 102.7% for rate hikes. Again, this difference is not statistically significant. Of particular interest are the findings concerning the savings rate, which demonstrate long-run downward asymmetry. In other words, the pass-through for rate increases is only 28.1%, compared with approximately 40.4% for rate cuts.

(c) For loan rates, the results indicate incomplete pass-through, ranging from 42.8% to 89.6%, with long-run upward asymmetric pass-through. In the case of MLR, the estimated long-run pass-through coefficients are 0.493 for rate cuts and 0.784 for rate hikes, and the test statistics confirms the difference between these two coefficients. This suggests that, in the long run, the policy rate is transmitted to the MLR at approximately 49.3% for rate cuts and 78.4% for rate hikes, indicating upward asymmetry or downward rigidity. A similar pattern holds for the MOR where the coefficient for the rate increase is approximately 0.775, whereas, for the rate decrease, it is approximately 0.428. Furthermore, the

Wald test for long-run asymmetry demonstrates statistical significance, signifying upward rigidity in loan rate adjustments. As for the MRR, rate hikes have coefficients of approximately 0.896, while rate cuts show an insignificant coefficient. These results support the asymmetric pass-through of loan rates in Thailand.

(d) The results reveal a notably higher degree of pass-through for both deposit and loan rates compared to findings in the existing literature on Thai data, such as the studies by Disyatat and Vongsinsirikul (2003) and Charoenseang and Manakit (2007), where pass-through rates range only between 21.9% and 70%. This disparity can be attributed to the use of more recent data encompassing the period after 2009, when the Thai banking system underwent further development. Furthermore, the literature was significantly influenced by the study period, which was characterized by the presence of exceptionally high excess liquidity in the banking system. According to Bank of Thailand, Monetary Policy Department (2008) and Gigineishvili (2011), elevated levels of excess liquidity tend to result in fewer interest-rate adjustments. The sample period of our data spans 2001–2023, encompassing the excess liquidity period 2007-2008 when banks exhibited a greater inclination to adjust their interest rates. Notably, Thai commercial banks' liquid asset ratio stood at 15.4% in March 2023, in contrast to approximately 25% in 2007-2008.

(e) Figure 5 provides the impulse-response functions of savings and loan rates to policy rate hikes. These results are shown in the left and right panels, respectively. The left panels offer additional insights into the short-run asymmetry, indicating a downward asymmetric pass-through for savings rates. In response to policy-rate hikes, savings rates exhibit only a 20% adjustment within 12 months before reaching a long-run rate adjustment. The results also reveal downward short-run asymmetry in term fixed deposit rates with a significant immediate adjustment following policy rate cuts, whereas the adjustment is less pronounced for rate hikes. It takes approximately 20 months for asymmetry to fade. For the loan rates displaced in the right panels, there is no evidence of short-run asymmetry, but there is upward long-run asymmetry, as indicated by the 90% confidence interval situated above zero.

5.2 Interest Rate Pass-through Across Bank Sizes

In this section, we conduct a comprehensive comparison of the long-run pass-through of interest rates while controlling for bank size. We present the estimation results in Table 4 and detailed regression results in Tables 7 and 8. These results are based on NARDL(2,2), NARDL(2,3), and NARDL(3,2), relying on the type of bank interest rates. We summarize the estimation and inference results as follows.

(a) Overall, the results indicate that loan rates exhibit greater pass-through for large banks than for small and medium-sized banks, whereas the findings are mixed for deposit rates. In terms of magnitude, as shown in Table 4, the pass-through level of savings rates for small and medium banks is approximately 35.9% for rate hikes and 46.4% for rate reductions. Conversely, the corresponding pass-through levels for large banks are only 18.2% and 0.8% (22.1% and 19.9%, respectively, using the single-step method). When it comes to three-month term-fixed deposit rates, the pass-through levels for small and medium-sized banks are 93.1% for rate increases and 89.8% for rate decreases. By contrast, for large banks, the pass-through levels are 83.6% for rate increases and 92.2% for rate cuts, indicating downward asymmetry. Regarding 12-month term-fixed deposit rates, large banks exhibit a pass-through degree exceeding 100% for both rate hikes and cuts, whereas, for small and medium-sized banks, it is evident that upward asymmetry exists.

(b) When it comes to lending rates, large banks exhibit an approximate adjustment of 78.9% to 81.6% in response to policy rate hikes, while small and medium-sized banks show an adjustment ranging from around 72.6% to 86.4%. In the case of rate cuts, large banks respond with adjustments ranging from 26.7% to 56.6%, whereas small and medium-sized banks exhibit a response of only 10.5% to 48.4%. These findings suggest that large banks have a greater capacity to adjust their loan rates in response to policy rate changes than small and medium-sized banks.

(c) From an intuitive perspective, these findings indicate that small and medium-sized banks, with lower market power than larger banks, may need to make more significant adjustments in their deposit rates to compete for market expansion and retain their deposit base, especially for longer-term fixed deposits. However, they may exercise caution and make less prominent adjustments to lending rates to enhance customer appeal, resulting in stickiness in lending rates. By contrast, larger banks, benefiting from a stronger market position, have more flexibility in reacting to policy rate changes. These results are consistent with the observations of Weth (2002) regarding stickiness in lending rate adjustments for small banks in Germany. Furthermore, the findings support the notion that bank market power influences the degree of interest rate pass-through, as mentioned by Andries and Billon (2016) and Gregor et al. (2021).

(d) Along with these findings, we acknowledge the possibility of a significant issue related to omitted variables in the cointegrating equation specification for loan rates. This is indicated by considerably low adjusted R^2 values, approximately 0.09 to 0.211 for the overall bank sample, 0.121 to 0.27 for the large bank sample, and 0.075 to 0.204 for the small and medium bank sample.

5.3 Results of Macroeconomic Impact

In this section, we discuss the NARDL estimation and inference results for macroeconomic impact. We present the estimation and inference results in Table 9 for the relationships between output and interest rates and between inflation and interest rates. These results are obtained by estimating NARDL (12,2), and it is evident from the results that interest rate increases have a more pronounced impact on output growth than rate decreases. Conversely, when it comes to inflation, interest rate decreases have a more substantial impact than rate increases. The key results are summarized as follows.

(a) For the impact on output, after controlling for inflation and lnTPGDP, the results suggest the presence of long-run cointegration among variables, as indicated by Phillips and Perron's (1988) test. Table 9 provides insights into these findings, revealing that policy rate increases have a more significant impact on output growth than rate reductions. Specifically, an increase in policy rates reduces output growth by 2.20% (1.018%, according to the single-step method), whereas a rate reduction stimulates output growth by only 1.627% (0.751%, based on the single-step method). To further support these results, the statistically significant Wald tests, $\chi^2_{LR\_ASYM}$ and $\chi^2_{SR\_ASYM}$, confirm the presence of asymmetric effects in both the long and short runs at a significance level of 1%.

(b) When examining the impact on inflation, as presented in Table 9, it becomes evident that policy rate decreases exert a more substantial impact on inflation compared to rate hikes. Specifically, a 1% increase in the policy rate translates into a decrease in the inflation rate of approximately 0.275% (0.133%, according to the single-step method). By contrast, a 1% decrease in the policy rate leads to an increase in inflation of approximately 0.924% (0.814%, based on the single-step method). However, the estimated coefficients of rate increases are statistically insignificant. The test statistics from Phillips and Perron (1988) for the fitted error correction term also confirm the existence of long-term cointegration.

(c) In terms of dynamic effects, Figure 8 illustrates the long-run asymmetric effects of policy rate changes on output and inflation. Notably, when the policy rate increases, its impact on output is more pronounced compared to rate decreases, whereas inflation demonstrates the opposite pattern. This asymmetry is evident in the bootstrap confidence interval, which consistently falls below zero for all horizons for output and above zero for inflation.

(d) For the impact on output, Figure 8 illustrates that rate cuts do not immediately affect output growth but instead take some time to influence it before stabilizing at the long-term impact level. In contrast, policy rate increases have immediate and significant effects on output. Initially, output growth declines but later rebounds and reaches the long-run impact level within 24 months. These findings align with insights from the new Keynesian framework, as discussed in Section 2.

(e) For inflation, as shown in Figure 9, there is a notable price puzzle[4] observed in the short run, consistent with findings in existing literature (e.g., see Sims, 1992; Eichenbaum, 1992; Christiano, Eichenbaum, and Evans, 1996). However, for longer time horizons, the results indicate that policy rate increases are associated with a decrease in inflation, while rate cuts lead to an increase in inflation. These findings align with the new Keynesian framework, which suggests that contractionary monetary policy results in decreased inflation and vice versa.

(f) Our results align with the existing empirical literature, which observes a more pronounced monetary policy effect on output for rate hikes than for rate cuts, while the results reverse for inflation (e.g., see Debortoli et al., 2020; Adelakun and Yousfi, 2020). Specifically, decreases in inflation resulting from policy rate hikes have a lesser effect than the increased inflation resulting from rate decreases, indicating the downward price rigidity of interest rate shocks. This phenomenon finds support in the presence of price adjustment costs, for example, the menu costs models of Ball and Mankiw (1994).

Based on these empirical results, we contribute to the literature by emphasizing the asymmetric effects of policy rates on inflation in the long term, specifically in Thailand. The results highlight that an increase in inflation resulting from a reduction in the policy rate is approximately 3.4 times as significant as the deflationary impact caused by an increase in the policy rate. In terms of output, the upward asymmetric impact of a policy rate increase is approximately 1.4 times greater than the impact magnitude of a policy rate decrease.

6 Conclusion

This study empirically analyzes data from Thailand using the NARDL model to investigate the asymmetric interest rate pass-through, pass-through variations across bank sizes, and the macroeconomic impact of monetary policy shocks on output and inflation.

In terms of the interest rate pass-through, we identify a long-run relationship or cointegration between the policy and retail rates. The results strongly indicate that loan rates exhibit an incomplete pass-through, showing an upward asymmetric pass-through in the long run with pass-through magnitudes ranging from 42.8% to 89.6%. However, evidence of an asymmetric pass-through for savings and deposit rates is limited

[4]The price puzzle refers to an unexpected association between a positive interest rate shock and an increase in inflation, contrary to the conventional expectation of rate hikes reducing demand and lowering prices. It was primarily introduced in Eichenbaum's (1992) commentary on Sims' (1992). Additionally, the link between the price puzzle and shock identification is summarized in Ramey (2016).

with magnitudes ranging from 28.1% to 102.7%. Notably, these findings demonstrate higher pass-through rates than those reported in the existing literature. This is likely attributable to the use of more recent data, including the post-2009 period, which is characterized by the absence of high excess liquidity.

Regarding the variations in pass-through across bank sizes, our findings indicate that large banks exhibited a more pronounced response to policy rate changes in loan rates than small and medium-sized banks. This finding suggests a higher degree of rigidity in adjusting loan rates among smaller banks. Conversely, larger banks demonstrate a greater ability to adjust their loan rates, highlighting the influence of market power on banks' pass-through behavior concerning lending rates. In the case of term-fixed deposit rates, smaller banks exhibit upward asymmetry or rigidity in reducing deposit rates, emphasizing the importance of retaining their deposit bases.

We also examine the macroeconomic implications of policy-rate changes on output and inflation. These results align with the principles of the new Keynesian framework and provide compelling evidence of asymmetric effects. Specifically, increases in interest rates had a more substantial impact on output growth compared to rate reductions. In contrast, interest rates hikes show less impact on inflation than interest rate deceases. Quantitatively, rate hikes had an effect 1.4 times larger than the impact of rate cuts on output growth, whereas rate cuts had an effect 3.4 times greater on inflation. These findings suggest that, in the context of Thailand, there exists downward price adjustment rigidity associated with monetary policy shocks.

Overall, this study contributes to the understanding of the transmission mechanism and macroeconomic impact of monetary policy in Thailand. The strong evidence of the asymmetric interest rate pass-through and the macroeconomic effects of monetary policy shocks offer valuable insights for policymakers in conducting monetary policy strategies to ensure macroeconomic and price stability.

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Table 1: Summary Statistics of Data

	Mean	SD	Min.	Max.	Median	Size
Policy Rate						
Policy Rate (RP)	2.020	1.066	0.500	5.000	1.310	270
RP^{+}	5.648	2.617	0.000	9.290	4.790	270
RP^{-}	-5.128	2.998	-8.790	0.000	-7.540	270
ΔRP^{+}	0.034	0.087	0.000	0.760	0.000	270
ΔRP^{-}	-0.033	0.098	-0.900	0.000	0.000	270
Commercial Banks						
Saving Rate (CB_SAV)	0.910	0.471	0.318	2.442	0.571	270
Time Deposits Rate (3 Months, CB_3M)	1.563	0.863	0.462	4.073	1.039	270
Time Deposits Rate (12 Months, CB_12M)	1.897	0.939	0.606	4.506	1.167	270
MLR (CB_MLR)	6.863	0.573	5.913	8.154	6.332	270
MOR (CB_MOR)	7.306	0.592	6.208	8.615	6.683	270
MRR (CB_MRR)	7.688	0.730	6.392	8.712	7.043	270
Large-sized Commercial Banks						
Saving Rate ($LARGE\_SAV$)	0.725	0.432	0.250	2.450	0.500	270
Time Deposits Rate (3 Months, $LARGE\_3M$)	1.409	0.843	0.331	3.938	0.925	270
Time Deposits Rate (12 Months, $LARGE\_12M$)	1.766	0.970	0.420	4.500	1.025	270
MLR ($LARGE\_MLR$)	6.458	0.655	5.360	7.900	5.863	270
MOR ($LARGE\_MOR$)	6.917	0.658	5.866	8.350	6.137	270
MRR ($LARGE\_MRR$)	7.191	0.741	5.997	8.350	6.412	270
Small and Medium-sized Commercial Banks						
Saving Rate (SM_SAV)	0.987	0.495	0.356	2.438	0.626	270
Time Deposits Rate (3 Months, SM_3M)	1.631	0.867	0.539	4.124	1.085	270
Time Deposits Rate (12 Months, SM_12M)	1.954	0.922	0.722	4.522	1.238	270
MLR (SM_MLR)	7.048	0.543	6.025	8.312	6.552	270
MOR (SM_MOR)	7.483	0.562	6.375	8.781	6.887	270
MRR (SM_MRR)	7.923	0.720	6.556	8.938	7.273	270
Economic Variables						
Real Output Growth (%YoY, CEI)	3.324	3.865	-13.23	21.24	2.113	270
Headline Inflation Rate (%YoY, $HCPI$)	2.091	2.172	-4.143	9.266	0.656	270
Log of Trading Partner GDP ($lnTPGDP$)	4.603	0.011	4.533	4.624	4.599	270
Log of Global Energy Price Index ($lnENERGYP$)	4.375	0.449	3.298	5.138	4.085	270

Source: The BOT, Thailand's Ministry of Commerce, OECD statistics, and FRED economic data, summarized by authors.

Table 2: Interest Rate Pass-through Results for Banks' Deposit Rate

Long-run Parameters

	Saving Rate				Term Deposits Rate 3 Months				Term Deposits Rate 12 Months			
	Two-step		Single-step		Two-step		Single-step		Two-step		Single-step	
	Estimate	S.E.	Estimate	S.E.	Estimate	S.E.	Estimate	S.E.	Estimate	S.E.	Estimate	S.E.
RP^+	0.281***	0.030	0.354***	0.080	0.910***	0.043	0.859***	0.092	1.027***	0.062	0.893***	0.106
RP^-	0.404***	0.041	0.400***	0.071	0.903***	0.060	0.866***	0.083	0.914***	0.087	0.878***	0.102
Constant	0.372***	0.132	-		-0.254	0.192	-		0.073	0.277	-	
R^2	0.674				0.760				0.666			
$Adj.R^2$	0.670				0.757				0.662			
LR Asym.												
$\chi^2_{LR\_ASYM}$	11.93		4.679		0.016		0.082		2.311		0.271	
P-value	0.001		0.031		0.898		0.775		0.128		0.603	
Coint.												
T-test	-3.949		-3.019		-5.180		-2.810		-5.637		-3.000	
P-value	0.002				0.000				0.000			

Note: (1) The single-step long-run (LR) parameters are determined as $\hat{\beta}^+_T := -\hat{\theta}^+_T/\hat{\rho}_T$ and $\hat{\beta}^-_T := -\hat{\theta}^-_T/\hat{\rho}_T$. For the two-step NARDL, the FM-OLS estimator are employed in the first-step, and the parameters are computed as $\ddot{\beta}^+_T := \ddot{\lambda}_T + \ddot{\eta}_T$ and $\ddot{\beta}^-_T := \ddot{\eta}_T$ using nonlinear combinations based on the delta method.
(2) In the single-step method, the intercept of the cointegrating equation is left unidentified, while bootstrap standard errors (S.E.) derived from 5,000 replications are reported. Asymptotic S.E. are reported for the FM-OLS parameters in the two-step model. The significant level: *** p<0.01, ** p<0.05, * p<0.1
(3) $\chi^2_{LR\_ASYM}$ denotes the Wald test statistic for LR asymmetric coefficients.
(4) Cointegration test: For the single-step method, the t_{BDM} proposed by Banerjee et al. (1998) for testing the null hypothesis of no cointegration is reported. Pesaran et al. (2001) tabulate 5% critical value of t_{BDM} as -3.22, -3.53, and -3.78 for $k = 1, 2,$ and 3, respectively. In the two-step model, the paper uses the Phillips and Perron (1988)'s unit-root test on LR equation residuals. Rejecting the null hypothesis indicates stationary residuals, implying a long-run relationship.

Short-run Parameters

	Saving Rate				Term Deposits Rate 3 Months				Term Deposits Rate 12 Months			
	Two-step		Single-step		Two-step		Single-step		Two-step		Single-step	
	Estimate	S.E.	Estimate	S.E.	Estimate	S.E.	Estimate	S.E.	Estimate	S.E.	Estimate	S.E.
Constant	0.031***	0.011	0.043**	0.017	0.067***	0.021	0.064**	0.026	0.024***	0.009	0.070***	0.027
Y_{t-1}			-0.045***	0.015			-0.053***	0.019			-0.050***	0.017
RP^+_{t-1}			0.016***	0.005			0.046***	0.015			0.045***	0.016
RP^-_{t-1}			0.018***	0.005			0.046***	0.015			0.044***	0.016
ECM_{t-1}	-0.031***	0.010			-0.049***	0.016			-0.032***	0.011		
ΔY_{t-1}	0.333***	0.100	0.267***	0.102	0.432***	0.081	0.436***	0.084	0.414***	0.074	0.417***	0.079
ΔY_{t-2}					-0.057	0.067	-0.053	0.071				
ΔRP^+_t	0.049*	0.029	0.062**	0.031	0.196***	0.071	0.191**	0.081	0.308***	0.081	0.291***	0.086
ΔRP^+_{t-1}	0.032	0.025	0.037	0.028	0.106	0.067	0.102	0.075	0.069	0.089	0.047	0.093
ΔRP^-_t	0.245***	0.047	0.247***	0.056	0.565***	0.049	0.561***	0.064	0.659***	0.064	0.663***	0.081
ΔRP^-_{t-1}	-0.033	0.041	-0.020	0.046	-0.043	0.066	-0.047	0.073	-0.136*	0.071	-0.133*	0.078
R^2	0.527		0.550		0.796		0.797		0.760		0.765	
$Adj.R^2$	0.516		0.536		0.791		0.789		0.754		0.757	
SR Asym.												
$\chi^2_{SR\_ASYM}$	5.113		3.686		4.918		3.960		1.816		2.423	
P-value	0.025		0.055		0.027		0.047		0.179		0.120	
$\chi^2_{S.Corr}$	0.052		1.033		1.959		1.718		3.381		3.569	
χ^2_{Hetero}	49.22		67.99		9.017		9.325		11.41		13.56	
	NARDL(2,2)				NARDL(3,2)				NARDL(2,2)			

Note: (1) In the single-step method, bootstrap S.E, derived from 5,000 replications are reported. Asymptotic S.E. are reported for the OLS parameters in the two-step model. The significant level: *** p<0.01, ** p<0.05, * p<0.1
(2) $\chi^2_{SR\_ASYM}$ denotes the Wald test statistic for SR asymmetry. $\chi^2_{S.Corr}$ and χ^2_{Hetero} denote the Breusch–Godfrey LM test for autocorrelation and the Breusch–Pagan/Cook–Weisberg test for heteroskedasticity, respectively.

Table 3: Interest Rate Pass-through Results for Banks' Loan Rates

Long-run Parameters

	Minimum Loan Rates MLR				**Minimum Overdraft Rate MOR**				**Minimum Retail Rate MRR**			
	Two-step		**Single-step**		**Two-step**		**Single-step**		**Two-step**		**Single-step**	
	Estimate	**S.E.**	**Estimate**	**S.E.**	**Estimate**	**S.E.**	**Estimate**	**S.E.**	**Estimate**	**S.E.**	**Estimate**	**S.E.**
RP^{+}	0.784***	0.064	0.957***	0.195	0.775***	0.078	1.072***	0.223	0.896***	0.114	1.261***	0.194
RP^{-}	0.493***	0.090	0.752***	0.153	0.428***	0.109	0.814***	0.180	0.127	0.160	0.973***	0.178
Constant	5.605***	0.288	-		6.176***	0.348	-		7.473***	0.512	-	
R^2	0.220				0.117				0.100			
$Adj.R^2$	0.211				0.107				0.090			
LR Asym.												
$\chi^2_{LR\_ASYM}$	14.06		8.635		13.71		10.65		31.08		18.82	
P-value	0.000		0.003		0.000		0.001		0.000		0.000	
Coint.												
T-test	-5.890		-2.639		-5.656		-2.842		-3.749		-3.470	
P-value	0.000				0.000				0.003			

Note: (1) The single-step long-run (LR) parameters are determined as $\hat{\beta}^{+}_{T} := -\hat{\theta}^{+}_{T}/\hat{\rho}_{T}$ and $\hat{\beta}^{-}_{T} := -\hat{\theta}^{-}_{T}/\hat{\rho}_{T}$. For the two-step NARDL, the FM-OLS estimator are employed in the first-step, and the parameters are computed as $\ddot{\beta}^{+}_{T} := \ddot{\lambda}_{T} + \ddot{\eta}_{T}$ and $\ddot{\beta}^{-}_{T} := \ddot{\eta}_{T}$ using nonlinear combinations based on the delta method.
(2) In the single-step method, the intercept of the cointegrating equation is left unidentified, while bootstrap standard errors (S.E.) derived from 5,000 replications are reported. Asymptotic S.E. are reported for the FM-OLS parameters in the two-step model. The significant level: *** p<0.01, ** p<0.05, * p<0.1
(3) $\chi^2_{LR\_ASYM}$ denotes the Wald test statistic for LR asymmetric coefficients.
(4) Cointegration test: For the single-step method, the t_{BDM} proposed by Banerjee et al. (1998) for testing the null hypothesis of no cointegration is reported. Pesaran et al. (2001) tabulate 5% critical value of t_{BDM} as -3.22, -3.53, and -3.78 for $k = 1, 2,$ and 3, respectively. In the two-step model, the paper uses the Phillips and Perron (1988)'s unit-root test on LR equation residuals. Rejecting the null hypothesis indicates stationary residuals, implying a long-run relationship.

Short-run Parameters

	Minimum Loan Rates MLR				**Minimum Overdraft Rate MOR**				**Minimum Retail Rate MRR**			
	Two-step		**Single-step**		**Two-step**		**Single-step**		**Two-step**		**Single-step**	
	Estimate	**S.E.**	**Estimate**	**S.E.**	**Estimate**	**S.E.**	**Estimate**	**S.E.**	**Estimate**	**S.E.**	**Estimate**	**S.E.**
Constant	-0.015***	0.006	0.167**	0.075	-0.021***	0.007	0.152**	0.065	-0.029**	0.011	0.157***	0.053
Y_{t-1}			-0.031***	0.012			-0.028***	0.010			-0.028***	0.008
RP^{+}_{t-1}			0.030***	0.009			0.030***	0.008			0.035***	0.009
RP^{-}_{t-1}			0.024***	0.008			0.023***	0.007			0.027***	0.008
ECM_{t-1}	-0.021***	0.006			-0.018***	0.006			-0.006**	0.002		
ΔY_{t-1}	0.368***	0.083	0.340***	0.086	0.350***	0.076	0.311***	0.078	0.324***	0.087	0.234**	0.097
ΔY_{t-2}	-0.046	0.058	-0.071	0.061	-0.027	0.057	-0.062	0.063				
ΔRP^{+}_{t}	0.191***	0.067	0.190***	0.073	0.169***	0.060	0.170***	0.064	0.163**	0.070	0.152**	0.063
ΔRP^{+}_{t-1}	0.130*	0.073	0.107	0.084	0.141*	0.073	0.115	0.081	0.180**	0.088	0.124	0.078
ΔRP^{-}_{t}	0.345***	0.075	0.367***	0.090	0.328***	0.076	0.357***	0.092	0.334***	0.088	0.380***	0.106
ΔRP^{-}_{t-1}	-0.037	0.058	-0.019	0.064	-0.031	0.058	-0.007	0.065	-0.039	0.064	-0.001	0.072
R^2	0.615		0.627		0.571		0.588		0.461		0.502	
$Adj.R^2$	0.605		0.614		0.560		0.574		0.449		0.486	
SR Asym.												
$\chi^2_{SR\_ASYM}$	0.027		0.232		0.028		0.347		0.232		0.731	
P-value	0.870		0.630		0.868		0.556		0.631		0.392	
$\chi^2_{S.Corr}$	1.791		5.260		2.398		6.025		1.464		0.006	
χ^2_{Hetero}	7.695		9.658		17.88		22.47		3.244		16.22	
	NARDL(3,2)				NARDL(3,2)				NARDL(2,2)			

Note: (1) In the single-step method, bootstrap S.E, derived from 5,000 replications are reported. Asymptotic S.E. are reported for the OLS parameters in the two-step model. The significant level: *** p<0.01, ** p<0.05, * p<0.1
(2) $\chi^2_{SR\_ASYM}$ denotes the Wald test statistic for SR asymmetry. $\chi^2_{S.Corr}$ and χ^2_{Hetero} denote the Breusch–Godfrey LM test for autocorrelation and the Breusch–Pagan/Cook–Weisberg test for heteroskedasticity, respectively.

Table 4: Summary of Interest Rate Pass-through Results

Long-run Pass-through (%)	Overall			Large Banks			SM Banks		
Deposit Rates:	SAV	3M	12M	SAV	3M	12M	SAV	3M	12M
Policy Rate Hikes	28.1	91.0	102.7	0.8#	83.6	104.8	35.9	93.1	100.7
Policy Rate Cuts	40.4	90.3	91.4	18.2	92.2	102.6	46.4	89.8	87.0
Long-run Asymmetry	yes	no	no	yes	yes	no	yes	no	yes
Short-run Asymmetry	yes	yes	no	no	yes	no	yes	yes	yes
Loan Rates:	MLR	MOR	MRR	MLR	MOR	MRR	MLR	MOR	MRR
Policy Rate Hikes	78.4	77.5	89.6	81.6	78.9	88.9	73.6	72.6	86.4
Policy Rate Cuts	49.3	42.8	12.7#	56.6	48.4	26.7	48.4	42.9	10.5#
Long-run Asymmetry	yes	yes	yes	yes	yes	yes	yes	yes	yes
Short-run Asymmetry	no	no	no	no	no	no	no	no	no

Note: # indicates the coefficient is not statistically significant.

Table 5: Interest Rate Pass-through Results: Large Banks' Deposit Rates

Long-run Parameters

	Saving Rate				**Term Deposits Rate 3 Months**				**Term Deposits Rate 12 Months**			
	Two-step		**Single-step**		**Two-step**		**Single-step**		**Two-step**		**Single-step**	
	Estimate	**S.E.**	**Estimate**	**S.E.**	**Estimate**	**S.E.**	**Estimate**	**S.E.**	**Estimate**	**S.E.**	**Estimate**	**S.E.**
RP^{+}	0.008	0.032	0.199*	0.108	0.836***	0.041	0.787***	0.090	1.048***	0.069	0.964***	0.129
RP^{-}	0.182***	0.045	0.221**	0.088	0.922***	0.057	0.814***	0.078	1.026***	0.097	0.932***	0.119
Constant	0.788***	0.145	-		-0.423**	0.183	-		-0.316	0.310	-	
R^2	0.481				0.763				0.589			
$Adj.R^2$	0.476				0.760				0.584			
LR Asym.												
$\chi^2_{LR\_ASYM}$	19.75		0.309		3.030		1.082		0.075		0.609	
P-value	0.000		0.578		0.082		0.298		0.785		0.435	
Coint.												
T-test	-2.738		-2.191		-3.784		-2.972		-4.922		-2.792	
P-value	0.068				0.003				0.000			

Note: (1) The single-step long-run (LR) parameters are determined as $\hat{\beta}^{+}_{T} := -\hat{\theta}^{+}_{T}/\hat{\rho}_T$ and $\hat{\beta}^{-}_{T} := -\hat{\theta}^{-}_{T}/\hat{\rho}_T$. For the two-step NARDL, the FM-OLS estimator are employed in the first-step, and the parameters are computed as $\ddot{\beta}^{+}_{T} := \ddot{\lambda}_T + \ddot{\eta}_T$ and $\ddot{\beta}^{-}_{T} := \ddot{\eta}_T$ using nonlinear combinations based on the delta method.
(2) In the single-step method, the intercept of the cointegrating equation is left unidentified, while bootstrap standard errors (S.E.) derived from 5,000 replications are reported. Asymptotic S.E. are reported for the FM-OLS parameters in the two-step model. The significant level: *** p<0.01, ** p<0.05, * p<0.1
(3) $\chi^2_{LR\_ASYM}$ denotes the Wald test statistic for LR asymmetric coefficients.
(4) Cointegration test: For the single-step method, the t_{BDM} proposed by Banerjee et al. (1998) for testing the null hypothesis of no cointegration is reported. Pesaran et al. (2001) tabulate 5% critical value of t_{BDM} as -3.22, -3.53, and -3.78 for $k = 1, 2,$ and 3, respectively. In the two-step model, the paper uses the Phillips and Perron (1988)'s unit-root test on LR equation residuals. Rejecting the null hypothesis indicates stationary residuals, implying a long-run relationship.

Short-run Parameters

	Saving Rate				**Term Deposits Rate 3 Months**				**Term Deposits Rate 12 Months**			
	Two-step		**Single-step**		**Two-step**		**Single-step**		**Two-step**		**Single-step**	
	Estimate	**S.E.**	**Estimate**	**S.E.**	**Estimate**	**S.E.**	**Estimate**	**S.E.**	**Estimate**	**S.E.**	**Estimate**	**S.E.**
Constant	0.005	0.006	0.017	0.013	0.087***	0.032	0.078**	0.032	0.070***	0.025	0.057*	0.030
Y_{t-1}			-0.029**	0.013			-0.066***	0.022			-0.052***	0.019
RP^{+}_{t-1}			0.006*	0.003			0.052***	0.016			0.050***	0.018
RP^{-}_{t-1}			0.006*	0.003			0.054***	0.017			0.048***	0.018
ECM_{t-1}	-0.009	0.008			-0.046***	0.016			-0.048***	0.016		
ΔY_{t-1}	0.316***	0.108	0.251**	0.118	0.332***	0.101	0.322***	0.103	0.312***	0.084	0.315***	0.088
ΔY_{t-2}					0.019	0.066	0.012	0.074				
ΔRP^{+}_{t}	0.019	0.014	0.028	0.020	0.216**	0.091	0.226**	0.089	0.421***	0.101	0.424***	0.112
ΔRP^{+}_{t-1}	0.031	0.021	0.033	0.023	0.046	0.071	0.050	0.072	-0.019	0.110	-0.009	0.123
ΔRP^{+}_{t-2}	0.026	0.025	0.027	0.028								
ΔRP^{-}_{t}	0.162*	0.092	0.165*	0.098	0.679***	0.086	0.671***	0.102	0.796***	0.089	0.785***	0.108
ΔRP^{-}_{t-1}	-0.071	0.052	-0.061	0.063	-0.179**	0.081	-0.173**	0.086	-0.211**	0.086	-0.214**	0.094
ΔRP^{-}_{t-2}	0.031	0.045	0.029	0.048								
R^2	0.295		0.333		0.690		0.700		0.670		0.671	
$Adj.R^2$	0.273		0.308		0.681		0.689		0.662		0.661	
SR Asym.												
$\chi^2_{SR\_ASYM}$	0.304		0.231		3.816		2.938		1.791		1.021	
P-value	0.582		0.631		0.052		0.087		0.182		0.312	
$\chi^2_{S.Corr}$	0.174		0.066		0.113		0.020		0.068		0.130	
χ^2_{Hetero}	300.8		341.9		3.344		0.718		5.535		6.980	
	NARDL(2,3)				NARDL(3,2)				NARDL(2,2)			

Note: (1) In the single-step method, bootstrap S.E, derived from 5,000 replications are reported. Asymptotic S.E. are reported for the OLS parameters in the two-step model. The significant level: *** p<0.01, ** p<0.05, * p<0.1
(2) $\chi^2_{SR\_ASYM}$ denotes the Wald test statistic for SR asymmetry. $\chi^2_{S.Corr}$ and χ^2_{Hetero} denote the Breusch–Godfrey LM test for autocorrelation and the Breusch–Pagan/Cook–Weisberg test for heteroskedasticity, respectively.

Table 6: Interest Rate Pass-through Results: Large Banks' Loan Rates

Long-run Parameters

	Minimum Loan Rates MLR				Minimum Overdraft Rate MOR				Minimum Retail Rate MRR			
	Two-step		Single-step		Two-step		Single-step		Two-step		Single-step	
	Estimate	**S.E.**	**Estimate**	**S.E.**	**Estimate**	**S.E.**	**Estimate**	**S.E.**	**Estimate**	**S.E.**	**Estimate**	**S.E.**
RP^+	0.816***	0.066	1.061***	0.250	0.789***	0.088	1.235***	0.284	0.889***	0.103	1.346***	0.286
RP^-	0.566***	0.093	0.882***	0.194	0.484***	0.123	0.952***	0.233	0.267*	0.144	1.083***	0.240
Constant	5.316***	0.296	-		5.843***	0.394	-		6.800***	0.462	-	
R^2	0.278				0.134				0.131			
$Adj.R^2$	0.270				0.125				0.121			
LR Asym.												
$\chi^2_{LR\_ASYM}$	9.815		3.381		8.248		7.004		25.07		7.744	
P-value	0.002		0.066		0.004		0.008		0.000		0.005	
Coint.												
T-test	-5.096		-2.338		-5.093		-2.743		-3.786		-2.939	
P-value	0.000				0.000				0.003			

Note: (1) The single-step long-run (LR) parameters are determined as $\hat{\beta}^+_T := -\hat{\theta}^+_T/\hat{\rho}_T$ and $\hat{\beta}^-_T := -\hat{\theta}^-_T/\hat{\rho}_T$. For the two-step NARDL, the FM-OLS estimator are employed in the first-step, and the parameters are computed as $\ddot{\beta}^+_T := \ddot{\lambda}_T + \ddot{\eta}_T$ and $\ddot{\beta}^-_T := \ddot{\eta}_T$ using nonlinear combinations based on the delta method.
(2) In the single-step method, the intercept of the cointegrating equation is left unidentified, while bootstrap standard errors (S.E.) derived from 5,000 replications are reported. Asymptotic S.E. are reported for the FM-OLS parameters in the two-step model. The significant level: *** p<0.01, ** p<0.05, * p<0.1
(3) $\chi^2_{LR\_ASYM}$ denotes the Wald test statistic for LR asymmetric coefficients.
(4) Cointegration test: For the single-step method, the t_{BDM} proposed by Banerjee et al. (1998) for testing the null hypothesis of no cointegration is reported. Pesaran et al. (2001) tabulate 5% critical value of t_{BDM} as -3.22, -3.53, and -3.78 for $k = 1, 2,$ and 3, respectively. In the two-step model, the paper uses the Phillips and Perron (1988)'s unit-root test on LR equation residuals. Rejecting the null hypothesis indicates stationary residuals, implying a long-run relationship.

Short-run Parameters

	Minimum Loan Rates MLR				Minimum Overdraft Rate MOR				Minimum Retail Rate MRR			
	Two-step		Single-step		Two-step		Single-step		Two-step		Single-step	
	Estimate	**S.E.**	**Estimate**	**S.E.**	**Estimate**	**S.E.**	**Estimate**	**S.E.**	**Estimate**	**S.E.**	**Estimate**	**S.E.**
Constant	-0.014**	0.006	0.140*	0.075	-0.021***	0.007	0.116**	0.056	-0.026**	0.012	0.121**	0.052
Y_{t-1}			-0.028**	0.012			-0.024***	0.009			-0.023***	0.008
RP^+_{t-1}			0.030***	0.009			0.030***	0.008			0.032***	0.008
RP^-_{t-1}			0.025***	0.009			0.023***	0.008			0.025***	0.007
ECM_{t-1}	-0.021***	0.006			-0.021***	0.006			-0.007**	0.003		
ΔY_{t-1}	0.257***	0.081	0.233***	0.085	0.268***	0.068	0.226***	0.072	0.324***	0.073	0.248***	0.073
ΔY_{t-2}	0.007	0.059	-0.014	0.060								
ΔRP^+_t	0.210***	0.071	0.208***	0.080	0.194***	0.068	0.193**	0.075	0.160**	0.069	0.152**	0.070
ΔRP^+_{t-1}	0.162*	0.088	0.138	0.101	0.167**	0.083	0.142	0.096	0.188**	0.094	0.138	0.092
ΔRP^-_t	0.483***	0.083	0.507***	0.101	0.439***	0.077	0.469***	0.098	0.448***	0.076	0.491***	0.093
ΔRP^-_{t-1}	-0.116*	0.068	-0.096	0.078	-0.076	0.063	-0.054	0.071	-0.116*	0.068	-0.076	0.074
R^2	0.557		0.567		0.526		0.538		0.510		0.538	
$Adj.R^2$	0.545		0.552		0.515		0.524		0.498		0.524	
SR Asym.												
$\chi^2_{SR\_ASYM}$	0.003		0.252		0.000		0.395		0.024		1.134	
P-value	0.959		0.615		0.989		0.529		0.877		0.287	
$\chi^2_{S.Corr}$	0.145		1.486		0.002		0.291		0.074		0.879	
χ^2_{Hetero}	3.874		3.935		11.09		11.67		4.430		5.565	
	NARDL(3,2)				NARDL(2,2)				NARDL(2,2)			

Note: (1) In the single-step method, bootstrap S.E, derived from 5,000 replications are reported. Asymptotic S.E. are reported for the OLS parameters in the two-step model. The significant level: *** p<0.01, ** p<0.05, * p<0.1
(2) $\chi^2_{SR\_ASYM}$ denotes the Wald test statistic for SR asymmetry. $\chi^2_{S.Corr}$ and χ^2_{Hetero} denote the Breusch–Godfrey LM test for autocorrelation and the Breusch–Pagan/Cook–Weisberg test for heteroskedasticity, respectively.

Table 7: Interest Rate Pass-through Results: Small and Medium Banks' Deposit Rates

Long-run Parameters

	Saving Rate				Term Deposits Rate 3 Months				Term Deposits Rate 12 Months			
	Two-step		**Single-step**		**Two-step**		**Single-step**		**Two-step**		**Single-step**	
	Estimate	**S.E.**	**Estimate**	**S.E.**	**Estimate**	**S.E.**	**Estimate**	**S.E.**	**Estimate**	**S.E.**	**Estimate**	**S.E.**
RP^+	0.359***	0.031	0.383***	0.070	0.931***	0.047	0.891***	0.098	1.007***	0.062	0.873***	0.104
RP^-	0.464***	0.044	0.435***	0.062	0.898***	0.065	0.889***	0.088	0.870***	0.087	0.860***	0.099
Constant	0.270*	0.139	-		-0.199	0.209	-		0.213	0.280	-	
R^2	0.726				0.757				0.711			
$Adj.R^2$	0.722				0.754				0.707			
LR Asym.												
$\chi^2_{LR\_ASYM}$	7.809		8.389		0.346		0.007		3.278		0.258	
P-value	0.005		0.004		0.556		0.934		0.070		0.611	
Coint.												
T-test	-4.167		-3.296		-5.303		-2.740		-5.399		-3.181	
P-value	0.001				0.000				0.000			

Note: (1) The single-step long-run (LR) parameters are determined as $\hat{\beta}_T^+ := -\hat{\theta}_T^+/\hat{\rho}_T$ and $\hat{\beta}_T^- := -\hat{\theta}_T^-/\hat{\rho}_T$. For the two-step NARDL, the FM-OLS estimator are employed in the first-step, and the parameters are computed as $\ddot{\beta}_T^+ := \ddot{\lambda}_T + \ddot{\eta}_T$ and $\ddot{\beta}_T^- := \ddot{\eta}_T$ using nonlinear combinations based on the delta method.
(2) In the single-step method, the intercept of the cointegrating equation is left unidentified, while bootstrap standard errors (S.E.) derived from 5,000 replications are reported. Asymptotic S.E. are reported for the FM-OLS parameters in the two-step model. The significant level: *** p<0.01, ** p<0.05, * p<0.1
(3) $\chi^2_{LR\_ASYM}$ denotes the Wald test statistic for LR asymmetric coefficients.
(4) Cointegration test: For the single-step method, the t_{BDM} proposed by Banerjee et al. (1998) for testing the null hypothesis of no cointegration is reported. Pesaran et al. (2001) tabulate 5% critical value of t_{BDM} as -3.22, -3.53, and -3.78 for $k = 1, 2,$ and 3, respectively. In the two-step model, the paper uses the Phillips and Perron (1988)'s unit-root test on LR equation residuals. Rejecting the null hypothesis indicates stationary residuals, implying a long-run relationship.

Short-run Parameters

	Saving Rate				Term Deposits Rate 3 Months				Term Deposits Rate 12 Months			
	Two-step		**Single-step**		**Two-step**		**Single-step**		**Two-step**		**Single-step**	
	Estimate	**S.E.**	**Estimate**	**S.E.**	**Estimate**	**S.E.**	**Estimate**	**S.E.**	**Estimate**	**S.E.**	**Estimate**	**S.E.**
Constant	0.050***	0.017	0.064***	0.023	0.058***	0.018	0.065**	0.027	0.016***	0.006	0.081***	0.027
Y_{t-1}			-0.060***	0.018			-0.053***	0.019			-0.054***	0.017
RP^+_{t-1}			0.023***	0.007			0.047***	0.016			0.047***	0.017
RP^-_{t-1}			0.026***	0.007			0.047***	0.016			0.046***	0.016
ECM_{t-1}	-0.046***	0.015			-0.047***	0.015			-0.029***	0.010		
ΔY_{t-1}	0.303***	0.110	0.261**	0.116	0.410***	0.086	0.409***	0.088	0.413***	0.076	0.408***	0.080
ΔY_{t-2}					-0.063	0.072	-0.059	0.076				
ΔRP^+_t	0.056	0.037	0.068*	0.040	0.182***	0.068	0.174**	0.078	0.253***	0.080	0.236***	0.078
ΔRP^+_{t-1}	0.024	0.030	0.030	0.037	0.154**	0.078	0.143	0.088	0.127	0.086	0.099	0.086
ΔRP^-_t	0.272***	0.039	0.270***	0.050	0.526***	0.044	0.528***	0.061	0.615***	0.058	0.623***	0.078
ΔRP^-_{t-1}	-0.013	0.043	-0.006	0.047	0.041	0.084	0.041	0.089	-0.071	0.080	-0.063	0.083
R^2	0.527		0.545		0.792		0.793		0.757		0.765	
$Adj.R^2$	0.517		0.531		0.786		0.785		0.751		0.757	
SR Asym.												
$\chi^2_{SR\_ASYM}$	7.689		5.049		3.960		3.955		1.842		2.994	
P-value	0.006		0.025		0.048		0.047		0.176		0.084	
$\chi^2_{S.Corr}$	3.004		6.847		3.881		5.217		8.849		10.40	
χ^2_{Hetero}	0.980		0.002		22.84		22.68		14.63		18.37	
	NARDL(2,2)				NARDL(3,2)				NARDL(2,2)			

Note: (1) In the single-step method, bootstrap S.E, derived from 5,000 replications are reported. Asymptotic S.E. are reported for the OLS parameters in the two-step model. The significant level: *** p<0.01, ** p<0.05, * p<0.1
(2) $\chi^2_{SR\_ASYM}$ denotes the Wald test statistic for SR asymmetry. $\chi^2_{S.Corr}$ and χ^2_{Hetero} denote the Breusch–Godfrey LM test for autocorrelation and the Breusch–Pagan/Cook–Weisberg test for heteroskedasticity, respectively.

Table 8: Interest Rate Pass-through Results: Small and Medium Banks' Loan Rates

Long-run Parameters

	Minimum Loan Rates MLR				**Minimum Overdraft Rate MOR**				**Minimum Retail Rate MRR**			
	Two-step		**Single-step**		**Two-step**		**Single-step**		**Two-step**		**Single-step**	
	Estimate	**S.E.**	**Estimate**	**S.E.**	**Estimate**	**S.E.**	**Estimate**	**S.E.**	**Estimate**	**S.E.**	**Estimate**	**S.E.**
RP^+	0.736***	0.067	0.913***	0.182	0.726***	0.074	0.986***	0.200	0.864***	0.116	1.180***	0.160
RP^-	0.484***	0.094	0.688***	0.142	0.429***	0.104	0.734***	0.161	0.105	0.162	0.879***	0.148
Constant	5.637***	0.299	-		6.223***	0.332	-		7.632***	0.518	-	
R^2	0.213				0.120				0.085			
$Adj.R^2$	0.204				0.110				0.075			
LR Asym.												
$\chi^2_{LR\_ASYM}$	9.786		12.45		10.99		14.09		29.54		24.29	
P-value	0.002		0.000		0.001		0.000		0.000		0.000	
Coint.												
T-test	-6.555		-2.669		-6.158		-3.024		-3.811		-3.385	
P-value	0.000				0.000				0.003			

Note: (1) The single-step long-run (LR) parameters are determined as $\hat{\beta}_T^+ := -\hat{\theta}_T^+/\hat{\rho}_T$ and $\hat{\beta}_T^- := -\hat{\theta}_T^-/\hat{\rho}_T$. For the two-step NARDL, the FM-OLS estimator are employed in the first-step, and the parameters are computed as $\ddot{\beta}_T^+ := \ddot{\lambda}_T + \ddot{\eta}_T$ and $\ddot{\beta}_T^- := \ddot{\eta}_T$ using nonlinear combinations based on the delta method.
(2) In the single-step method, the intercept of the cointegrating equation is left unidentified, while bootstrap standard errors (S.E.) derived from 5,000 replications are reported. Asymptotic S.E. are reported for the FM-OLS parameters in the two-step model. The significant level: *** p<0.01, ** p<0.05, * p<0.1
(3) $\chi^2_{LR\_ASYM}$ denotes the Wald test statistic for LR asymmetric coefficients.
(4) Cointegration test: For the single-step method, the t_{BDM} proposed by Banerjee et al. (1998) for testing the null hypothesis of no cointegration is reported. Pesaran et al. (2001) tabulate 5% critical value of t_{BDM} as -3.22, -3.53, and -3.78 for $k = 1, 2$, and 3, respectively. In the two-step model, the paper uses the Phillips and Perron (1988)'s unit-root test on LR equation residuals. Rejecting the null hypothesis indicates stationary residuals, implying a long-run relationship.

Short-run Parameters

	Minimum Loan Rates MLR				**Minimum Overdraft Rate MOR**				**Minimum Retail Rate MRR**			
	Two-step		**Single-step**		**Two-step**		**Single-step**		**Two-step**		**Single-step**	
	Estimate	**S.E.**	**Estimate**	**S.E.**	**Estimate**	**S.E.**	**Estimate**	**S.E.**	**Estimate**	**S.E.**	**Estimate**	**S.E.**
Constant	-0.008**	0.004	0.182**	0.079	-0.016***	0.006	0.174***	0.067	-0.035***	0.013	0.202***	0.066
Y_{t-1}			-0.034***	0.013			-0.031***	0.010			-0.035***	0.010
RP^+_{t-1}			0.031***	0.009			0.030***	0.007			0.041***	0.011
RP^-_{t-1}			0.023***	0.007			0.022***	0.006			0.031***	0.010
ECM_{t-1}	-0.029***	0.008			-0.024***	0.006			-0.008***	0.003		
ΔY_{t-1}	0.376***	0.079	0.356***	0.086	0.352***	0.075	0.321***	0.078	0.255**	0.121	0.171	0.133
ΔY_{t-2}	-0.073	0.051	-0.093	0.059	-0.041	0.055	-0.071	0.061				
ΔRP^+_t	0.183***	0.063	0.184**	0.075	0.154***	0.055	0.155**	0.062	0.173**	0.079	0.154**	0.065
ΔRP^+_{t-1}	0.113*	0.064	0.097	0.077	0.125*	0.064	0.103	0.072	0.191**	0.093	0.121*	0.073
ΔRP^-_t	0.290***	0.069	0.309***	0.085	0.287***	0.075	0.313***	0.090	0.288***	0.095	0.339***	0.117
ΔRP^-_{t-1}	0.019	0.053	0.032	0.059	0.004	0.056	0.023	0.062	0.022	0.068	0.055	0.080
R^2	0.627		0.635		0.578		0.592		0.375		0.425	
$Adj.R^2$	0.617		0.622		0.566		0.578		0.361		0.407	
SR Asym.												
$\chi^2_{SR\_ASYM}$	0.032		0.377		0.025		0.575		0.241		0.826	
P-value	0.858		0.539		0.874		0.448		0.624		0.364	
$\chi^2_{S.Corr}$	5.629		8.441		6.113		9.368		3.172		0.312	
χ^2_{Hetero}	7.961		8.608		21.29		24.44		0.406		23.00	
	NARDL(3,2)				NARDL(3,2)				NARDL(2,2)			

Note: (1) In the single-step method, bootstrap S.E, derived from 5,000 replications are reported. Asymptotic S.E. are reported for the OLS parameters in the two-step model. The significant level: *** p<0.01, ** p<0.05, * p<0.1
(2) $\chi^2_{SR\_ASYM}$ denotes the Wald test statistic for SR asymmetry. $\chi^2_{S.Corr}$ and χ^2_{Hetero} denote the Breusch–Godfrey LM test for autocorrelation and the Breusch–Pagan/Cook–Weisberg test for heteroskedasticity, respectively.

Table 9: Macroeconomic Impact of Policy Rate (Long-run Parameters)

	Dep: Output (CEI growth)				**Dep: Inflation (HCPI growth)**			
	Two-step		**Single-step**		**Two-step**		**Single-step**	
	Estimate	**S.E.**	**Estimate**	**S.E.**	**Estimate**	**S.E.**	**Estimate**	**S.E.**
RP^+	-2.200***	0.446	-1.018***	0.330	-0.275***	0.192	-0.133	0.141
RP^-	1.627**	0.798	0.751*	0.420	0.924***	0.184	0.814***	0.116
Constant	-1448***	162.9	-		-13.85***	2.574	-	
R^2	0.530				0.558			
$Adj.R^2$	0.517				0.546			
LR Asym.								
$\chi^2_{LR\_ASYM}$	20.56		9.093		36.54		40.43	
P-value	0.000		0.003		0.000		0.000	
Coint.								
T-test	-6.468		-5.490		-4.918		-5.234	
P-value	0.000				0.000			

Note: (1) The single-step long-run (LR) parameters are determined as $\hat{\beta}_T^+ := -\hat{\theta}_T^+/\hat{\rho}_T$ and $\hat{\beta}_T^- := -\hat{\theta}_T^-/\hat{\rho}_T$. For the two-step NARDL, the FM-OLS estimator are employed in the first-step, and the parameters are computed as $\ddot{\beta}_T^+ := \ddot{\lambda}_T + \ddot{\eta}_T$ and $\ddot{\beta}_T^- := \ddot{\eta}_T$ using nonlinear combinations based on the delta method.

(2) In the single-step method, the intercept of the cointegrating equation is left unidentified, while bootstrap standard errors (S.E.) derived from 5,000 replications are reported. Asymptotic S.E. are reported for the FM-OLS parameters in the two-step model. The significant level: *** p<0.01, ** p<0.05, * p<0.1

(3) $\chi^2_{LR\_ASYM}$ denotes the Wald test statistic for LR asymmetric coefficients.

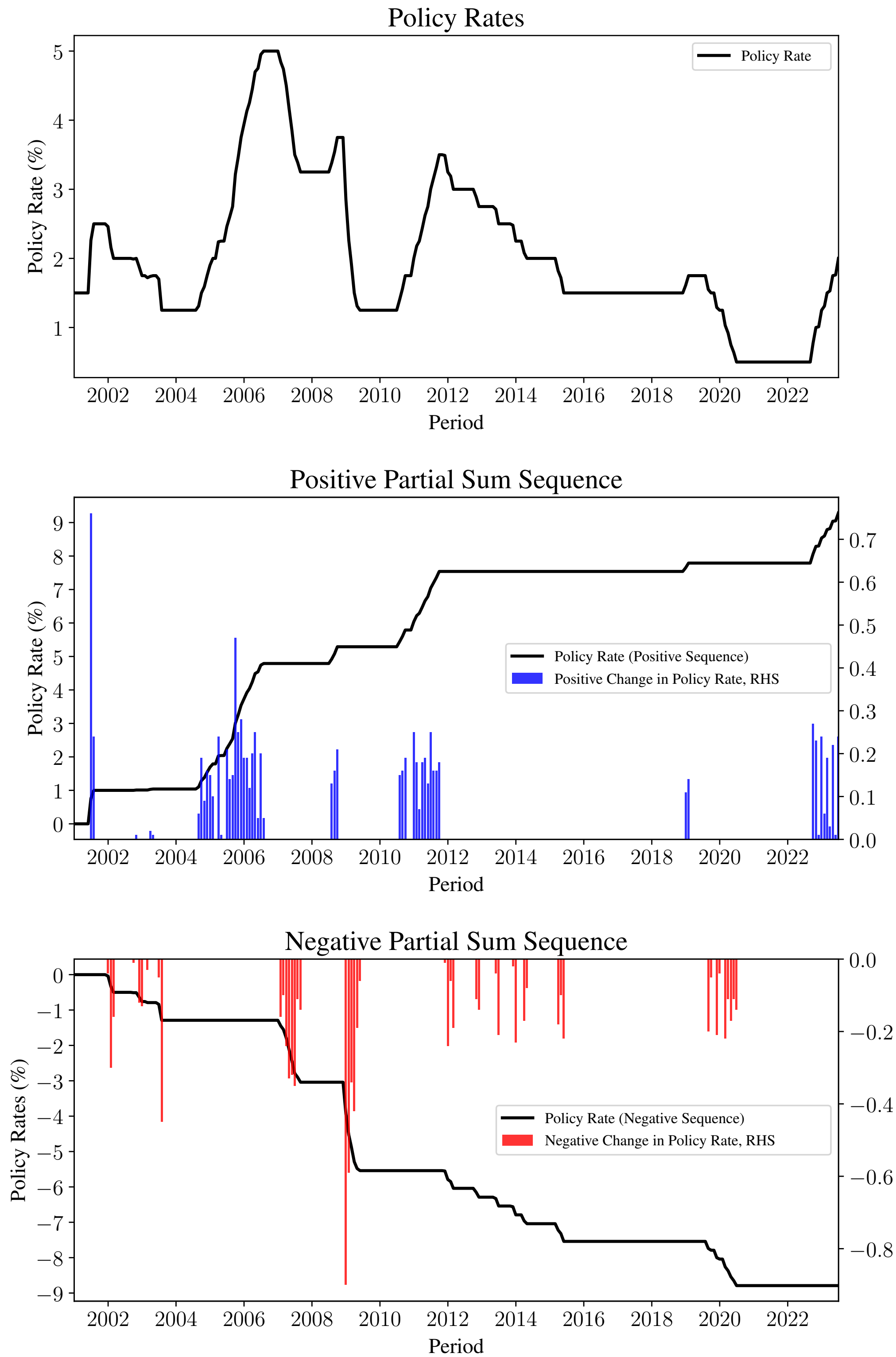
(4) Cointegration test: For the single-step method, the t_{BDM} proposed by Banerjee et al. (1998) for testing the null hypothesis of no cointegration is reported. Pesaran et al. (2001) tabulate 5% critical value of t_{BDM} as -3.22, -3.53, and -3.78 for $k = 1, 2,$ and 3, respectively. In the two-step model, the paper uses the Phillips and Perron (1988)'s unit-root test on LR equation residuals. Rejecting the null hypothesis indicates stationary residuals, implying a long-run relationship.

Table 10: Macroeconomic Impact of Policy Rate (Short-run Parameters)

	Dep: Output (CEI growth)				**Dep: Inflation (HCPI growth)**			
	Two-step		**Single-step**		**Two-step**		**Single-step**	
	Estimate	**S.E.**	**Estimate**	**S.E.**	**Estimate**	**S.E.**	**Estimate**	**S.E.**
Constant	0.212	0.133	4.179***	0.878	0.010	0.044	1.084***	0.189
Y_{t-1}			-0.644***	0.118			-0.514***	0.091
RP^+_{t-1}			-0.655**	0.283			-0.069	0.070
RP^-_{t-1}			0.484	0.310			0.419***	0.099
ECM_{t-1}	-0.344***	0.104			-0.347***	0.073		
ΔY_{t-1}	0.187	0.117	0.373***	0.118	0.364***	0.093	0.440***	0.106
ΔY_{t-2}	0.041	0.103	0.215**	0.092	0.057	0.099	0.174*	0.099
ΔY_{t-3}	-0.008	0.079	0.191**	0.089	0.042	0.098	0.143	0.103
ΔY_{t-4}	-0.102	0.083	0.085	0.085	0.094	0.081	0.185**	0.090
ΔY_{t-5}	0.137*	0.079	0.289***	0.087	0.099	0.065	0.191**	0.080
ΔY_{t-6}	0.083	0.069	0.253***	0.075	0.094	0.060	0.187**	0.074
ΔY_{t-7}	-0.058	0.069	0.134*	0.071	0.042	0.055	0.125*	0.065
ΔY_{t-8}	0.063	0.073	0.217***	0.081	0.019	0.078	0.096	0.082
ΔY_{t-9}	0.262***	0.071	0.407***	0.085	0.106	0.073	0.177**	0.084
ΔY_{t-10}	0.013	0.063	0.160**	0.077	0.171**	0.084	0.259***	0.091
ΔY_{t-11}	0.017	0.075	0.166**	0.073	-0.095	0.073	0.013	0.088
ΔRP^+_t	-4.509***	1.372	-2.318	1.515	1.151*	0.614	1.730**	0.703
ΔRP^+_{t-1}	-1.099	1.683	-0.567	1.851	-0.955	0.761	-0.529	0.740
ΔRP^-_t	-0.344	1.237	0.785	1.643	1.234***	0.301	1.386***	0.396
ΔRP^-_{t-1}	1.292	1.139	2.380*	1.315	-0.608*	0.315	-0.613	0.394
$HCPI^+_{t-1}$			0.234**	0.104				
$HCPI^-_{t-1}$			-0.060	0.093				
$ln(TPGDP)^+_{t-1}$			0.413**	0.191				
$ln(TPGDP)^-_{t-1}$			0.962***	0.294				
CEI^+_{t-1}							-0.009	0.015
CEI^-_{t-1}							-0.007	0.015
$ln(ENERGYP)^+_{t-1}$							0.019***	0.003
$ln(ENERGYP)^-_{t-1}$							0.015***	0.003
R^2	0.265		0.361		0.324		0.382	
$Adj.R^2$	0.216		0.302		0.279		0.325	
SR Asym.								
$\chi^2_{SR\_ASYM}$	12.64		6.009		0.365		0.151	
P-value	0.000		0.014		0.546		0.698	
$\chi^2_{S.Corr}$	8.385				2.237			
χ^2_{Hetero}	3.010				5.697			
	NARDL(12,2)				NARDL(12,2)			

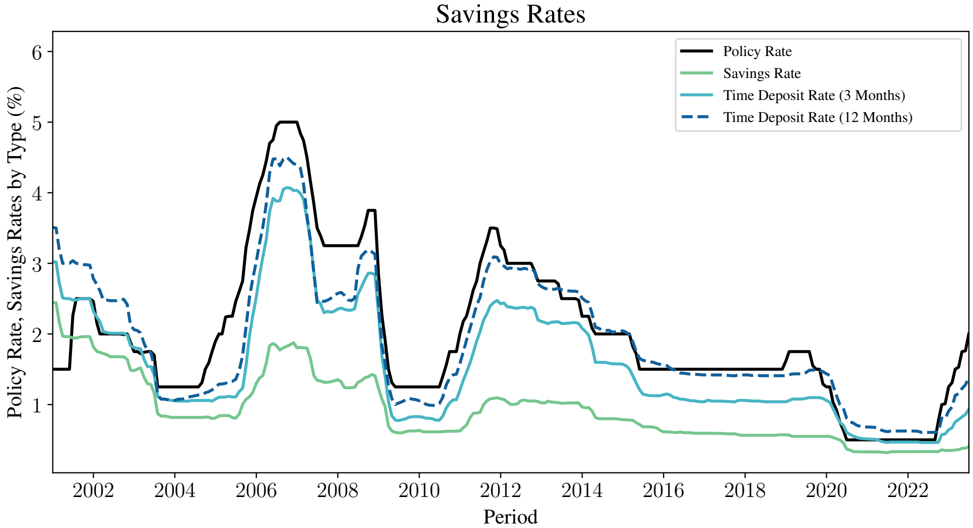
Note: (1) In the single-step method, bootstrap S.E, derived from 5,000 replications are reported. Asymptotic S.E. are reported for the OLS parameters in the two-step model. The significant level: *** p<0.01, ** p<0.05, * p<0.1
(2) $\chi^2_{SR\_ASYM}$ denotes the Wald test statistic for SR asymmetry. $\chi^2_{S.Corr}$ and χ^2_{Hetero} denote the Breusch–Godfrey LM test for autocorrelation and the Breusch–Pagan/Cook–Weisberg test for heteroskedasticity, respectively.

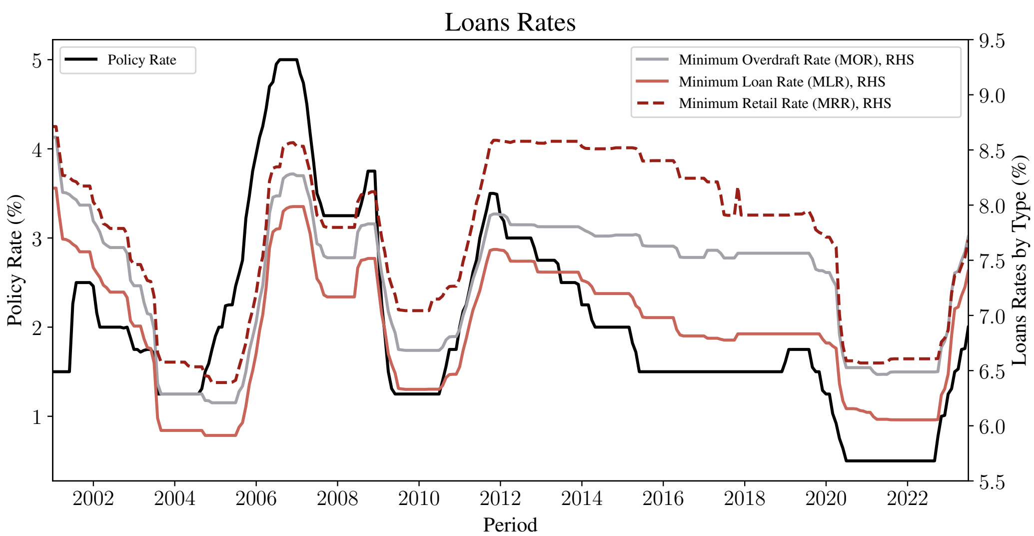
Figure 1: Thailand's Policy Rate



Source: The BOT.
Notes: The BOT has changed its policy interest rate from 14-day RP to 1-day RP since 2007.

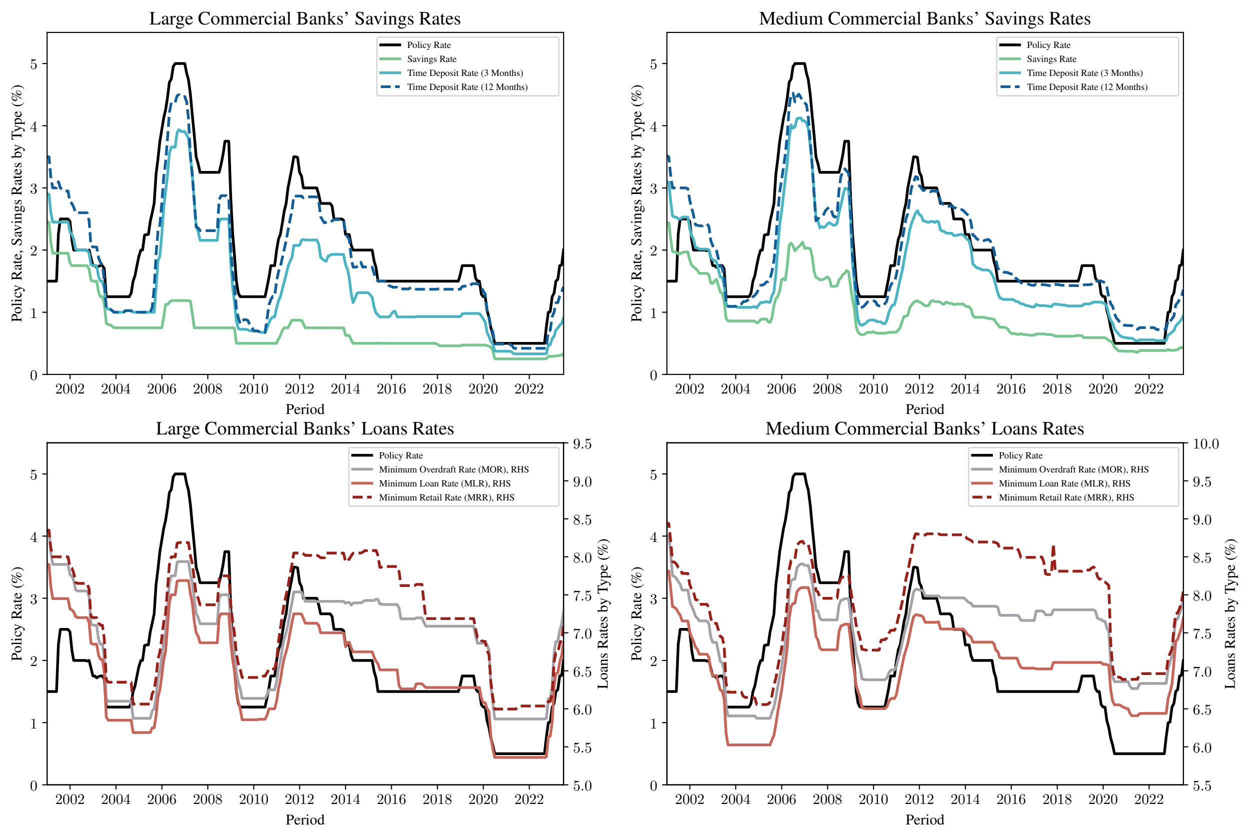
Figure 2: Thailand's Policy Rate and Commercial Banks' Interest Rate





Source: The BOT.
Notes: (1) The BOT has changed its policy interest rate from 14-day RP to 1-day RP since 2007.
(2) Each bank's interest rates are calculated as the average between the minimum and maximum rates reported to the BOT, which is represented as $0.5 \times (min. + max.)$.
(3) For average data, it is computed as the average from individual banks' daily data.

Figure 3: Thailand's Policy Rate and Commercial Banks' Interest Rate
Classified by Bank Sizes

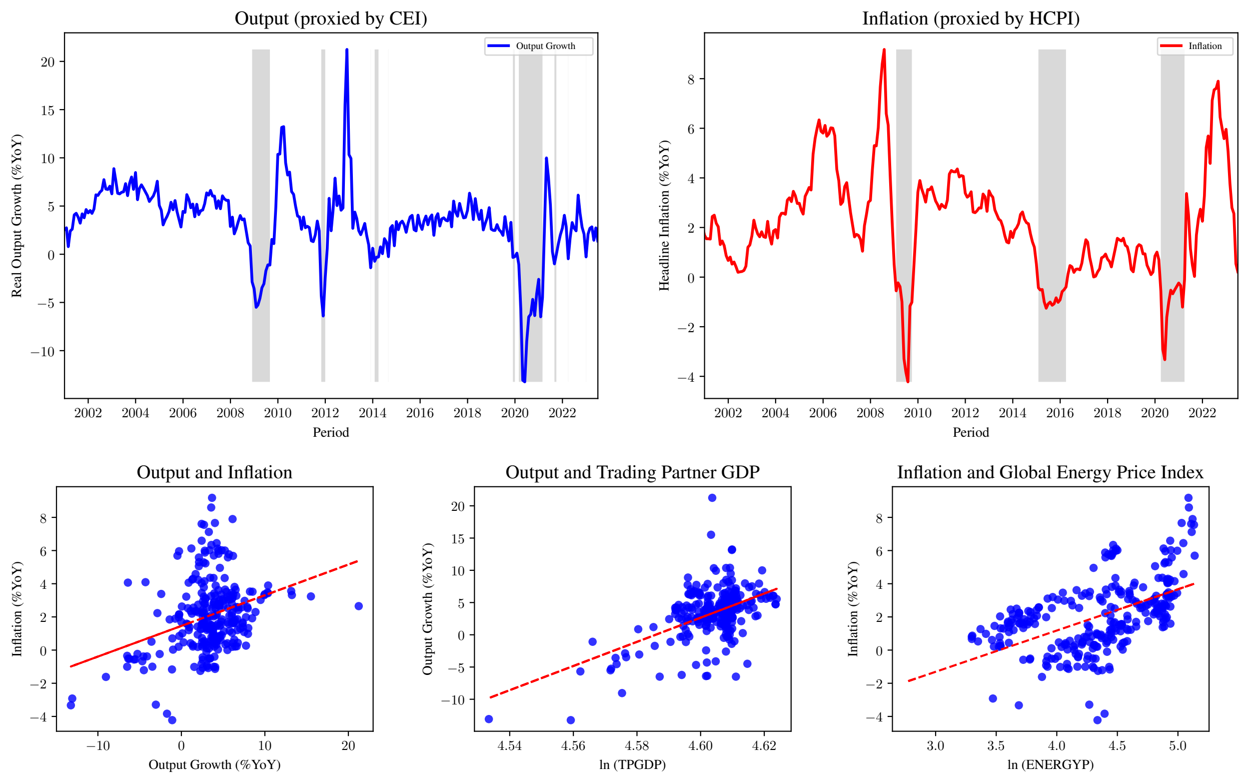


Source: Daily Interest Rates of Commercial Banks Reported on the BOT's Website

Notes: (1) Each bank's interest rates are calculated as the average between the minimum and maximum rates, $rate_i = 0.5 \times (min_i + max_i)$, while the data shown in the graph is the simple average of all banks classified by group.

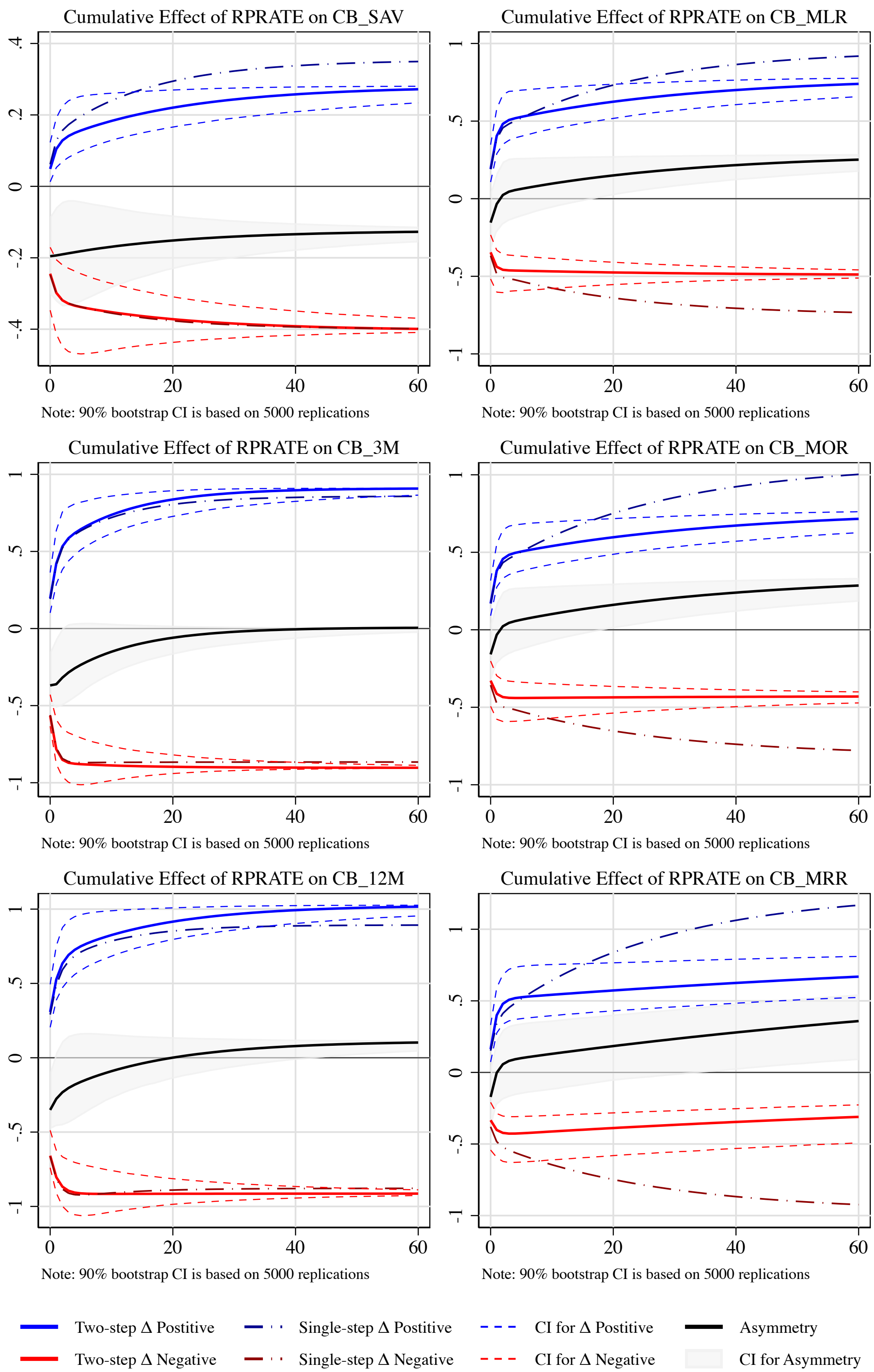
(2) The data was based on daily interest rate of individual banks data provided by the BOT, which can be accessed via: https://www.bot.or.th/en/statistics/interest-rate.html

Figure 4: Thailand's Macroeconomic Development



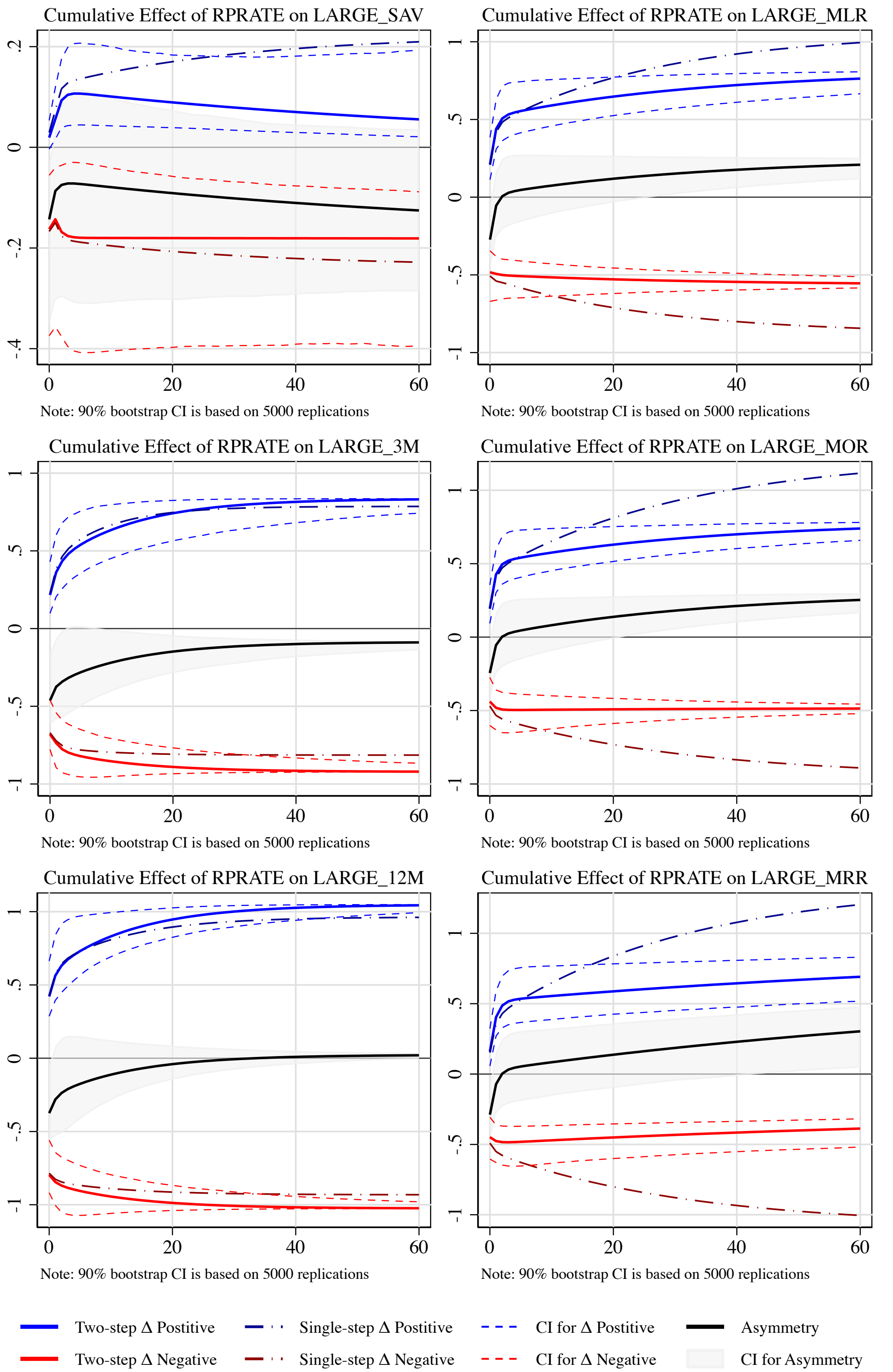
Source: The BOT, Thailand's Ministry of Commerce, FRED data, and OECD dataset.

Figure 5: Dynamic Interest Rate Pass-through



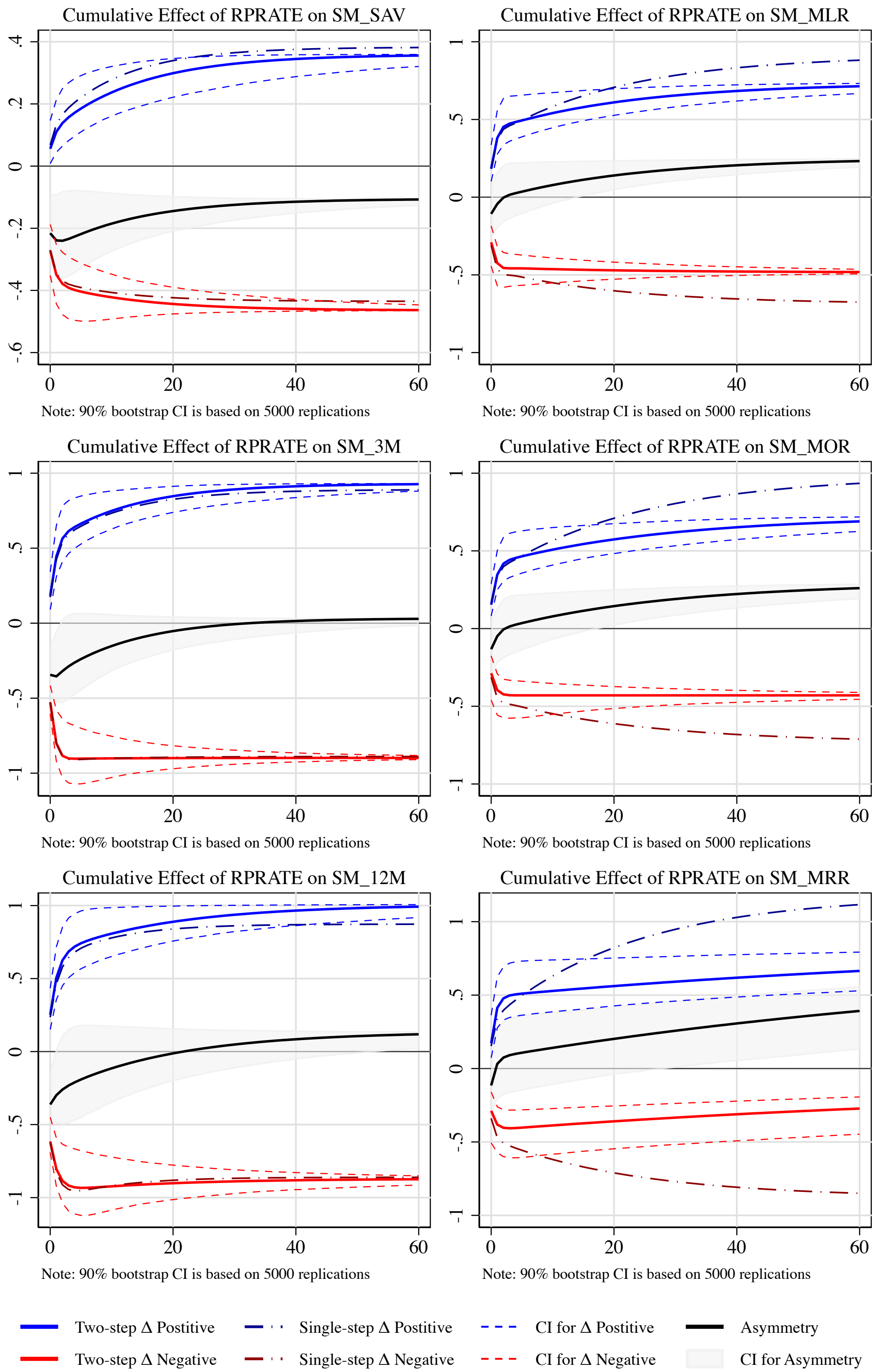
Note: For ease of interpretation, the paper reports the inverse impact results of negative changes. Therefore, it refers to a change in interest rate as -1.

Figure 6: Dynamic Interest Rate Pass-through: Large Banks



Note: For ease of interpretation, the paper reports the inverse impact results of negative changes. Therefore, it refers to a change in interest rate as -1.

Figure 7: Dynamic Interest Rate Pass-through: Small and Medium Banks



Note: For ease of interpretation, the paper reports the inverse impact results of negative changes. Therefore, it refers to a change in interest rate as -1.

Figure 8: Dynamic Effects of Policy Rate Shocks on Output

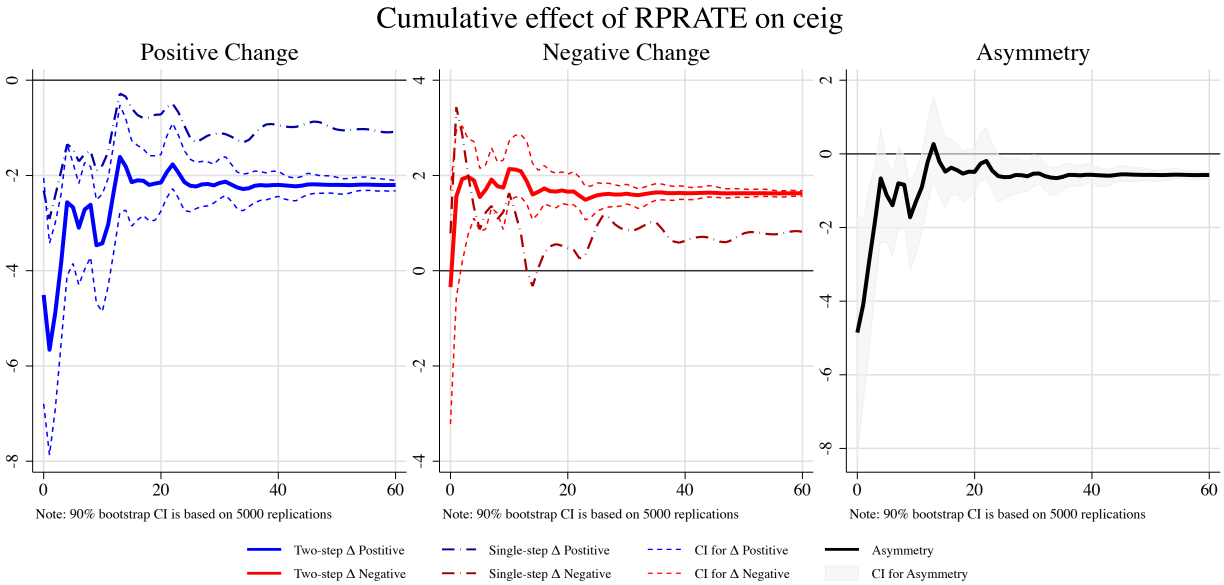
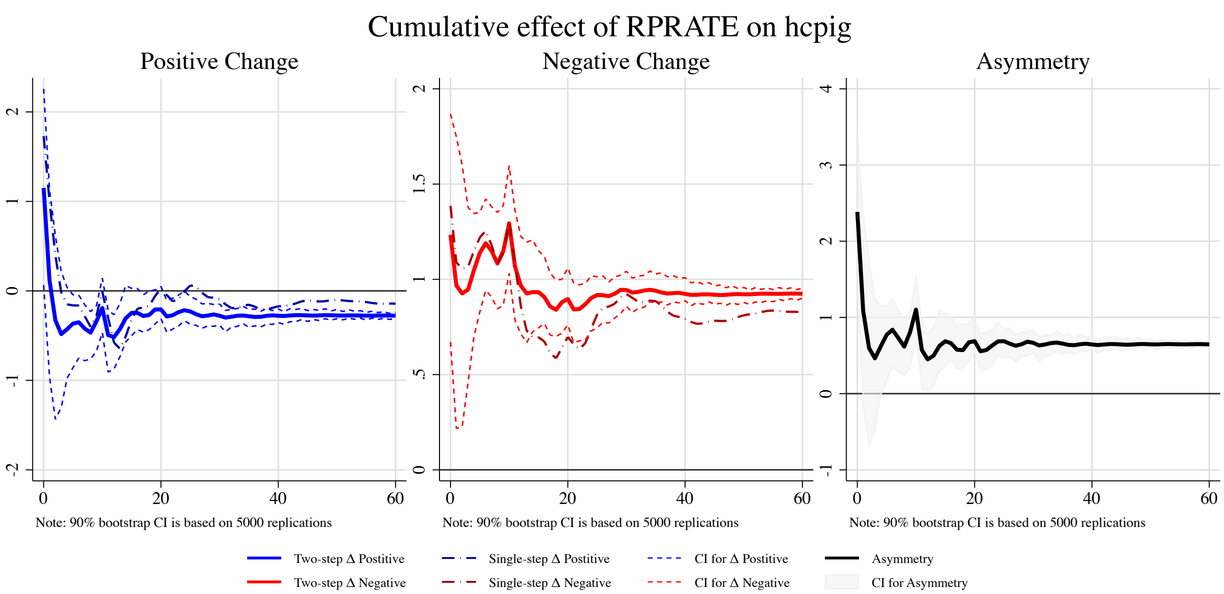


Figure 9: Dynamic Effects of Policy Rate Shocks on Inflation



A Appendix

In the Appendix, we present preliminary testing results before applying the 2SNARDL method. We apply Dickey and Fuller (1981) augmented test to the variables used in this study to assess their stationarity. We test the nonstationary data hypothesis by including only the intercept or by including both the intercept and time trend. The test results are presented in Table A.1. During the test, we determine the lag of the model using the SIC and set the level of significance to 0.01. Consequently, all the investigated variables turn out to be nonstationary.

Table A.1: The Unit-root Test Results (Augmented Dickey-fuller Test)

Variables	Level						First Difference					
	Intercept			Intercept with Trend			Intercept			Intercept with Trend		
	Test-stat	Lags	Results	Test-stat	Lags	Results	Test-stat	Lags	Results	Test-stat	Lags	Results
RP	-2.460	4	Unit-root	-2.905	4	Unit-root	-5.600	3	Stationary	-5.583	3	Stationary
CB_SAV	-1.996	2	Unit-root	-2.483	2	Unit-root	-9.491	1	Stationary	-9.488	1	Stationary
CB_3M	-2.800	4	Unit-root	-3.279	4	Unit-root	-5.557	3	Stationary	-5.553	3	Stationary
CB_12M	-3.028	4	Unit-root	-3.400	4	Unit-root	-5.298	3	Stationary	-5.298	3	Stationary
CB_MLR	-3.475	4	Stationary	-3.387	4	Unit-root	-4.436	3	Stationary	-4.490	3	Stationary
CB_MOR	-3.307	4	Unit-root	-3.258	4	Unit-root	-4.569	3	Stationary	-4.611	3	Stationary
CB_MRR	-2.603	4	Unit-root	-2.589	4	Unit-root	-5.001	3	Stationary	-5.005	3	Stationary
$LARGE\_SAV$	-2.990	2	Unit-root	-2.770	2	Unit-root	-11.37	1	Stationary	-11.51	1	Stationary
$LARGE\_3M$	-3.078	4	Unit-root	-3.725	4	Unit-root	-4.955	3	Stationary	-4.955	3	Stationary
$LARGE\_12M$	-3.381	4	Unit-root	-3.812	4	Unit-root	-6.433	1	Stationary	-6.423	1	Stationary
$LARGE\_MLR$	-3.271	4	Unit-root	-3.315	4	Unit-root	-4.618	3	Stationary	-4.674	3	Stationary
$LARGE\_MOR$	-3.265	4	Unit-root	-3.167	4	Unit-root	-7.238	1	Stationary	-7.269	1	Stationary
$LARGE\_MRR$	-2.234	2	Unit-root	-2.190	2	Unit-root	-7.202	1	Stationary	-7.203	1	Stationary
SM_SAV	-2.293	4	Unit-root	-3.157	4	Unit-root	-5.978	3	Stationary	-5.982	3	Stationary
SM_3M	-2.660	2	Unit-root	-3.267	2	Unit-root	-6.160	1	Stationary	-6.143	1	Stationary
SM_12M	-2.618	2	Unit-root	-3.000	2	Unit-root	-6.632	1	Stationary	-6.620	1	Stationary
SM_MLR	-2.448	2	Unit-root	-2.397	2	Unit-root	-6.891	1	Stationary	-6.920	1	Stationary
SM_MOR	-3.320	4	Unit-root	-3.336	4	Unit-root	-7.158	1	Stationary	-7.164	1	Stationary
SM_MRR	-1.993	2	Unit-root	-2.025	2	Unit-root	-7.800	1	Stationary	-7.789	1	Stationary
CEI	-4.904	1	Stationary	-5.459	1	Stationary	-6.405	12	Stationary	-6.393	12	Stationary
$HCPI$	-3.657	2	Stationary	-3.765	2	Unit-root	-6.502	12	Stationary	-6.462	12	Stationary
$lnTPGDP$	-4.621	5	Stationary	-4.660	5	Stationary	-5.525	7	Stationary	-5.525	7	Stationary
$lnENERGYP$	-2.393	2	Unit-root	-2.325	2	Unit-root	-9.969	1	Stationary	-9.977	1	Stationary

Note: The unit-root test results are interpreted at a significant level of 0.01. Lags selection is based on SIC.