Asymmetric Interest Rate Pass-through and Its Effects on Macroeconomic Variables in 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

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Abstract

This study employs the two-step Nonlinear Autoregressive Distributed Lag (NARDL) model, as proposed by Cho, Greenwood-Nimmo, and Shin (2019), to identify the asymmetric impact of monetary policy on economic variables using monthly data from Thailand spanning the period from 2001 to 2023. Its primary objective is to investigate the effects of policy rate shocks on the economy. The paper examines three key aspects: (i) the asymmetric pass-through of policy rates to commercial bank deposit and loan rates; (ii) pass-through variations across bank sizes; and (iii) asymmetric macroeconomic effects on output and inflation. The empirical findings reveal the presence of asymmetry within the relationships of the variables. Firstly, the study identifies incomplete interest rate pass-through, with deposit rates ranging from 28.1% to 102.7% and loan rates from 12.7% to 89.6%. Notably, long-term upward asymmetry is observed for loan rates, while evidence for deposit rates is limited. Secondly, concerning bank sizes, large banks exhibit a greater pass-through effect on loan rates, whereas small and medium-sized banks display higher responsiveness in short-term deposit or savings rates. Lastly, the study provides strong evidence of long-term asymmetric macroeconomic impact. Quantitatively, rate hikes have a more substantial effect, being 1.3 times larger than the impact of rate cuts, on both output growth and inflation. These findings emphasize the significant role of contractionary monetary policy in controlling inflation in Thailand, surpassing the impact of expansionary policy.

Key Words: Interest rate pass-through; asymmetric impact; macroeconomic effects; NARDL model.

Subject Classification: E43, E51, E52.

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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 asymmetric impacts on the economy. For instance, Bernanke (1996) has noted that financial frictions exist in the financial market, resulting in varied effects on the economy when policy interest rates change. Another example can be seen in the survey conducted by Égert and MacDonald (2009) which explores various monetary transmission mechanisms in the Central and Eastern European financial market, revealing their complexity that cannot be comprehensively understood within a single framework.

The identification of asymmetric monetary policy effects is often a 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.

The primary objective of this paper is to examine the asymmetric effects of interest rate shocks on the economy. Our main focus is to analyze the differences in effects between increases and decreases in policy rates. To accomplish this, we employ the Nonlinear Autoregressive Distributed Lag (NARDL) model to investigate the transmission of policy rates and their impact on macroeconomic variables. Specifically, we aim to empirically apply the two-step NARDL (2SNARDL) estimation method as proposed by Cho et al. (2019), demonstrating 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 their 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 asymmetric pass-through, drawing on the foundational markup pricing model by Rousseas (1985). Subsequently, we explore variations in pass-through across banks of different sizes. Finally, we investigate how key macroeconomic variables, including inflation and economic growth, respond to interest rate changes. This analysis takes into account economic indicators of trading partners and global energy prices, providing a comprehensive understanding of these relationships. Consistent with recent literature, we anticipate observing asymmetric effects of interest rate shocks on these macroeconomic variables.

Recent studies have challenged the traditional view of complete and symmetric monetary policy pass-through, 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., Gigineishvili, 2011; Hannan and Berger, 1991), economic uncertainty (e.g., Égert and MacDonald, 2009), liquidity conditions (e.g., Cho, Greenwood-Nimmo, and Shin, 2023; Gigineishvili, 2011), financial frictions (e.g., Bernanke, Gertler, and Gilchrist, 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 impact of interest rates. The first strand examines state-dependent asymmetry, aiming to differentiate the effects of interest rate shocks during economic slowdowns and expansions (e.g., see Lo and Piger, 2005; Peersman and Smets, 2005). The second strand focuses on the direction of interest rate changes. For example, Debortoli, Forni, Gambetti, and Sala (2020) and Hayford (2006) revealed 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 our study.

In terms of methodology, we have chosen the NARDL model due to its practicality, simplicity, and effectiveness in analyzing asymmetric phenomena, as highlighted by Cho et al. (2023). This choice diverges from most existing literature on interest rate pass-through, which primarily relies on error-correction models (ECMs), dynamic OLS, and autoregressive distributed lag (ARDL) models. These conventional models have limited capacity to capture 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 through cumulative dynamic multipliers. As a result, the NARDL model has gained popularity in prior literature for analyzing both interest rate pass-through (e.g., Greenwood-Nimmo, Shin, van Treeck, and Yu, 2013; Apergis and Cooray, 2015; Galindo and Steiner, 2022; Yu, Chun, and Kim, 2013) and investigating 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 ordinary least squares (OLS) estimation, as proposed by Shin, Yu, and Greenwood-Nimmo (2014). However, in this paper, we employ the 2SNARDL method introduced by Cho et al. (2019). We have made this choice due to the recognition that the single-step method may encounter issues related to the asymptotic singular matrix problem. The 2SNARDL model overcomes this challenge and offers the advantage of analytical tractability. In summary, the 2SNARDL approach involves initially estimating the long-run relationship using the fully-modified OLS (FM-OLS) estimator by Phillips (please provide the full citation), depending on the number of explanatory variables. Subsequently, the short-run dynamic parameters are estimated using OLS, as detailed in Section 3.

To the best of our knowledge, research on interest rate pass-through in Thailand is limited, both in terms of 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, it was estimated that both deposit and loan rates exhibited a pass-through rate of less than 50% (e.g., Disyatat and Vongsinsirikul, 2003; Charoenseang and Manakit, 2007). More recently, Yu et al. (2013) explored asymmetric interest rate pass-through using Thai data from 2000 to 2009, revealing 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 Waiquamdee and Boonyatotin, 2008). Therefore, it is imperative to provide a more comprehensive and up-to-date analysis.

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

The paper is organized as follows. Section 2 provides a summary of the related theoretical background and empirical studies, along with the motivation for this research. In Section 3, we introduce the details of the NARDL model. Section 4 provides insights into the data used in this study. We present the empirical results in Section 5, and Section 6 concludes the paper. In the Appendix, we include 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 motivate the current study. We separately review the empirical literature after first reviewing the theoretical literature.

2.1 Theoretical Literature Review

For the goal of this study, it is essential to understand the interest rate channel thoroughly within the 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 the incomplete and asymmetric interest rate pass-through, we first briefly review the literature on the interest channel.

The central bank employs its 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 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. These policy-controlled interest rates are often approximated using central bank policy rates and money market rates. This linkage is described by the markup pricing model:

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

where r_t^B represents 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 though lending and deposit rates, thereby adjusting consumption and investment, ultimately affecting aggregate demand and prices. Clarida, Galí, and Gertler (1999) provide a macroeconomic model describing the connection between 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 (AS-AD) 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 current inflation rate, $\hat{y}_t$ is the output gap, $\hat{r}_t$ is the interest rate, $\hat{r}_t - E_t \hat{\pi}_{t+1}$ implies real interest rate, ϵ_t^s is a cost-push shock, ϵ_t^d is a demand shock, and ϵ_t^R is a monetary policy shock.

The economic intuitions of this framework are as follows. First, the New Keynesian Phillips Curve (NKPC) plays the role of aggregate supply. It signifies that current inflation depends on anticipated future inflation, the output gap, and other economic shocks. The NKPC is typically derived from the pricing decisions of firms, 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 (IS) curve illustrates how private agents decide on consumption and investment based on their forward-looking economic outlook from 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 (MP) schedule represents how the interest rate reacts to fluctuations in inflation and the output gap. This response is governed by the monetary policy rule, typically the Taylor rule. For instance, if inflation rate exceeds the target while 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 the aggregate money supply. Its applications are widely observed 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 the delayed consumption of households, 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, following Taylor's rule, after the initial shock period.

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. In particular, when policy interest rate changes quickly and have a full impact on retail rates, they 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 the conventional monetary theory, which assumes pass-through to be symmetrical, rapid, and complete. Comprehensive surveys conducted by Andries and Billon (2016) and Gregor et al. (2021) have identified several factors influencing the magnitude of pass-through. These factors include marginal pricing costs, the bank's market power (or degree of competition), the bank's balance sheet, liquidity constraints, the development of financial markets, and prevailing economic conditions. In addition to these factors, Grigoli and Mota (2017) also provide a theoretical background to explain the stickiness and asymmetry of retail rate adjustments in response to policy rate changes. These phenomena are primarily attributed to incomplete market competition and information asymmetry.

We can summarize these factors as follows:

(a) First, market structure and the degree of competitiveness significantly influence the interest rate channel. In markets characterized by limited competition, banks may have more market power. 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) investigated price rigidity within the banking industry and concluded 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 financial instability, banks often adopt a more cautious stance, becoming 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, as prices are adjusted more frequently.

(c) Third, liquidity conditions also play a significant role. The availability of funds in the 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, as it depends on how monetary policy affects the broader liquidity landscape, the cost of funds, and the cost of credit (see Cho et al., 2023). Empirically, Gigineishvili (2011) found 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 a significant role in the stickiness of retail rates. When banks perceive an increased risk of default, they tend to maintain a spread between lending and deposit rates. Moreover, they may resist raising their rates, even when their funding costs increase, in order to mitigate the heightened default risk within their loan portfolio. This implies upward rigidity in loan rate adjustments. This notion finds support in the research conducted by De Bondt (2005), who also suggested that banks may be reluctant to fully transmit an increase in the policy rate 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 a vicious cycle effect of interest rate changes on the economy, primarily through their influence on firm balance sheets and credit conditions. These

[1] The paper includes 398 bank data in 132 markets and measures security-rate by 3-month T-bill rates.

factors can 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. In practical terms, 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 decreases. Regarding lending rates, a cross-country analysis conducted by Borio and Fritz (1995) demonstrated that the adjustment of lending rates is swifter in response to upward rate changes in certain countries.

2.2 Empirical Literature Review

Motivated by the theoretical studies on asymmetric interest rate pass-through, numerous studies have investigated its properties both qualitatively and quantitatively. These studies aimed to elucidate how long-term retail interest rates, including lending and deposit rates, respond to changes in policy rates by estimating empirical models stemming from (1). It is important to note that empirical evidence on this topic may vary depending on factors such as the bank location, the sample period, and the estimation methods used. Most of the studies applied cointegration analysis in the form of error-correction models, dynamic OLS, and autoregressive distributed lag (ARDL) models. It is commonly observed that the pass-through process is incomplete, meaning that the adjustment of 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 to be 0.803, with a median of 0.854.

Recent studies have begun to recognize a nonlinear cointegration structure in (1) and have estimated asymmetric models of monetary policy pass-through. For instance, Greenwood-Nimmo et al. (2013) applied NARDL to U.S. data to examine how banks adjust deposit and lending rates in response to changes in policy-controlled rates.[2] Their results provide strong evidence of asymmetry. Specifically, they found that in the long run, rate cuts are passed through more completely than rate hikes, while in the short run, rate hikes are transmitted more rapidly than rate cuts. Additionally, the degree of pass-through has diminished during the great moderation. Similarly, Apergis and Cooray (2015) applied NARDL to data from the U.S., the UK, and Australia, observing positive asymmetry for lending rates and noting negative asymmetry regarding deposit rates. In other words, for deposit interest rates, decreases exhibit a stronger degree of interest rate pass-through, whereas increases in lending rates demonstrate a greater degree of pass-through than decreases. Furthermore, Galindo and Steiner (2022) examined the case of Colombia, suggesting upward rigidity in

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

lending rates, indicating that they respond more to rate cuts than to hikes. This 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 into 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 22 to 50% for savings rates. For instance, Disyatat and Vongsinsirikul (2003) examined the degree of pass-through of money market rates to the minimum lending rate (MLR) and 3-month deposit rates, using the dynamic multiplier (DM) model and error-correction model (ECM) with Thai data from 1989 to 2002. In pre-crisis 1997 samples, deposit rates adjusted by around 70% and 50% in response to RP14 rate changes by DM and ECM, respectively, while lending rates adjusted by 56% (DM) and 40% (ECM). However, the results were considerably lower for the 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) examined 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 the IT monetary policy by the MPC. Their findings indicated that commercial banks charged a higher loan markup than deposits, resulting in an incomplete pass-through. Specifically, the average pass-through for the loan rate was 26.5%, whereas for the savings deposit rate, the average pass-through was 21.9%. Furthermore, pass-through was higher for longer-term rates, exceeding 50%.

In another empirical study on asymmetric pass-through, Yu et al. (2013) examined 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 indicated that the short-term market rate is cointegrated with the bank lending rate. The estimated coefficients suggested a pass-through degree ranging from 69% to 82% when assuming long-run symmetry. However, when assuming asymmetric pass-through over the long run, positive increases in market rates were found to pass through at around 76% to 77%, whereas negative changes passed through more than one-to-one, at around 103%. This implies that Thailand's lending rates exhibit negative asymmetric pass-through over the long run.

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

Recent studies have empirically revealed the presence of asymmetric effects resulting from monetary policy shocks, considering factors such as confidence levels, wage rigidity, prevailing economic conditions, credit conditions, and price inflexibility (see Pierre-Richard, 2001). There are two distinct lines of research that stand out. The first strand examines asymmetry across different economic states, showing that monetary policy shocks have a more pronounced influence on economic activity during economic slowdowns compared to expansions. This phenomenon has been supported by studies conducted by Lo and Piger (2005) and Peersman and Smets (2005), who explored data from the U.S. and the Euro area, 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) applied SVAR with exogenous variables (SVARX) to U.S. data and found 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 suggested that positive shocks to the Federal funds rate have a more considerable 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) applied the NARDL model to South African data and found that positive monetary policy shocks had a more significant effect on output growth, while the results were reversed for inflation. Furthermore, the asymmetric impact of monetary policy shocks intensified when government borrowing was taken into account. In a separate investigation, Gocer and Ongan (2020) examined UK data to explore Fisher effect. They identified an asymmetric coefficient relationship between inflation and interest rates during the period from 1995 to 2008, although this evidence weakened after 2008. Additionally, Claus and Nguyen (2020) utilized the NARDL model on Australia's consumer survey data to analyze consumer responses to monetary policy shocks. Their findings provide substantial evidence of asymmetric reactions, both in the short and long run.

The developments in the literature motivate 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 aim to explore the asymmetric features of the interest rate channel while considering the unique aspects of the Thai economy. For this purpose, we specifically consider the bank size effect in Thailand as we detail below.

3 Methodology

Before examining the Thai economy by NARDL, we first briefly outline the NARDL model and its estimation procedure. The development of the NARDL model estimation is comprehensively summarized by Cho et al. (2023). The NARDL model builds upon the 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), which assumed a linear long-run relationship. This was followed by the NARDL model introduced by Shin et al. (2014), any they estimate long-run and short-run parameters simultaneously using the ordinary least squares (OLS) method. This approach allows for partial sum decompositions of the explanatory variables, accommodating asymmetrical relationships. In addition, Cho et al. (2019) propose a two-step estimation method to address an asymptotic singularity problem that arises when the OLS estimation is applied.

There are several advantages to using the NARDL model. First, the researcher can utilize partial sum decompositions of 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 the researcher to jointly analyze data with non-stationary and nonlinearity issues, such as the mixing of I(1) or I(0) data. Second, as highlighted by Cho et al. (2023) and Shin et al. (2014), NARDL also allows researchers to construct cumulative dynamic multipliers, providing easily interpretable visualizations of the dynamic adjustment toward an equilibrium position following a shock. Moreover, cumulative dynamic multipliers do not rely on controversial procedures for structural shock identification.

The NARDL model captures a nonlinear cointegrating relationship by decomposing the positive and negative influences of exogenous variables. Suppose 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 $\Delta \boldsymbol{x}_j^- := \min[\mathbf{0}, \Delta \boldsymbol{x}_j]$. In other words, $\boldsymbol{x}_t$ is decomposed into $\boldsymbol{x}_t^+$ and $\boldsymbol{x}_t^-$ around zero. We suppose 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 a zero mean and constant variance. We 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 equation above captures 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) into an 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 $\boldsymbol{\beta}_*^{+\prime} := -(\boldsymbol{\theta}_*^{+\prime}/\rho_*)$ and $\boldsymbol{\beta}_*^{-\prime} := -(\boldsymbol{\theta}_*^{-\prime}/\rho_*)$ are the asymmetric long-run parameters, and $\boldsymbol{\pi}_{j*}^{+\prime}$ and $\boldsymbol{\pi}_{j*}^{-\prime}$ are the parameters capturing nonlinear short-run dynamics.

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

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

This estimation is done 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 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 exist, viz., $k > 1$, Cho et al. (2019) further 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 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. Then, they 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 for the long-run parameters do not suffer from singularity issues. The long-run parameter estimates obtained this way are super-consistent.

Next, 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 distinct convergence rates of the two-step estimators, resulting in consistency, and the long-run and short-run parameters asymptotically follow mixed normal and normal distributions, respectively. Specifically, as the long-run coefficients are estimated using FM-OLS with faster convergence rates than OLS for short-run parameters, we can treat the long-run parameters as known when estimating 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. For cointegration testing, as the 2SNARDL method is based on the two-step ECM form, the test can be performed by examining the stationarity of the fitted residuals obtained from the cointegrating equation. Cho et al. (2019) employs the 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 in our analysis of the Thai economy data, which involves a three-step process. Initially, we validate the applicability of the 2SNARDL estimation to our collected data. In this regard, we furnish 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 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 interest rate pass-through and its broader effects on the economy.

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

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 itself.

(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: the log trading partner GDP for output growth and the 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 the existence of a long-run 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-run and short-run asymmetry.

(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 t_{BDM}-test as proposed by Banerjee, Dolado, and Mestre (1998).

(f) Finally, we compute dynamic multipliers 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 for the empirical study before discussing the empirical estimation and inference results.

We obtained monthly data required for this study from Thai authorities, covering the period from 2001 to 2022. This timeframe aligns with the adoption of the IT framework by the Bank of Thailand (BOT), where the policy interest rate served as its primary tool[3]. Table 1 provides a summary of the data statistics.

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

We partition the economic variables into three groups and first detail Thailand's policy rates (RP) given in Table 1. The policy rates have averaged approximately 2.01% annually over the past two decades. The lowest recorded rate stands at around 0.5%, while the highest reaches approximately 5%. The median rate is below the mean, registering at 1.5%. When considering incremental rate changes, rate increases and decreases 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 the historical trend in the policy rates. There have been three distinct cycles of interest rate increases and four significant periods of rate cuts. We can classify them into five categories as follows:

(a) First, until 2004, policy rates remained relatively low, fluctuating between approximately 1.25% and 2.5%. However, from 2004 to 2006, the MPC raised policy rates significantly, implementing three rate hikes in the second half of 2004, six in 2005, and four in 2006. This led to a policy rate increase from around 2% in 2004 to 4% by the end of 2005, peaking at 5% in June 2006. These measures aimed to address upward price pressures resulting from improved domestic economic growth, persistently high oil prices, and the global tightening of monetary policy 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, inflationary pressures eased. Consequently, the MPC reduced the policy interest rate five times in the first half of 2007, lowering it from 5% to 3.25%. As 2008 approached, the global economy was anticipated to experience slow growth due to 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 by 1.5% over three consecutive meetings in 2009, ultimately bringing the policy rate down 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%. Subsequently, in 2011, the MPC implemented an additional six consecutive rate hikes, totaling a 1.25% increase. By October 2011, the policy rate had reached 3.5%. However, from 2012 to 2014, internal political uncertainty disrupted the domestic economy. In response, the MPC decided to gradually reduce policy rates, ultimately lowering them to 2% by the end of 2014. In 2015, the MPC made another adjustment by cutting the rate by 0.5% to 1.5%. This action was aimed at supporting economic recovery, particularly in response to the global economic slowdown, which had a notable impact on China and

Asia.

(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 due to structural issues, including declining competitiveness in certain export sectors. This period was marked by slow growth rate and persistently low inflation rate, occasionally even dipping into negative territory. Moreover, during this timeframe, global financial markets experienced excess liquidity and sustained low global interest rates. In response, the MPC gradually reduced the policy rate from 2019 to 2020, ultimately reaching its lowest point at 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 adverse effects from the COVID-19 situation 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. Continuing their efforts in 2023, the MPC implemented additional rate hikes, bringing the policy rate to 2% in the first half of the year. These decisions were taken 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 also align with the global trend of interest rate normalization.

Next, we detail the interest rates offered by commercial banks in Thailand. Thailand's financial system has a substantial dependence on commercial banks, which comprised approximately 47%, nearly half, of the total assets of financial institutions as of Quarter 3 of 2021. Furthermore, they account for about 70% of corporate loans and approximately 44% of consumer loans. Consequently, commercial banks play a pivotal role in the transmission of interest rates to the real economy. We summarize the key features 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. In 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 with longer fixed-term periods. In terms of loan rates, the MLR, offered to well-established corporate customers, is typically set at a lower level compared to 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 provides a comparison of the time series of policy rates and interest rates. It becomes evident that longer-term fixed deposit rates have moved into closer alignment with the policy rate. 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. In contrast, the savings rate exhibits a lower correlation of only 0.5. Among loan rates, the MLR shows the highest positive correlation with the policy rate, approximately 0.7, while the MOR and the MRR have correlations around 0.6. This pattern remains consistent when considering the size of commercial banks, as depicted in Figure 3. However, it is noteworthy that small and medium-sized commercial banks tend to adjust 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 that period, even though the MPC significantly increased the policy rate, interest rates at commercial banks for both deposits and lending remained unchanged for approximately half a year. This phenomenon was attributed to the bank balance sheet positions, which had substantial excess liquidity, particularly among the four largest banks (see Waiquamdee and Boonyatotin, 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 becoming 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. The situation of excess liquidity was resolved in late Quarter 2 of 2005, leading commercial banks to subsequently adjust their deposit and lending rates accordingly.

Finally, we detail the macroeconomic variables utilized in this study. 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). Here is a summary of these variables:

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

(b) Inflation is quantified by the annual growth of the HCPI, which serves as the current indicator for IT. The current target range is set at 1%-3%. Over the same sample period, inflation rate averaged 2.09%, with lows of -4.14% and highs of 9.07%. Figure 4 also illustrates a positive correlation between inflation rate and output, aligning with conventional macroeconomic theory. Thai inflation rate remained at a low level since 2015 due to structural issues and low oil prices. However, it experienced a significant uptick due to a sharp drop in growth during the 2020 pandemic. Following the pandemic, the inflation rate enjoyed a remarkable rebound thanks to robust economic growth.

(c) The other two variables serve as control variables used for the NARDL model. Given the Thai economy's heavy reliance on export, we exercise control over the output equation by incorporating the trading partner's 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, which are represented by global energy prices, in assessing the impact of monetary policy on inflation rate.

5 Empirical Results

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

For various models, our presentation plans remain consistent. We initiate with a cointegration test by employing Phillips and Perron's (1988) test on the fitted error-correction term. Rejecting the null hypothesis indicates the presence of a long-run relationship. In the case of the single-step method, we apply the residual bootstrap to compute the standard errors. Additionally, we apply the Wald test to assess long-run and short-run symmetries. Finally, we provide dynamic impulse response functions of the variables of interest by 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. In both models, we estimate the NARDL(2,2) model. Here is a summary of the estimation results:

(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 on 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 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 3-month term-fixed deposit rates is approximately 90.3% for rate reductions, slightly lower than for rate increases, at around 91%. However, this difference is not statistically significant. In contrast, 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. Once again, their difference is statistically insignificant. Of particular interest are the findings concerning the saving rate, which demonstrate a long-run downward asymmetry. In other words, the pass-through for rate increases is only 28.1%, compared to 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 the MLR, the estimated long-run pass-through coefficients are 0.493 for rate cuts and 0.784 for rate hikes, and the test statistic 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 rate increases is approximately 0.775, while for rate decreases, it stands around 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 around 0.896, while for rate cuts, it shows an insignificant coefficient. These results provide evidence supporting 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 ranged only between 21.9% and 70%. This disparity can be attributed to the utilization of more recent data, encompassing the period after 2009, during which the Thai banking system underwent further development. Furthermore, the existing literature was significantly influenced by the study period characterized by the presence of exceptionally high excess liquidity in the banking system. According to the studies by Waiquamdee and Boonyatotin (2008) and Gigineishvili (2011), elevated levels of excess liquidity tend to result in fewer adjustments in interest rates. The sample period of our data spans from 2001 to 2023, encompassing the excess liquidity period of 2007-2008 when banks exhibited a greater inclination to adjust their interest rates. Notably, the Thai commercial bank liquid assets ratio stood at 15.4% in March 2023, in contrast to approximately 25% during 2007-2008.

(e) Figure 5 provides the impulse-response functions of savings and loan rates to policy rate hikes. They are displayed in the left and right panels, respectively. The left panels offer additional insights into short-run asymmetry, indicating a downward asymmetric pass-through for savings rates. In response to policy rate hikes, savings rates exhibit only around a 20% adjustment within 12 months before reaching a long-run rate adjustment. The results also reveal a downward short-run asymmetry in term-fixed deposit rates, with a significant immediate adjustment following policy rate cuts, while the adjustment is less pronounced for rate hikes. It takes around 20 months for the asymmetry to fade away. For loan rates displaced in the right panels, there is no evidence of short-run asymmetry, but it shows upward long-run asymmetry, as indicated by the 90% confidence interval being 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 further controlling for bank size. We present the estimation results in Table 4, along with detailed regression results in Tables 7 and 8. These results are based on NARDL (2,2), except for the large bank saving rate, which is based on NARDL (2,3). We summarize the estimation and inference results as follows:

(a) Overall, the results indicate that loan rates exhibit a greater degree of pass-through for large banks compared to 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 around 35.9% for rate hikes and 46.4% for rate reductions. Conversely, the corresponding pass-through level for large banks is only 18.2% and 0.8% (22.1% and 19.9% from the single-step method), respectively. When it comes to 3-month term-fixed deposit rates, pass-through levels for small and medium-sized banks are at 93.1% for rate increases and 89.8% for rate decreases. In 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 rate cuts, while 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 an adjustment 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 demonstrate a greater capacity to adjust their loan rates in response to policy rate changes compared to small and medium-sized banks.

(c) From an intuitive perspective, these findings indicate that small and medium-sized banks, with lower market power compared to 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 in lending rates to enhance customer appeal, resulting in stickiness in lending rates. In 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 made by Weth (2002) regarding the 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 the macroeconomic impact. We present the estimation and inference results in Tables 9 regarding 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 economic variables than rate decreases. We summarize the key results as follows:

(a) For the impact on output, after controlling for inflation and log trading partner GDP, the results suggest the presence of long-run cointegration among variables, as indicated by the 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 compared to rate reductions. Specifically, an increase in policy rates reduces output growth by 1.903% (0.913% according to the single-step method), whereas a rate reduction stimulates output growth by only 1.556% (0.733% 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-run and short-run, at a significance level of 1%.

(b) When examining the impact on inflation, as presented in Table 9, it becomes evident that policy rate hikes exert a more substantial impact on inflation compared to rate cuts. To elaborate, a 1% increase in the policy rate translates into a decrease in the inflation rate of approximately 0.716% (0.525% according to the single-step method). In contrast, a 1% decrease in the policy rate leads to an increase in inflation of approximately 0.536% (0.45% based on the single-step method). The test statistic from Phillips and Perron (1988) on the fitted error-correction term also confirms 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 both output and inflation is more pronounced compared to rate decreases. This asymmetry is evident in the bootstrap confidence interval, which consistently falls below zero for all horizons.

(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 more 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 the 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.

From these empirical results, we contribute to the literature by emphasizing the asymmetric effects of policy rates on inflation in the long term, specifically in the context of Thailand. The results highlight that a reduction in inflation resulting from an increase in the policy rate is approximately 1.33 times as significant as the inflationary impact caused by a decrease in the policy rate. In terms of output, the upward asymmetric

[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 linkage between the price puzzle and shock identifications is well summarized in Ramey (2016).

impact of a policy rate increase is approximately 1.27 times greater than the impact magnitude of a policy rate decrease.

6 Conclusion

In this study, we empirically analyze the data from Thailand using the NARDL model to investigate 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 interest rate pass-through, we have identified a long-run relationship or cointegration between the policy rate and retail rates. The results strongly indicate that loan rates exhibit incomplete pass-through, showing upward asymmetric pass-through in the long run, with pass-through magnitudes ranging from 42.8% to 89.6%. However, evidence of asymmetric pass-through for savings rates and deposit rates is limited, with magnitudes ranging from 28.1% to 102.7%. Notably, these findings demonstrate higher pass-through rates compared to those reported in the existing literature. This is likely attributed to the utilization of more recent data, including the post-2009 period, characterized by the absence of high excess liquidity.

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

We also examined the macroeconomic implications of policy rate changes on output and inflation. The 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 both output growth and inflation compared to rate reductions. Quantitatively, rate hikes had an effect 1.3 times larger than the impact of rate cuts on both output growth and inflation. These findings suggest that, in the context of Thailand, monetary policy measures aimed at controlling inflation have a more pronounced impact on macroeconomic variables than measures aimed at stimulating economic growth.

Overall, this study contributes to understanding the transmission mechanism and macroeconomic impact of monetary policy in Thailand. The strong evidence of 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)	2.02	1.07	0.50	5.00	1.31	270
RP^+	5.65	2.62	0.00	9.29	4.79	270
RP^-	-5.13	3.00	-8.79	0.00	-7.54	270
ΔRP^+	0.03	0.09	0.00	0.76	0.00	270
ΔRP^-	-0.03	0.10	-0.90	0.00	0.00	270
Commercial Banks						
Saving Rate	0.91	0.47	0.32	2.44	0.57	270
Time Deposits Rate (3 Months)	1.56	0.86	0.46	4.07	1.04	270
Time Deposits Rate (12 Months)	1.90	0.94	0.61	4.51	1.17	270
MOR	7.31	0.59	6.21	8.62	6.68	270
MLR	6.86	0.57	5.91	8.15	6.33	270
MRR	7.69	0.73	6.39	8.71	7.04	270
Large-sized Commercial Banks						
Saving Rate	0.72	0.43	0.25	2.45	0.50	270
Time Deposits Rate (3 Months)	1.41	0.84	0.33	3.94	0.92	270
Time Deposits Rate (12 Months)	1.77	0.97	0.42	4.50	1.02	270
MOR	6.92	0.66	5.87	8.35	6.14	270
MLR	6.46	0.66	5.36	7.90	5.86	270
MRR	7.19	0.74	6.00	8.35	6.41	270
Small and Medium-sized Commercial Banks						
Saving Rate	0.99	0.50	0.36	2.44	0.63	270
Time Deposits Rate (3 Months)	1.63	0.87	0.54	4.12	1.09	270
Time Deposits Rate (12 Months)	1.95	0.92	0.72	4.52	1.24	270
MOR	7.48	0.56	6.38	8.78	6.89	270
MLR	7.05	0.54	6.03	8.31	6.55	270
MRR	7.92	0.72	6.56	8.94	7.27	270
Economic Variables						
Real Output Growth (%YoY)	3.32	3.87	-13.24	21.24	2.11	270
Headline Inflation Rate (%YoY)	2.09	2.17	-4.14	9.27	0.66	270
Log of Trading Partner GDP	4.60	0.01	4.53	4.62	4.60	270
Log of Global Energy Price Index	4.37	0.45	3.30	5.14	4.09	270

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

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***	0.0295	0.354***	0.0802	0.910***	0.0428	0.859***	0.0923	1.027***	0.0617	0.893***	0.106
RP^-	0.404***	0.0413	0.400***	0.0708	0.903***	0.0600	0.866***	0.0830	0.914***	0.0865	0.878***	0.102
Constant	0.372***	0.132	-		-0.254	0.192	-		0.0733	0.277	-	
R^2	0.674				0.760				0.666			
$Adj.R^2$	0.670				0.757				0.662			
LR Asym.												
$\chi^2_{LR\_ASYM}$	11.931		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	0.0312***	0.0111	0.0426**	0.0172	0.0673***	0.0210	0.0643**	0.0263	0.0235***	0.00871	0.0699***	0.0268
Y_{t-1}			-0.0449***				-0.0532***				-0.0501***	
RP^+_{t-1}			0.0159***	0.00480			0.0457***	0.0150			0.0448***	0.0164
RP^-_{t-1}			0.0180***	0.00535			0.0461***	0.0153			0.0440***	0.0162
ECM_{t-1}	-0.0305***				-0.0492***				-0.0324***			
ΔY_{t-1}	0.333***		0.267***		0.432***		0.436***		0.414***		0.417***	
ΔY_{t-2}					-0.0572		-0.0528					
ΔRP^+_t	0.0489*	0.0294	0.0618**	0.0312	0.196***	0.0711	0.191**	0.0806	0.308***	0.0812	0.291***	0.0857
ΔRP^+_{t-1}	0.0324	0.0253	0.0371	0.0284	0.106	0.0665	0.102	0.0751	0.0689	0.0893	0.0466	0.0929
ΔRP^+_{t-2}												
ΔRP^-_t	0.245***	0.0471	0.247***	0.0559	0.565***	0.0493	0.561***	0.0643	0.659***	0.0640	0.663***	0.0812
ΔRP^-_{t-1}	-0.0328	0.0408	-0.0200	0.0457	-0.0425	0.0663	-0.0467	0.0728	-0.136*	0.0705	-0.133*	0.0775
ΔRP^-_{t-2}												
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}	49.217		67.989		9.017		9.325		11.412		13.556	

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***	0.0642	0.957***	0.195	0.775***	0.0776	1.072***	0.223	0.896***	0.114	1.261***	0.194
RP^-	0.493***	0.0900	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.												
$\chi^2_{LR\_ASYM}$	14.064		8.635		13.713		10.646		31.075		18.822	
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	-0.0152***	0.00570	0.167**	0.0750	-0.0211***	0.00741	0.152**	0.0645	-0.0286**	0.0114	0.157***	0.0526
Y_{t-1}			-0.0314***				-0.0279***				-0.0281***	
RP^+_{t-1}			0.0301***	0.00863			0.0299***	0.00782			0.0354***	0.00940
RP^-_{t-1}			0.0236***	0.00768			0.0227***	0.00673			0.0273***	0.00823
ECM_{t-1}	-0.0210***				-0.0183***				-0.00594**			
ΔY_{t-1}	0.368***		0.340***		0.350***		0.311***		0.324***		0.234**	
ΔY_{t-2}	-0.0463		-0.0714		-0.0271		-0.0621					
ΔRP^+_t	0.191***	0.0668	0.190***	0.0734	0.169***	0.0598	0.170***	0.0639	0.163**	0.0699	0.152**	0.0627
ΔRP^+_{t-1}	0.130*	0.0732	0.107	0.0843	0.141*	0.0726	0.115	0.0813	0.180**	0.0875	0.124	0.0783
ΔRP^+_{t-2}												
ΔRP^-_t	0.345***	0.0747	0.367***	0.0904	0.328***	0.0756	0.357***	0.0924	0.334***	0.0880	0.380***	0.106
ΔRP^-_{t-1}	-0.0369	0.0575	-0.0188	0.0635	-0.0313	0.0577	-0.00712	0.0649	-0.0389	0.0641	-0.00140	0.0724
ΔRP^-_{t-2}												
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		17.876		22.465		3.244		16.218	

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	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
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^{+}	0.00825	0.0323	0.199*	0.108	0.836***	0.0408	0.787***	0.0898	1.048***	0.0691	0.964***	0.129
RP^{-}	0.182***	0.0453	0.221**	0.0878	0.922***	0.0572	0.814***	0.0778	1.026***	0.0968	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.												
$\chi^2_{LR\_ASYM}$	19.748		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	0.00461	0.00643	0.0174	0.0134	0.0871***	0.0322	0.0776**	0.0318	0.0695***	0.0249	0.0568*	0.0301
Y_{t-1}			-0.0285**				-0.0658***				-0.0516***	
RP^{+}_{t-1}			0.00568*	0.00320			0.0518***	0.0161			0.0497***	0.0183
RP^{-}_{t-1}			0.00631*	0.00329			0.0536***	0.0170			0.0481***	0.0183
ECM_{t-1}	-0.00903				-0.0455***				-0.0484***			
ΔY_{t-1}	0.316***		0.251**		0.332***		0.322***		0.312***		0.315***	
ΔY_{t-2}					0.0189		0.0124					
ΔRP^{+}_{t}	0.0194	0.0136	0.0282	0.0199	0.216**	0.0910	0.226**	0.0887	0.421***	0.101	0.424***	0.112
ΔRP^{+}_{t-1}	0.0309	0.0205	0.0328	0.0227	0.0457	0.0707	0.0498	0.0715	-0.0193	0.110	-0.00949	0.123
ΔRP^{+}_{t-2}	0.0258	0.0250	0.0270	0.0278								
ΔRP^{-}_{t}	0.162*	0.0917	0.165*	0.0983	0.679***	0.0860	0.671***	0.102	0.796***	0.0890	0.785***	0.108
ΔRP^{-}_{t-1}	-0.0711	0.0524	-0.0613	0.0631	-0.179**	0.0808	-0.173**	0.0864	-0.211**	0.0858	-0.214**	0.0941
ΔRP^{-}_{t-2}	0.0310	0.0450	0.0288	0.0480								
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}	300.811		341.919		3.344		0.718		5.535		6.980	

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***	0.0661	1.061***	0.250	0.789***	0.0880	1.235***	0.284	0.889***	0.103	1.346***	0.286
RP^-	0.566***	0.0926	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		25.070		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	-0.0142**	0.00589	0.140*	0.0748	-0.0212***	0.00713	0.116**	0.0557	-0.0261**	0.0115	0.121**	0.0518
Y_{t-1}			-0.0281**				-0.0240***				-0.0234***	
RP^+_{t-1}			0.0298***	0.00946			0.0296***	0.00827			0.0315***	0.00839
RP^-_{t-1}			0.0248***	0.00911			0.0228***	0.00755			0.0253***	0.00745
ECM_{t-1}	-0.0209***				-0.0209***				-0.00702**			
ΔY_{t-1}	0.257***		0.233***		0.268***		0.226***		0.324***		0.248***	
ΔY_{t-2}	0.00658		-0.0144									
ΔRP^+_t	0.210***	0.0714	0.208***	0.0796	0.194***	0.0675	0.193**	0.0753	0.160**	0.0689	0.152**	0.0698
ΔRP^+_{t-1}	0.162*	0.0875	0.138	0.101	0.167**	0.0827	0.142	0.0955	0.188**	0.0940	0.138	0.0916
ΔRP^+_{t-2}												
ΔRP^-_t	0.483***	0.0827	0.507***	0.101	0.439***	0.0771	0.469***	0.0984	0.448***	0.0763	0.491***	0.0929
ΔRP^-_{t-1}	-0.116*	0.0679	-0.0959	0.0779	-0.0758	0.0627	-0.0540	0.0710	-0.116*	0.0675	-0.0759	0.0737
ΔRP^-_{t-2}												
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		11.085		11.669		4.430		5.565	

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***	0.0311	0.383***	0.0696	0.931***	0.0467	0.891***	0.0978	1.007***	0.0624	0.873***	0.104
RP^-	0.464***	0.0436	0.435***	0.0622	0.898***	0.0654	0.889***	0.0876	0.870***	0.0874	0.860***	0.0986
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	0.0504***	0.0165	0.0641***	0.0229	0.0581***	0.0181	0.0647**	0.0271	0.0161***	0.00610	0.0806***	0.0271
Y_{t-1}			-0.0602***				-0.0527***				-0.0540***	
RP^+_{t-1}			0.0230***	0.00658			0.0469***	0.0155			0.0471***	0.0165
RP^-_{t-1}			0.0262***	0.00747			0.0468***	0.0157			0.0464***	0.0162
ECM_{t-1}	-0.0459***				-0.0468***				-0.0288***			
ΔY_{t-1}	0.303***		0.261**		0.410***		0.409***		0.413***		0.408***	
ΔY_{t-2}					-0.0628		-0.0591					
ΔRP^+_t	0.0558	0.0367	0.0678*	0.0401	0.182***	0.0684	0.174**	0.0781	0.253***	0.0798	0.236***	0.0783
ΔRP^+_{t-1}	0.0239	0.0295	0.0297	0.0365	0.154**	0.0775	0.143	0.0878	0.127	0.0858	0.0990	0.0857
ΔRP^+_{t-2}												
ΔRP^-_t	0.272***	0.0388	0.270***	0.0500	0.526***	0.0438	0.528***	0.0605	0.615***	0.0580	0.623***	0.0776
ΔRP^-_{t-1}	-0.0131	0.0426	-0.00569	0.0465	0.0411	0.0837	0.0406	0.0885	-0.0712	0.0802	-0.0628	0.0831
ΔRP^-_{t-2}												
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		10.397	
χ^2_{Hetero}	0.980		0.002		22.838		22.678		14.632		18.370	

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***	0.0668	0.913***	0.182	0.726***	0.0740	0.986***	0.200	0.864***	0.116	1.180***	0.160
RP^-	0.484***	0.0936	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		12.454		10.985		14.091		29.543		24.288	
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	-0.00843**	0.00407	0.182**	0.0791	-0.0162***	0.00559	0.174***	0.0674	-0.0351***	0.0132	0.202***	0.0664
Y_{t-1}			-0.0335***				-0.0305***				-0.0347***	
RP^+_{t-1}			0.0306***	0.00858			0.0301***	0.00740			0.0410***	0.0113
RP^-_{t-1}			0.0230***	0.00728			0.0224***	0.00616			0.0305***	0.00962
ECM_{t-1}	-0.0294***				-0.0241***				-0.00771***			
ΔY_{t-1}	0.376***		0.356***		0.352***		0.321***		0.255**		0.171	
ΔY_{t-2}	-0.0727		-0.0927		-0.0405		-0.0710					
ΔRP^+_t	0.183***	0.0627	0.184**	0.0754	0.154***	0.0549	0.155**	0.0623	0.173**	0.0792	0.154**	0.0649
ΔRP^+_{t-1}	0.113*	0.0638	0.0966	0.0765	0.125*	0.0635	0.103	0.0724	0.191**	0.0925	0.121*	0.0731
ΔRP^+_{t-2}												
ΔRP^-_t	0.290***	0.0692	0.309***	0.0845	0.287***	0.0750	0.313***	0.0898	0.288***	0.0947	0.339***	0.117
ΔRP^-_{t-1}	0.0193	0.0527	0.0322	0.0593	0.00442	0.0556	0.0231	0.0617	0.0219	0.0681	0.0549	0.0798
ΔRP^-_{t-2}												
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		21.285		24.444		0.406		23.003	

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^+	-1.903***	0.422	-0.913***		-0.716***	0.209	-0.525***	0.194
RP^-	1.556**	0.731	0.733*		0.536***	0.199	0.450***	0.166
Constant	-1,476***	163.2	-		-15.45***	2.796	-	
R^2	0.530				0.558			
$Adj.R^2$	0.517				0.546			
LR Asym.								
$\chi^2_{LR\_ASYM}$	20.557		9.093		36.538		40.425	
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			

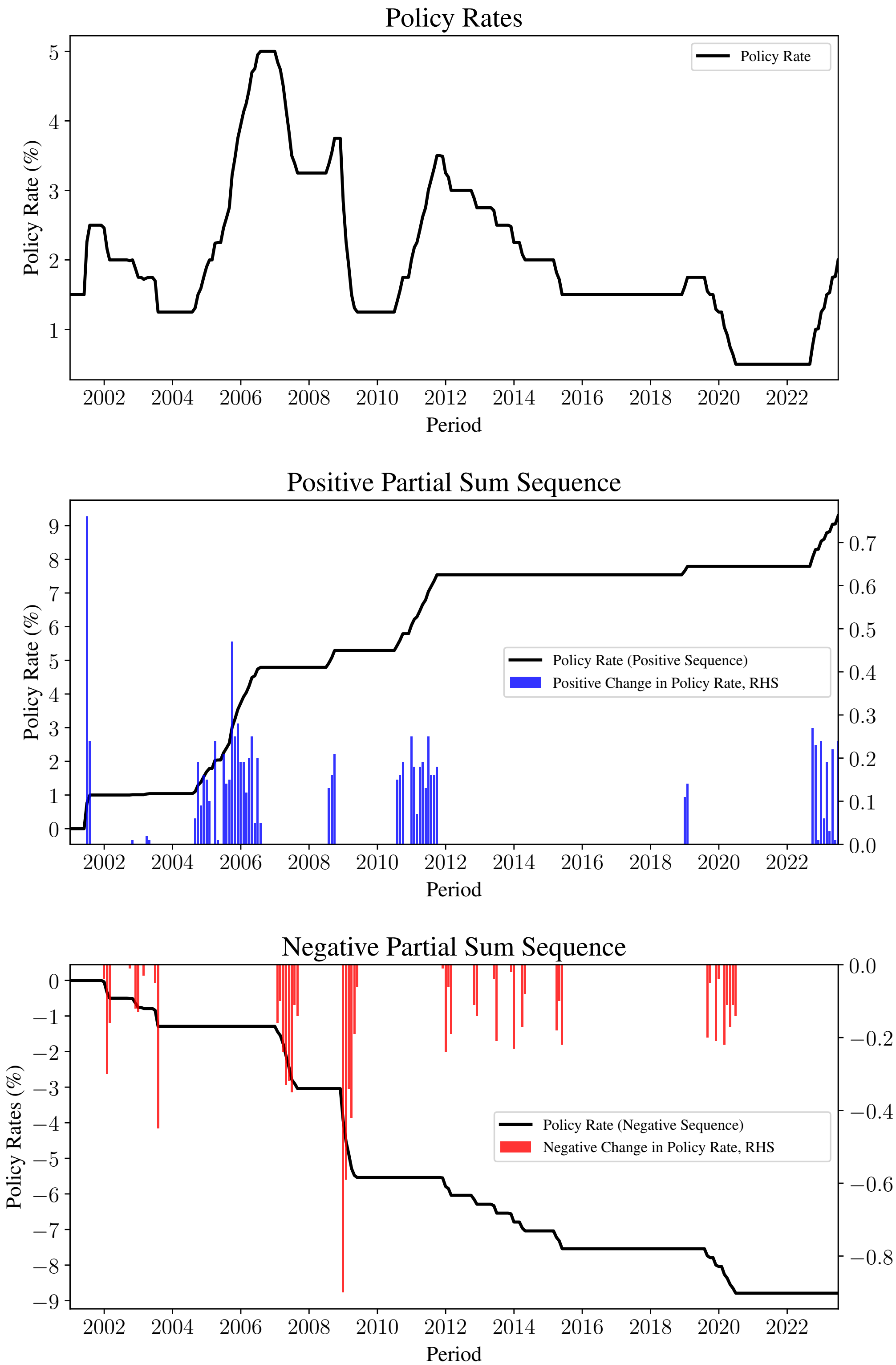
Note: (1) The single-step long-run (LR) parameters are determined as $\hat{\beta}_T^+ := -\hat{\theta}_T^+/\hat{\rho}_T$ and $\hat{\beta}_T^- := -\hat{\theta}_T^-/\hat{\rho}_T$. For the two-step NARDL, the FM-OLS estimator are employed in the first-step, and the parameters are computed as $\ddot{\beta}_T^+ := \ddot{\lambda}_T + \ddot{\eta}_T$ 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.**
Intercept	0.208	0.133	4.284***	0.911	0.0229	0.0432	1.219***	0.222
Y_{t-1}			-0.650***	0.118			-0.416***	0.0795
RP^{+}_{t-1}			-0.593**	0.258			-0.218***	0.0796
RP^{-}_{t-1}			0.476*	0.286			0.187**	0.0811
ECM_{t-1}	-0.341***	0.104			-0.280***	0.0620		
ΔY_{t-1}	0.186	0.118	0.377***	0.120	0.324***	0.0932	0.382***	0.104
ΔY_{t-2}	0.0391	0.104	0.218**	0.0938	0.0335	0.0982	0.126	0.0995
ΔY_{t-3}	-0.00719	0.0803	0.195**	0.0919	0.0151	0.0972	0.0953	0.101
ΔY_{t-4}	-0.102	0.0839	0.0882	0.0847	0.0889	0.0820	0.163*	0.0911
ΔY_{t-5}	0.138*	0.0797	0.292***	0.0881	0.0747	0.0681	0.149*	0.0793
ΔY_{t-6}	0.0835	0.0702	0.256***	0.0781	0.0648	0.0582	0.142*	0.0728
ΔY_{t-7}	-0.0582	0.0694	0.137*	0.0708	0.0377	0.0568	0.105	0.0662
ΔY_{t-8}	0.0628	0.0729	0.219***	0.0829	-0.00329	0.0771	0.0588	0.0848
ΔY_{t-9}	0.261***	0.0707	0.408***	0.0848	0.101	0.0752	0.161*	0.0885
ΔY_{t-10}	0.0131	0.0631	0.162**	0.0773	0.152*	0.0861	0.224**	0.0917
ΔY_{t-11}	0.0172	0.0743	0.168**	0.0737	-0.126*	0.0762	-0.0366	0.0927
ΔRP^{+}_{t}	-4.423***	1.359	-2.309	1.598	0.749	0.655	1.313*	0.722
ΔRP^{+}_{t-1}	-1.127	1.685	-0.611	1.843	-0.675	0.729	-0.206	0.748
ΔRP^{-}_{t}	-0.419	1.228	0.747	1.663	1.133***	0.279	1.257***	0.367
ΔRP^{-}_{t-1}	1.205	1.151	2.301*	1.342	-0.398	0.307	-0.282	0.403
$HCPI^{+}_{t-1}$			0.239**	0.100				
$HCPI^{-}_{t-1}$			-0.0525	0.0911				
$ln(TPGDP)^{+}_{t-1}$			0.429**	0.191				
$ln(TPGDP)^{-}_{t-1}$			0.981***	0.298				
CEI^{+}_{t-1}							-0.00824	0.0160
CEI^{-}_{t-1}							-0.00905	0.0156
$ln(ENERGYP)^{+}_{t-1}$							0.0159***	0.00314
$ln(ENERGYP)^{-}_{t-1}$							0.0136***	0.00271
R^2	0.263		0.362		0.314		0.366	
$Adj.R^2$	0.214		0.303		0.269		0.307	
SR Asym.								
$\chi^2_{SR\_ASYM}$	11.946		5.988		0.938		0.016	
P-value	0.001		0.014		0.334		0.901	
$\chi^2_{S.Corr}$	8.920				5.283			
χ^2_{Hetero}	2.617				9.205			

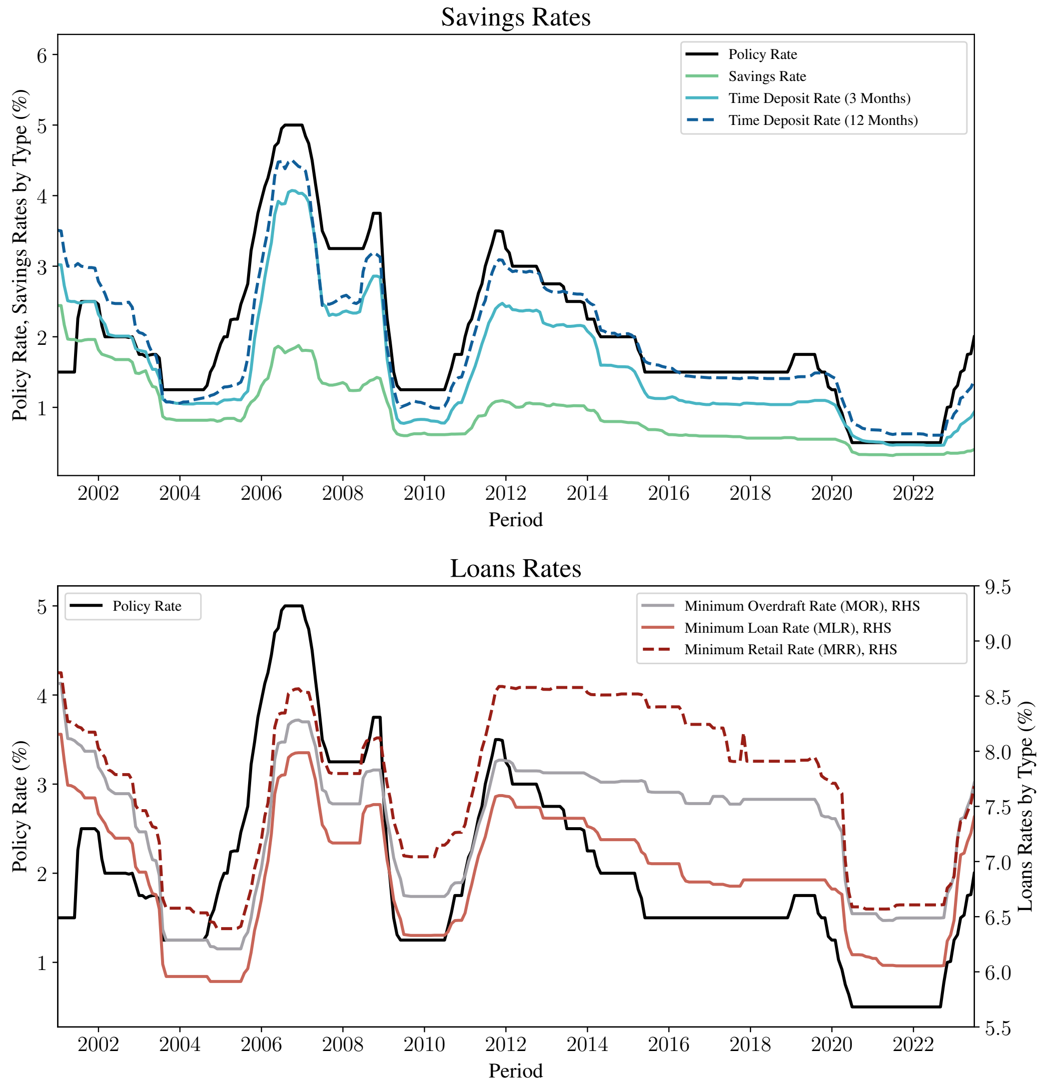
Note: (1) In the single-step method, bootstrap S.E, derived from 5,000 replications are reported. Asymptotic S.E. are reported for the OLS parameters in the two-step model. The significant level: *** p<0.01, ** p<0.05, * p<0.1
(2) $\chi^2_{SR_ASYM}$ denotes the Wald test statistic for SR asymmetry. $\chi^2_{S.Corr}$ 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



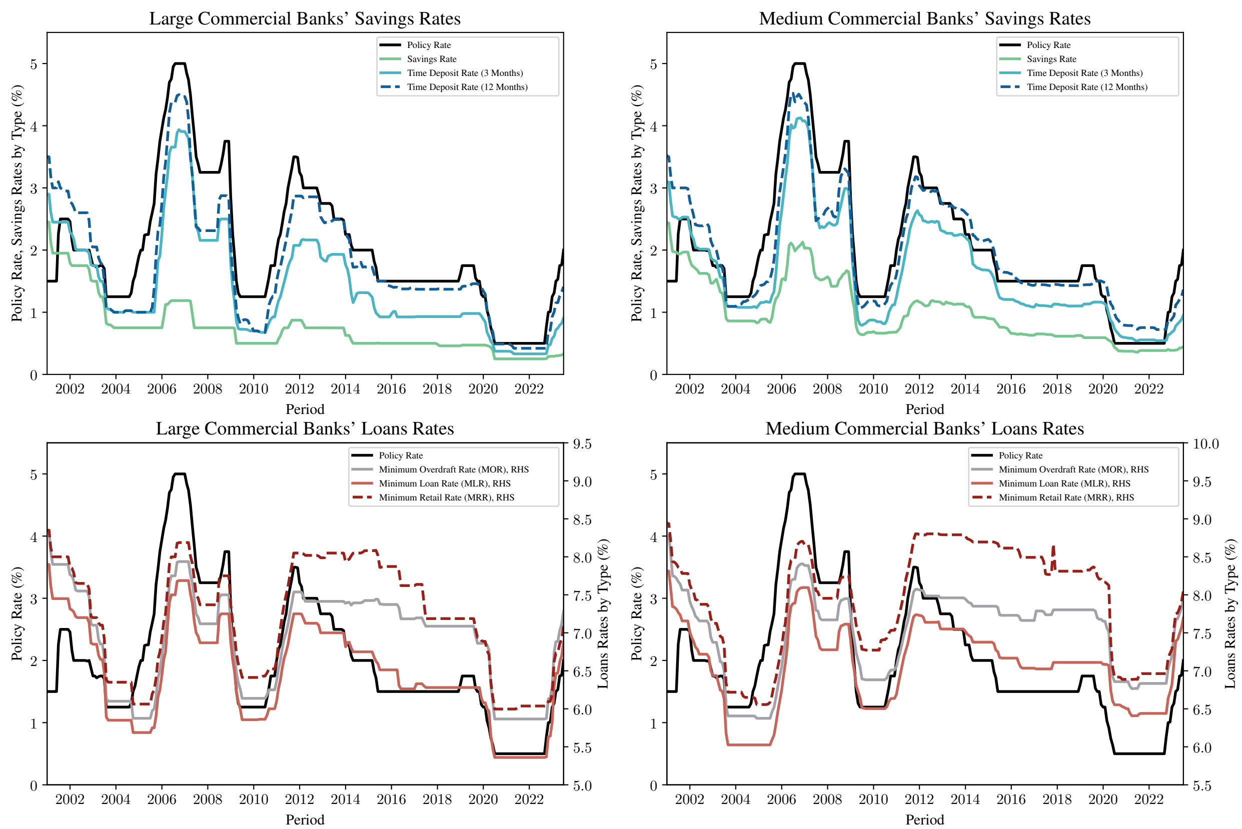
Source: The BOT.
Notes: The BOT has changed its policy interest rate from 14-day RP to 1-day RP since 2007.

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



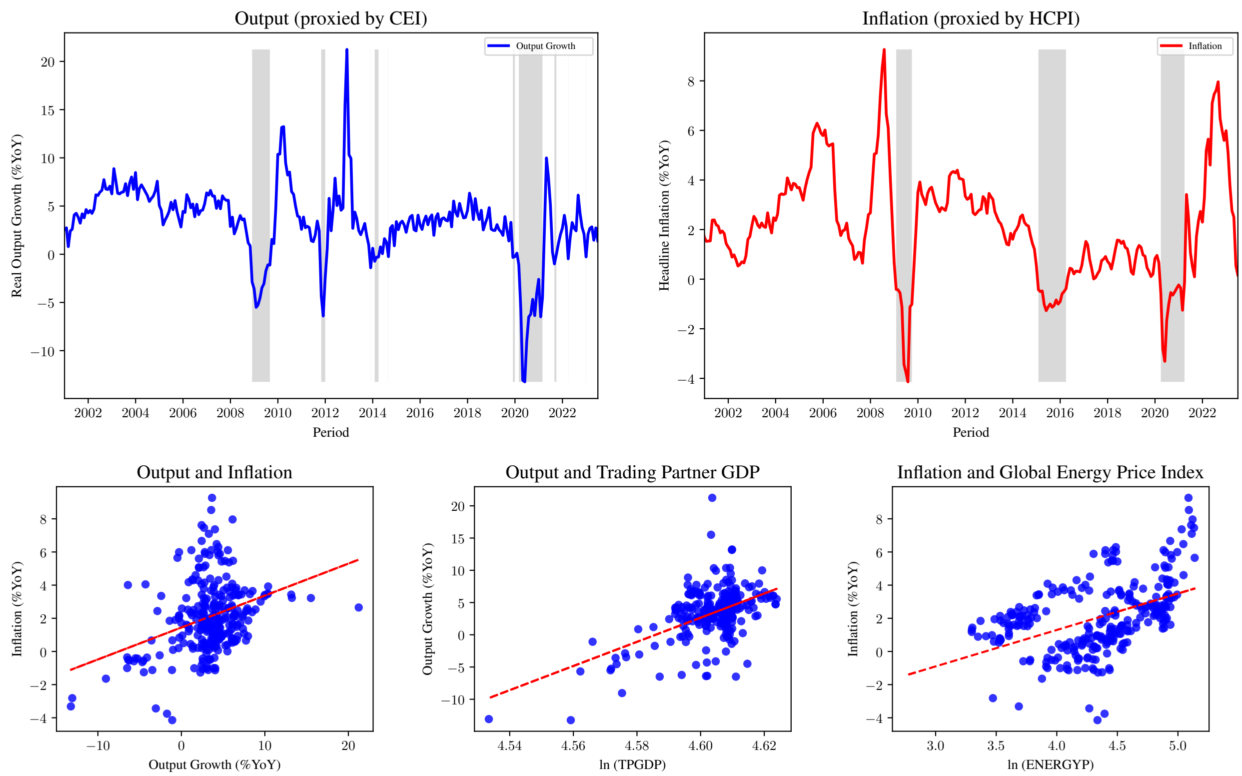
Source: The BOT.
Notes: (1) The BOT has changed its policy interest rate from 14-day RP to 1-day RP since 2007.
(2) Each bank's interest rates are calculated as the average between the minimum and maximum rates reported to the BOT, which is represented as $0.5 \times (min. + max.)$.
(3) For average data,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



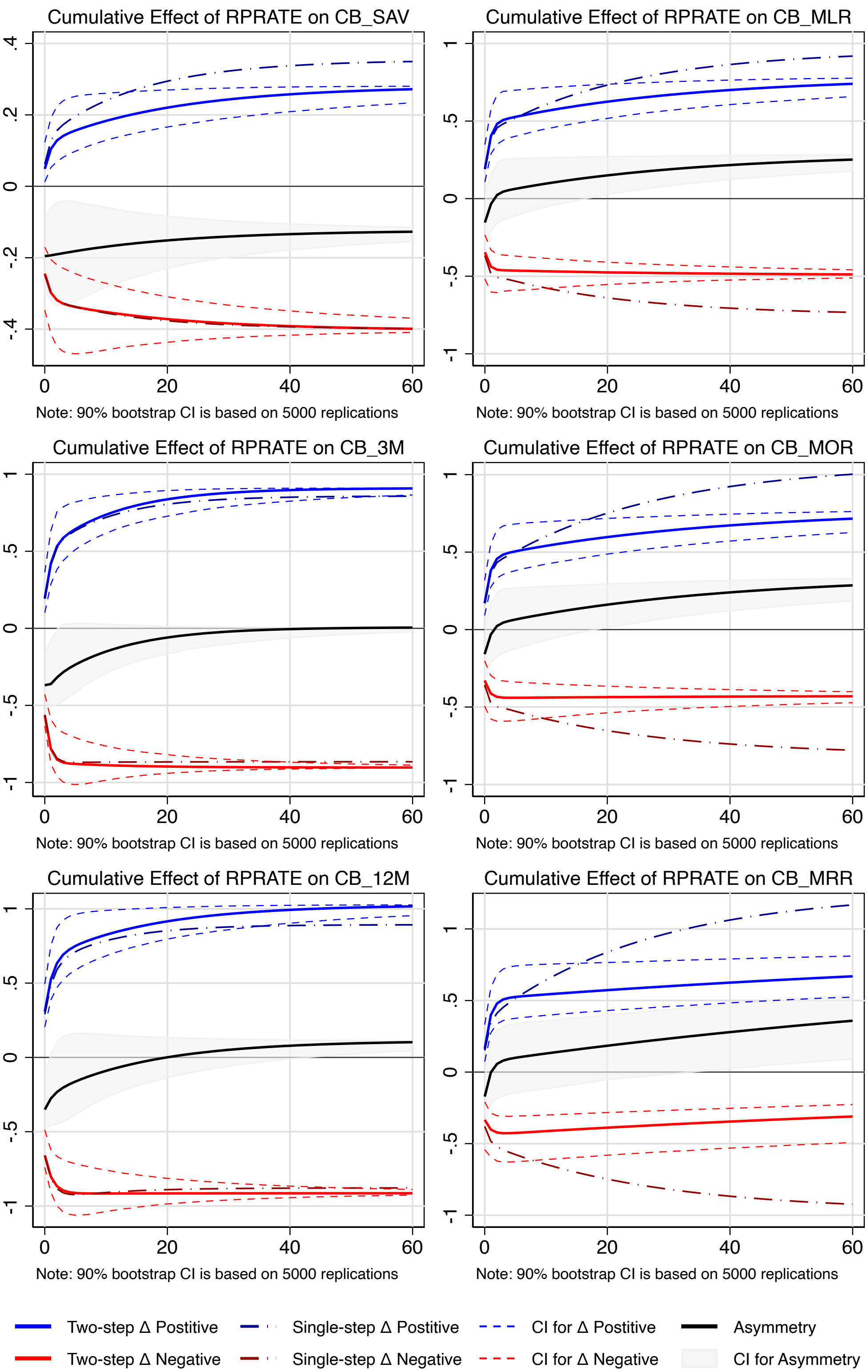
Source: Daily Interest Rates of Commercial Banks Reported on the BOT's Website
Notes: (1) Each bank's interest rates are calculated as the average between the minimum and maximum rates, $rate_i = 0.5 \times (min_i + max_i)$, while the data shown in the graph is the simple average of all banks classified 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



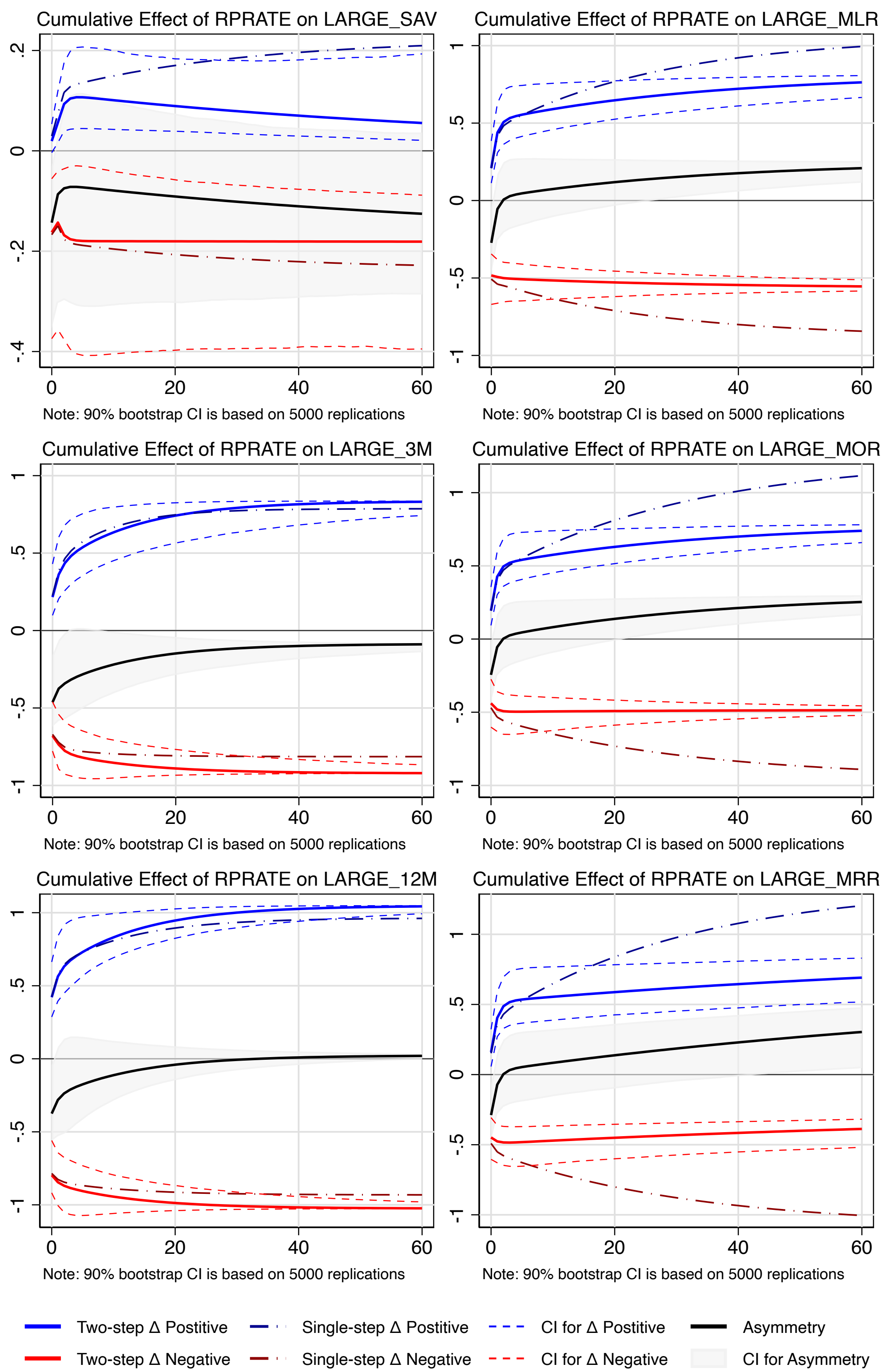
Source: The BOT, Thailand's Ministry of Commerce, FRED data, and OECD dataset.

Figure 5: Dynamic Interest Rate Pass-through



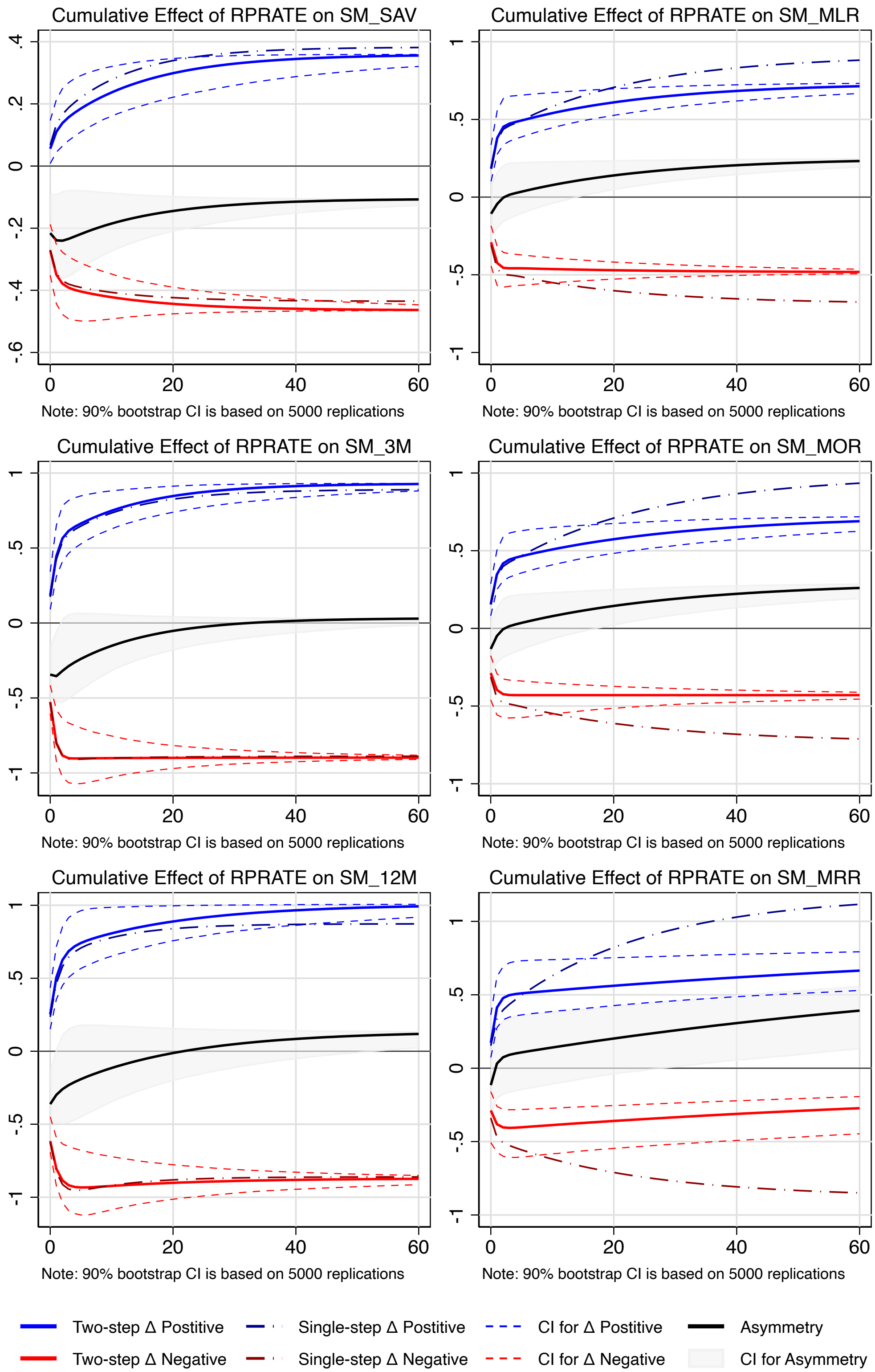
Note: For ease of interpretation, the paper reports the inverse impact results of negative changes. Therefore, it refers to a change in interest rate as -1.

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

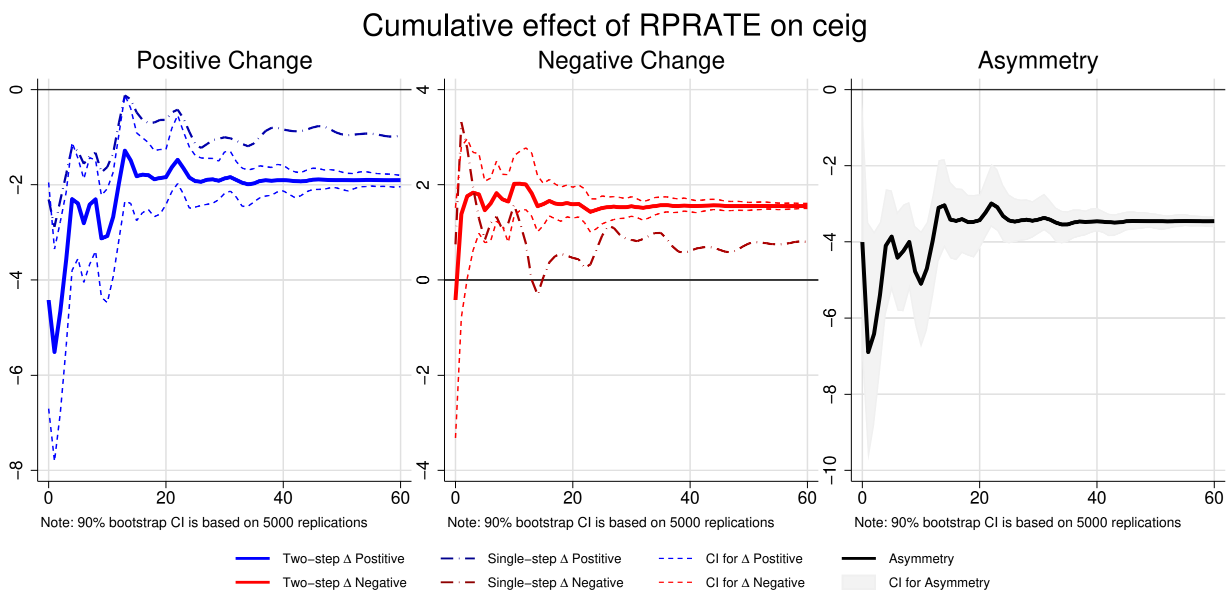
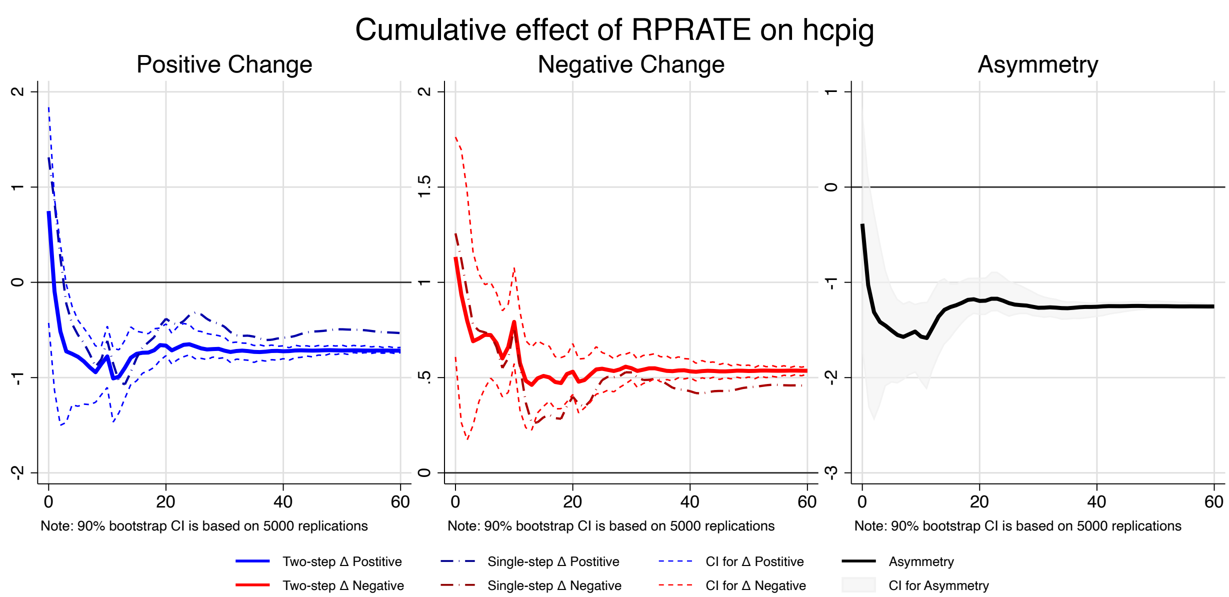


Figure 9: Dynamic Effects of Policy Rate Shocks on Inflation



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 Schwarz Information Criterion (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	-2.9049	4	Unit-root	-5.600	3	Stationary	-5.5832	3	Stationary
CB_SAV	-1.996	2	Unit-root	-2.4827	2	Unit-root	-9.491	1	Stationary	-9.4881	1	Stationary
CB_3M	-2.800	4	Unit-root	-3.2788	4	Unit-root	-5.557	3	Stationary	-5.5528	3	Stationary
CB_12M	-3.028	4	Unit-root	-3.4001	4	Unit-root	-5.298	3	Stationary	-5.2976	3	Stationary
CB_MLR	-3.475	4	Stationary	-3.3872	4	Unit-root	-4.436	3	Stationary	-4.4900	3	Stationary
CB_MOR	-3.307	4	Unit-root	-3.2578	4	Unit-root	-4.569	3	Stationary	-4.6105	3	Stationary
CB_MRR	-2.603	4	Unit-root	-2.5893	4	Unit-root	-5.001	3	Stationary	-5.0051	3	Stationary
$LARGE\_SAV$	-2.990	2	Unit-root	-2.7701	2	Unit-root	-11.370	1	Stationary	-11.5106	1	Stationary
$LARGE\_3M$	-3.078	4	Unit-root	-3.7249	4	Unit-root	-4.955	3	Stationary	-4.9547	3	Stationary
$LARGE\_12M$	-3.381	4	Unit-root	-3.8123	4	Unit-root	-6.433	1	Stationary	-6.4230	1	Stationary
$LARGE\_MLR$	-3.271	4	Unit-root	-3.3147	4	Unit-root	-4.618	3	Stationary	-4.6739	3	Stationary
$LARGE\_MOR$	-3.265	4	Unit-root	-3.1671	4	Unit-root	-7.238	1	Stationary	-7.2689	1	Stationary
$LARGE\_MRR$	-2.234	2	Unit-root	-2.1902	2	Unit-root	-7.202	1	Stationary	-7.2031	1	Stationary
SM_SAV	-2.293	4	Unit-root	-3.1569	4	Unit-root	-5.978	3	Stationary	-5.9816	3	Stationary
SM_3M	-2.660	2	Unit-root	-3.2669	2	Unit-root	-6.160	1	Stationary	-6.1426	1	Stationary
SM_12M	-2.618	2	Unit-root	-3.0002	2	Unit-root	-6.632	1	Stationary	-6.6201	1	Stationary
SM_MLR	-2.448	2	Unit-root	-2.3969	2	Unit-root	-6.891	1	Stationary	-6.9202	1	Stationary
SM_MOR	-3.320	4	Unit-root	-3.3358	4	Unit-root	-7.158	1	Stationary	-7.1645	1	Stationary
SM_MRR	-1.993	2	Unit-root	-2.0253	2	Unit-root	-7.800	1	Stationary	-7.7887	1	Stationary
CEI	-4.904	1	Stationary	-5.4595	1	Stationary	-6.405	12	Stationary	-6.3929	12	Stationary
$HCPI$	-3.716	2	Stationary	-3.8655	2	Unit-root	-6.489	12	Stationary	-6.4490	12	Stationary
$lnTPGDP$	-4.621	5	Stationary	-4.6603	5	Stationary	-5.525	7	Stationary	-5.5251	7	Stationary
$lnENERGYP$	-2.393	2	Unit-root	-2.3254	2	Unit-root	-9.969	1	Stationary	-9.9769	1	Stationary

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