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Feldstein-Horioka Puzzle in Thailand and China: Evidence from the ARDL Bounds Testing

Warawut RUANKHAM[1], Phoommhiphat PONGPRUTTIKUL[2]

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Abstract

This study aimed to investigate the existence of the Feldstein-Horioka (1980) puzzle in international macroeconomics by applying the conditional Autoregressive Distributed Lag (ARDL) model to examine the long-run relationship between national savings and investments in Thailand and China. The input of this study relied on annual national savings and investments as a fraction of GDP during 1980–2019 which was collected from China National Bureau of Statistics (NBS) and Thailand National Economic and Social Development Council (NESDC). Hypothetically, Augmented Dickey-Fuller (ADF) and Phillips-Perron (PP) unit root tests were applied to test the stationary properties and to investigate the integration level of selected time series. The empirical results, confirmed by cumulative sum (CUSUM) and cumulative sum square (CUSUMSQ), maintained no serial correlation and structural break problems. The finding of this study suggested that the Feldstein-Horioka puzzle in Thailand did not exist significantly. Thailand's national savings and investments nexus was independent, following the classic economic idea that financial liberalization, or perfect capital mobility, allowed national savings and investments to flow freely to countries with better interest rates. Whereas, a strong significant correlation was found in the case of China during the fixed exchange rate regime switching in 1994 and post WTO participation after 2001–2019.

Keywords: Feldstein Horioka Puzzle, ARDL Bounds Testing, Cointegration, Saving, Investment

JEL Classification Code: B23, E44, F21

1. Introduction

In the standard economic literature, national savings and investments played an important role to create a higher level of economic growth and development through the capital formation and accumulation process. For example, Keynes' general theory of employment, interest rates, and money and Capitalist Economics (1933–1970) both demonstrated the importance of liquidity in economic expansion. The injection of capital into the economy was considered as the key resource to increase investment and output production (Asimakopulos, 1983). The same could be said of International Macroeconomics, which believed that finance and trade deregulation could boost domestic and global economic growth through foreign direct and portfolio investment (Mohamed Mustafa, 2019; Tung & Thang, 2020). Therefore, domestic capital accumulation and capital mobility became the main source of economic expansion.

Obstfeld and Rogoff (2000) raised the issue of six major puzzles to challenge the international macroeconomic phenomena. They questioned the six puzzles that were unanswered in some industrial countries. For example, the home bias puzzle in trading between the individual in the same country and foreigners, the pricing puzzles, the exchange rate disconnect puzzle, home bias in equity portfolios, and consumption correlation. Besides the puzzles in trading, Obstfeld and Rogoff also pointed out the puzzle in international capital mobility, called by the name of two scholars Feldstein-Horioka (F-H). The F-H is a puzzle because they assumed the dependent relationship

[1]First Author and Corresponding Author. Office of Border Economy and Logistic Study (OBELS), School of Management, Mae Fah Luang University, Thailand [Postal Address: 333 M.1, Tasud, Muang Chiag Rai, Chiang Rai, 57100, Thailand]
Email: Warawut.rua@mfu.ac.th
[2]Research and Development Specialist, Research Development and Innovation (RDI), Dhurakij Pundit University, Bangkok, Thailand. Email: phoommhiphat.min@dpu.ac.th

between national savings and investments that contradicts the standard international macroeconomic. Hence, we should expect perfect capital mobility as a result of trade and investment liberalization, in which individual investment and savings are allowed to be invested in any country with a higher marginal product of investment until the returns of the two countries are equal. When interest rates are arbitrarily set, savers in the United States would invest in foreign countries to get a larger return. This suggested, theoretically, that national investments and savings should have been correlated at a low level, or independent, depending on the extent of financial liberalization and arbitrary interest rate (Obstfeld & Rogoff, 2000).

However, empirically, evidence shows the existence of the contradiction between standard international macroeconomics and real-world situation. There are numerous empirical proofs that the relationship between economic variables may not follow standard economic theory. This phenomenon was widely known as a "puzzle". Obstfeld and Rogoff (2000) found that there are still "many puzzles" existing in economic literature, particularly in international macroeconomics and finance. For instance, the studies of Feldstein and Horioka (1980) initially found a strong correlation between domestic investment and saving (S-I) in some industrial countries. This puzzle is become well known as the Feldstein-Horioka puzzle. They reasoned that if the saving and investment (S-I) assumption is correct and capital mobility is perfect, we should examine the correlation between domestic investment and savings. The results from their study found a strong correlation between saving and investment in some developed countries during 1960–74. (i.e., Australia, UK, Germany).

Feldstein and Horioka (1980) described five rationales behind this correlation of S-I. First, it is because the domestic tax regime discourages investors to invest abroad. This tax regime discounts the return of investment and makes national savings and investment significantly correlated. If the investors invest their savings in other countries, there will be no correlation between S-I. Second, the cost of the transaction, fees, custody, or dealing costs in different countries are volatile. This discourages investors to invest outside and instead invest in their home country, where these costs are calculable and lower. As a result, domestic savings are retained within the country, and we end up with an S-I correlation. Third, there are some exchange rate risks when investing in foreign countries, particularly when withdrawing money back to one's home country. This fluctuation in the exchange rate deters risk-averse investors from investing abroad. Following that, information asymmetry issues may occur outside of the country, causing domestic investors to incur additional costs of information seeking and forcing them to invest only within the country. Finally, foreign investment may be subject to unanticipated regulatory risks imposed by a foreign government. Five reasons as mentioned provide insights behind the international flow of capital during the 1980s (Byrne et al., 2009).

The interpretation of Feldstein and Horioka (1980) inspired the later scholars to study the nexus between saving and investment (S-I) and made the Feldstein and Horioka puzzle receive more interest later on. This puzzle has been widely tested by a number of studies. They found that capital immobility happened due to the country's capital constraints and some restrictions imposed on the outward investment promotion.

Unfortunately, most studies were aimed at advancing economies and industrial countries. Only a small number of studies explained this puzzle in less developing countries like Asia, Africa, and developing Latin American (Narayan, 2005). Nevertheless, from 2005 to 2010, there were some empirical studies (Jiranyakul & Brahmasrene, 2008; Narayan, 2005) that tested the validity of such a theoretical puzzle in less developing countries but not in two fast-growing economies like Thailand and China. Thailand as the home country for Tom-Yum-Kung or the Asian financial crisis in 1997, which changed their exchange rate regime from pegged to a managed float system, needs to be tested whether their financial market has fully operated globally. China on the other hand, as one of the powerful investor of the world, mainly after the exchange-rate regime switching in 1998 and WTO participation in 2001, also need to be studied. Because, if the massive investment and saving in China outflow to other countries, this magnificent amount will structurally change the landscape of investment of the whole region, particularly, countries along the one-belt-one-road routes where their S-I shared a certain level of correlation with China.

This study, therefore, aimed to provide additional empirical evidence to test the existence of the F-H puzzle in Thailand and China during the period of structural change. Most studies employed the classical model (Granger cointegration), only a few studies employed the modern method. The methodology of this study therefore employed modern time-series econometrics which is ARDL Bounds Testing from the study of Narayan (2005). This approach is considered as a better way to test the cointegration between the two non-stationary variables (Pesaran et al., 2001), particularly data with different levels of integration. The main contributions of this paper would consist of 1) provision of additional empirical evidence on major puzzles in international macroeconomics literature in Asia, and 2) empirical evidence supporting the approach of ARDL Bounds Testing.

2. Data and Methodology

The aim of this study was to test the existence of the F-H puzzle by looking at capital mobility. This objective, in particular, can be achieved by testing the cointegration between national savings and investments (S-I). If a strong correlation exists, it implies that the F-H puzzle exists. That is the national savings and investments share a certain level of correlation if investors in that country are not inclined to invest abroad, or there may have been regulations in place to prevent money outflow. If the S-I correlation is independent, the F-H puzzle does not exist. The S-I can then be invested independently in any country with higher returns. For the data time point, this study collected yearly national savings and investments as a fraction of GDP during 1980–2019. The methodology follows four stages of ARDL-Bound Testing for integration. In the first stage, the unit root test for the order of integration was undertaken. The main point was to diagnose the basic assumption of stationary and to ensure the absence of structural break problem. The second stage was about the ARDL-Bound testing for integration of Thailand and China. In the case of Thailand, the purpose of this study was to extend the empirical findings of Jirayakul and Brahmasrene (2008) and to accomplish a longer period of research from 1990 to 2006 and from 1980–2019, which included the 1997 financial crisis.

As shown in Figure 1, the data depicted the movement of national savings and investments during the previous study in 1990–2006 and the extension 1980–2019. This study covered three significant periods, namely, pre-crisis, tipping point (structural break of investment during the financial crisis in 1997), and post-crisis.

While in the case of China, this study extended data coverage of Narayan (2005) from 1952–1998 and 1980–2019. The main finding will provide more evidence to the previous studies and to support the existence of the F-H puzzle in China after passing through two magnificent structural changes, which are the exchange rate regime switching and the WTO participation. Figure 2 depicted the movement of savings and investments (S-I) as a fraction of real GDP in China during 2 significant periods: 1952–1994 and 1994–2019.

In the third stage, this study aimed to achieve the Error Correction Model (ECM) if cointegration was found. The ECM would help to investigate a short-run equilibrium adjustment in their long-run model. In the final stage, the cumulative sum (CUSUM) and cumulative sum square (CUSUMSQ) are adopted to satisfy the structural break assumption.

Annual national savings and investments in Thailand were retrieved from the National Economic and Social Development Council (NESDC). The authors then calculated the savings and the private investments as the fraction of the real GDP of Thailand following Jiranyakul and Brahmasrene (2008). The authors, however, extended the time period starting from 1980 to 2019 or 39 time periods. In the case of China, the model was formulated using the annual data from 1994–2018, which was retrieved from the China National Bureau of Statistics (NBS). The authors extended the empirical finding from Narayan (2005) whose study was based on the period of fixed exchange rate regime (1952–1994 and 1952–1998), to see if there was a structural change relative to 1952–1994 or after the China WTO 2001 participation. The data for both countries is presented in Table 1.

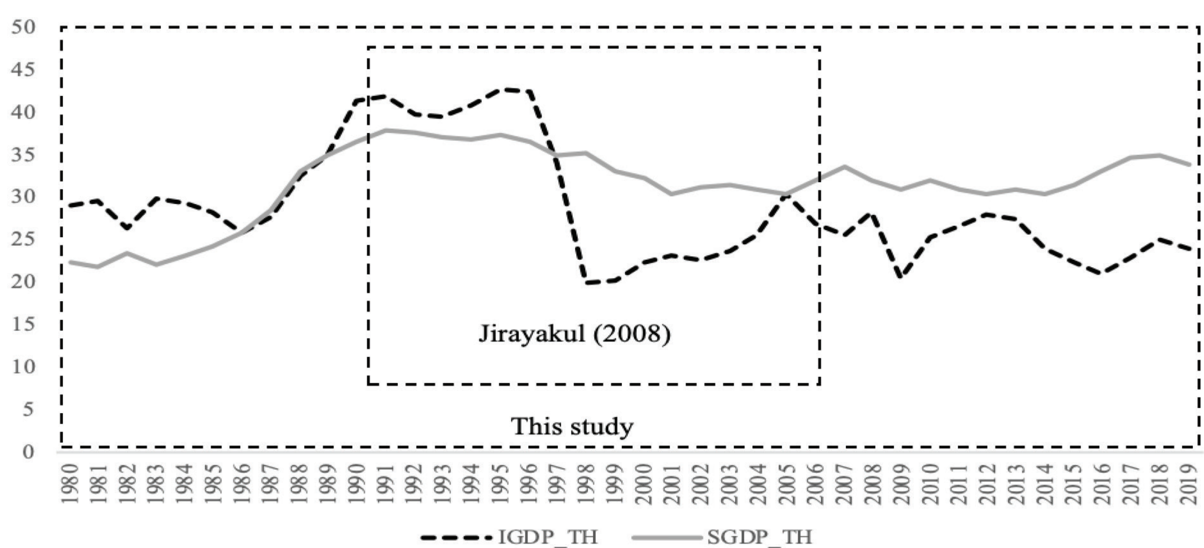

Figure 1: Saving and Investment as a Percentage of Real GDP in Thailand (1980–2019)

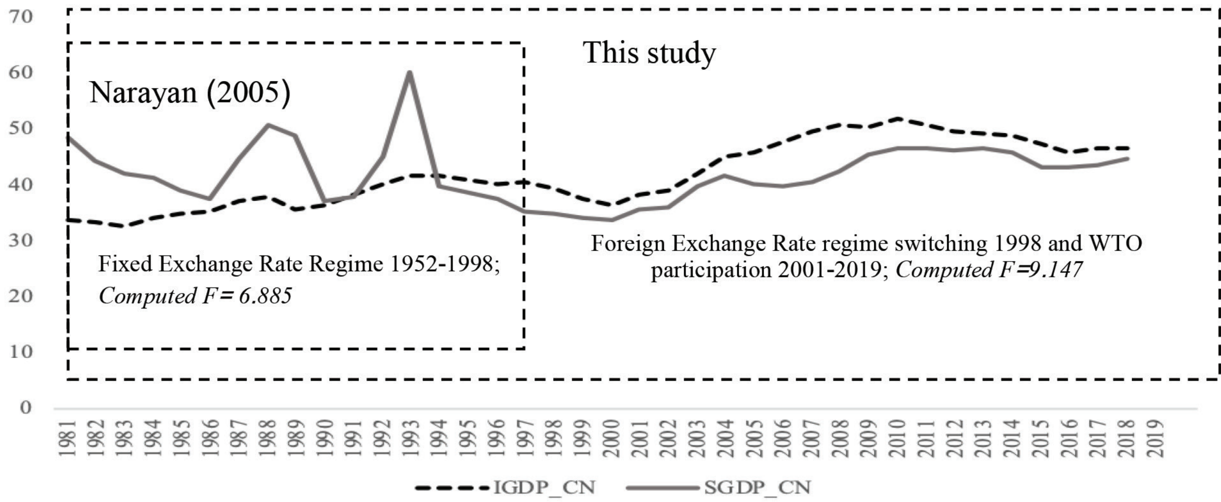

Figure 2: Saving and Investment as a Percentage of Real GDP in China (1980–2019)

Table 1: Selected Annual Investment and Savings as Percent of GDP (1980–2019)

Year	Thailand		China	
	Investment (%)	**Saving (%)**	**Investment (%)**	**Saving (%)**
1980	29.14	22.26	35.33	55.43
:	:	:	:	:
1990	41.54	36.61	36.41	37.25
:	:	:	:	:
2000	22.28	32.29	36.51	33.57
:	:	:	:	:
2010	25.36	32.02	51.71	46.57
2011	26.79	30.90	50.60	46.65
2012	28.02	30.38	49.66	46.25
2013	27.46	31.08	49.35	46.76
2014	23.92	30.49	48.80	45.86
2015	22.36	31.53	47.19	43.03
2016	21.10	33.13	45.96	43.01
2017	22.93	34.77	46.74	43.62
2018	25.19	34.95	46.65	44.74
2019	23.95	33.82		
Average	28.87	31.55	41.77	42.43

Source: Authors calculated from Thailand NESDC and China NBS.

3. Model Construction

Under the modern econometric era, multiples tools for long-run correlation were available and widely applied in various fields of literature (Pesaran et al., 2001; Engle & Granger, 1987). To set up the empirical model, this study followed previous studies (Obstfeld & Rogoff, 2000; Pesaran et al., 2001; Jiranyakul & Brahmasrene, 2008), which was explained by the following stages. First, we indicated the nexus between S-I in regard to Feldstein-Horioka regression, the relation of S-I is expressed as

$$\left(\frac{I}{Y}\right)_t = a_0 + a_1\left(\frac{\text{NS}}{Y}\right)_t + \varepsilon_t \quad (1)$$

Engle and Granger (1987) strictly required that all variables must be integrated of order one or $I(1)$ and must be known before performing the cointegration test. This study thus assumed a null hypothesis that national private investment (I) would be contingent on domestic saving (S) if savings and investment were integrated of the similar order I(1). Therefore, to trace the long-run relationship, we first tested the order of integration under the Augmented Dickey-Fuller (ADF) and Phillips-Perron (PP) and then estimated the ordinary least square (OLS) of two variables in equation (1) where Y represents the real GDP, I represent the investments, and NS represents the national savings.

However, if the same integrated order of stationary by Engle and Granger failed to achieve, we should proceed to a more relaxed approach. Pesaran et al. (2001) then proposed a new method that required a more relaxed assumption of the integration. This approach was called ARDL bounds testing procedure. This approach is mainly based on conditional autoregressive distributed lag (ARDL) and error correction mechanism (ECM). The ARDL cointegration technique is used in determining the long-run relationship between series with different order of integration (Narayan, 2005; Jiranyakul & Brahmasrene, 2008; Saleem & Sial, 2015). That is, ARDL could examine cointegration relationships amongst variables regardless of their order of integration, I(0), I(1), or a mixture of both.

In the second stage, we indicated the model specification based on the ARDL instead of Engle and Granger (1987). The model was specified as:

$$\Delta\left(\frac{I}{Y}\right)_t = \alpha_0 + \sum_{i=1}^{p} \beta_t \Delta\left(\frac{I}{Y}\right)_{t-1} + \sum_{i=1}^{q} \gamma_t \Delta\left(\frac{\text{NS}}{Y}\right)_{t-j} + \lambda\Delta\left(\frac{\text{NS}}{Y}\right)_t + \varepsilon_t \quad (2)$$

Where the expression (3) and (4) provided an auto-regressive part (lag of endogenous variables) and the distributed lag (lags of exogenous variables) by adding the lagged level variables into equation (2). The α, β, γ are the parameters of short-run variables and λ was the parameter of the long-run variables. The study then obtained the F-statistic by estimating equations (3) and (4) for both countries as following equations;

Thailand:

$$\begin{aligned}\Delta(I/Y)\_\text{TH}_t = {} & \alpha_{01} + \alpha_{11}(I/Y)\_\text{TH}_{t-1} \\ & + \alpha_{12}(\text{NS}/Y)\_\text{TH}_{t-1} \\ & + \sum_{i=1}^{p} \beta_t \Delta(I/Y)\_\text{TH}_{t-i} \\ & + \sum_{i=1}^{q} \gamma_j \Delta(\text{NS}/Y)\_\text{TH}_{t-j} \\ & + \lambda\Delta(\text{NS}/Y)\_\text{TH}_t + \varepsilon_t \end{aligned} \quad (3)$$

China:

$$\begin{aligned}\Delta(I/Y)\_\text{CN}_t = {} & \alpha_{02} + \alpha_{21}(I/Y)\_\text{CN}_{t-1} \\ & + \alpha_{22}(\text{NS}/Y)\_\text{CN}_{t-1} \\ & + \sum_{i=1}^{p} \beta_t \Delta(I/Y)\_\text{CN}_{t-i} \\ & + \sum_{i=1}^{q} \gamma_j \Delta(\text{NS}/Y)\_\text{CN}_{t-j} \\ & + \lambda\Delta(\text{NS}/Y)\_\text{CN}_t + \varepsilon_t \end{aligned} \quad (4)$$

Next, we compared the computation of F-statistics from (3) and (4) with the critical bound values obtained from the Narayan (2005) (Case 3: (unrestricted intercepts; no trends) $a_0 \neq 0$ and $a_1 = 0$. According to Pesaran et al. (2001), there are 5 cases provided for testing the cointegrating bound test). This was because it fit better with a small number of observations (Jiranyakul & Brahmasrene, 2008). The hypotheses for integration were presented in Table 2.

Table 2: Hypotheses for Thailand and China

Thailand	
H_0: $\alpha_{01} = \alpha_{11} = 0$	No cointegration
H_1: $\alpha_{01} \neq \alpha_{11} \neq 0$	Cointegration exists
China	
H_0: $\alpha_{02} = \alpha_{21} = 0$	No cointegration
H_1: $\alpha_{02} \neq \alpha_{21} \neq 0$	Cointegration exists

In addition to the ARDL, if cointegration exists, this study could further find the short-run equilibrium adjustment obtained from the error correction model (ECM) by replacing the lagged level variables in equations (3) and (4) with the one-period lagged residuals from the estimate of equation (1).

4. Results

In the first stage, we tested the unit root under the Augmented Dickey-Fuller (ADF) and Phillips-Perron (PP). The results, as shown in Table 3, concluded that S-I as a fraction of real GDP was non-stationary in level with trend and without trend in Thailand. However, both variables were stationary after the first difference or $I(1)$. While in the case of China, the saving ratio was stationary at the level $I(0)$ in all cases. The investment ratio was nonstationary until we took the first difference.

5. ARDL-ECM Cointegration Study

The bounds testing procedure simply estimates OLS regression with the first difference of the variables in equation (2). Then, put the lagged level variables and conduct the additional test. Finally, compute the F-statistics and compare them with the critical values (Narayan 2005) (Case 3: (unrestricted intercepts; no trends)). The result displayed in Table 4 showed the ARDL equation for Thailand and China with the computed F-statistics and χ^2. The estimated result shows the computed F-statistics for Thailand and China was 2.753 and 9.147 respectively. In addition, the LM test was applied to diagnose the serial correlation. The result indicated no serial correlation problem since the p-value of $\chi^2(2)$ was greater than 0.05 for both Thailand and China.

Now, to conclude the cointegration, we compared the computed F-statistics shown in Table 4 with the Bounds critical value in Table 5. In Thailand's case, the study

Table 3: Tests for Unit Root

	ADF Statistic		Phillips-Perron	
	Without Trend	**With Trend**	**Without Trend**	**With Trend**
Thailand				
Savings Ratio (level) (S/Y)	−2.071 [0] (0.257)	−2.235 [1] (0.457)	−2.112 {4} (0.241)	−1.788 {3} (0.691)
Savings Ratio (1st difference) Δ(S/Y)	−3.832*** [0] (0.005)	−4.023** [0] (0.016)	−3.888*** {3} (0.005)	−4.067** {3} (0.014)
Investment Ratio (level) (I/Y)	−1.673 [0] (0.436)	−2.243 [1] (0.453)	−1.673 {0} (0.437)	−2.261 {1} (0.448)
Investment Ratio (1st difference) Δ(I/Y)	−4.981*** [0] (0.000)	−3.906** [0] (0.022)	−4.874*** {5} (0.000)	−4.802*** {5} (0.002)
China				
Savings Ratio (level) (S/Y)	−3.794*** [0] (0.006)	−3.699** [0] (0.035)	−3.832*** {7} (0.005)	−3.685** {8} (0.035)
Savings Ratio (1st difference) Δ(S/Y)	−5.956*** [2] (0.000)	−5.942*** [2] (0.000)	−9.111*** {17} (0.000)	−11.494*** {16} (0.000)
Investment Ratio (level) (I/Y)	−1.623 [1] (0.461)	−2.243 [1] (0.453)	−1.091 {3} (0.710)	−2.009 {3} (0.578)
Investment Ratio (1st difference) Δ(I/Y)	−3.904*** [0] (0.004)	−3.906** [0] (0.022)	−3.882*** {2} (0.005)	−3.868** {2} (0.024)

Note: a) The number in [square brackets] is the optimal lag length determined by Schwartz information criterion (SIC) with a maximum lag of 9. b) The number in {curly brackets} is the optimal bandwidth determined by the Newly-West using Bartlett Kernel. c) The number in (parentheses) is the p-value provided by MacKinnon (1996). d) *, **, and *** denotes 10, 5, and 1 percent significance levels, respectively.

Table 4: Results of ARDL Test for Cointegration

Thailand	**China**
$\Delta I_t = 1.449 - 0.239 I_{t-1} + 0.167 S_t + 0.298 \Delta I_{t-1}$	$\Delta I_t = 4.483 - 0.453 I_{t-1} + 0.388 S_t + 0.381 \Delta I_{t-1}$
(0.303) (−2.346)** (1.010) (1.815)**	(2.143) ** (−4.171) *** (3.384) *** (2.211)**
Computed ***F* = 2.753,** $\chi^2(2)$ = 0.593 (p = 0.7433)	Computed ***F* = 9.147 ,** $\chi^2(2)$ = 0.109 (p = 0.947)

Note: a. The number in parenthesis is t-statistics. b. *, **, and *** denotes 10, 5, and 1 percent significance levels, respectively.
c. LM $\chi^2(2)$ is statistics for testing no residual serial correlation against orders 2.

concluded that the *F*-statistics was below the lower bound for both studies (Pesaran et al., 2001; Narayan, 2005). Thus, the cointegration between national savings and investment did not exist. The result of Thailand matched with Jiranyakul and Brahmasrene (2008). Whereas, China exhibited a strong and significant cointegration between national savings and investments. The computed *F*-statistic was 9.147, which was located above the upper bound for both studies (Pesaran et al., 2001; Narayan, 2005). Our result was similar to Narayan (2005) that cointegration existed after the fixed exchange rate regime (1994–2019).

6. The Long Run Estimation Results (Case of China)

Since the cointegration of S-I existed in China, the authors further constructed the long-run model using the estimated equation (5) to ensure the robustness of the model. The estimated long-run coefficients, as well as the robustness test which included the serial correlation, normality, and heteroscedasticity tests, were presented in Table 6.

$$\frac{I}{Y} = \alpha_0 + \sum_{i=0}^{m} \alpha_1 \left(\frac{I}{Y}\right)_{t-i} + \sum_{i=0}^{n} \alpha_2 \left(\frac{\text{NS}}{Y}\right)_{t-1} + \mu_t \qquad (5)$$

Results in Table 6 suggested that the long-run model was robust as it presented no serial correlation and heteroscedasticity problem. The model also indicated strong and significant long-run relationships between S-I during 1994–2019. We concluded the existence of the F-H puzzle in China during a particular period as we significantly found cointegration of S-I evidenced from both ARDL Bound testing and the long-run model.

7. The Estimated Error Correction Model (Case of China)

Table 7 presented the short-run dynamic results between national savings and investments along with the diagnostic test. The error correction (EC_{t-1}) was selected based on adjusted *R*-square corresponding diagnostic statistics. The error correction EC_{t-1} represented the speed of adjustment

Table 5: The Bounds Critical Value

***F*-statistics Lower to Upper Bound**	**Pesaran et al. (2001) Critical Bounds**	***F*-statistics Lower to Upper Bound**	**Narayan, P. K. (2005) Critical Bounds**
6.87 to 7.84	1%	5.20 to 6.84	1%
4.94 to 5.73	5%	3.62 to 4.91	5%
4.04 to 4.78	10%	2.96 to 4.10	10%
Criteria:		**Conclusion:**	
Above the Upper Bound		Cointegrated	
Between the Lower and Upper Bound		Inconclusive	
Below the Lower Bound		No cointegrated	

Note: Critical value obtained from the study of Pesaran, Shin, and Smith (2001) and Narayan (2005), case III Unrestricted intercept and no trend (fit better with 30–80 observations).

Table 6: Estimated Long Run Coefficients (ARDL, 1,1)

Variables	***c***	**$(I/Y)\_CN_{t-1}$**	**$(S/Y)\_CN_{t-2}$**
Coefficient	4.483**	−0.453**	0.388***
t-statistic	2.143	−4.171	3.384
Diagnostic tests			
	Statistic	***p*-value**	
Serial Correlation	0.334	0.967	
Normality (Jarque-Bera)	1.205	0.547	
Heteroscedasticity	0.854	0.533	

Note: *, **, and ***denotes 10, 5, and 1 percent significance level, respectively.

Table 7: Error Correction Model (ECM), Dependent Variable ΔI_CN_t

Variables	c	$\Delta(I/Y)\_CN_{t-1}$	$\Delta(I/Y)\_CN_{t-2}$	$\Delta(I/Y)\_CN_{t-3}$	EC_{t-1}
Coefficient	4.483***	0.382***	0.105	0.471***	−0.453***
t-statistic	4.421	2.284	0.541	2.610	−4.417
Diagnostic Tests					
R^2 adjusted	0.629		SIC	3.056	
F-statistic	9.483		AIC	2.808	
DW-statistic	2.098		RSS	12.657	

Note: a. Cointegration equation EC = IGDP_CN - (0.8573*SGDP_CN). b. *, **, and *** denotes 10, 5, and 1 percent significance level, respectively.

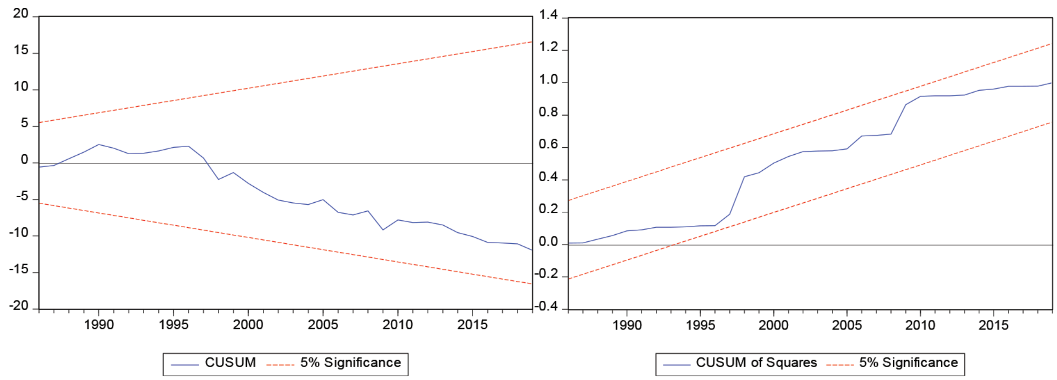

Figure 3: Stability Test for Thailand (TH)

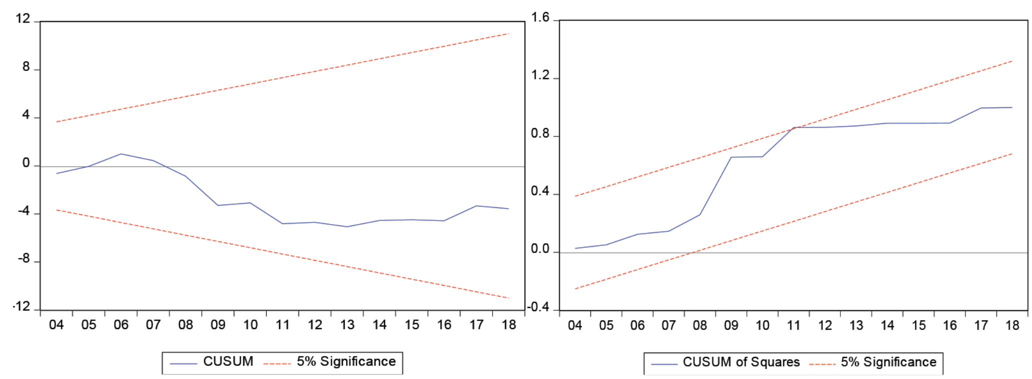

Figure 4: Stability Test for China (CN)

or equilibrium reversion in the dynamic model after any disequilibrium (Pesaran et al., 2001). The ECM suggested that the result was negative and statistically significant with a 95% confidence level, which was similar to the findings of Narayan (2005). This meant that the result confirmed the existence of a long-run relationship of S-I and strong equilibrium reversion in China.

8. Stability Test

In addition to the unit root test, the study plotted the cumulative sum (CUSUM) and cumulative sum square (CUSUMSQ) to ensure the model passed the necessary assumptions of stationary and the structural break. The stability test such as the cumulative sum (CUSUM) and cumulative sum square (CUSUMSQ) tests was used to investigate the stability of the equation (2). There appear to be no structural breaks in the two countries. The results were depicted in Figures 3 and 4.

9. Conclusion

This study aimed to provide more empirical results to international macroeconomic and finance literature by investigating the F-H hypothesis. We carefully tested the existence of cointegration between national savings and investments (S-I) in Thailand and China by ARDL-Bounds testing procedure. Annual national savings and investments data was retrieved from the National Economic and Social Development Board (NESDB) of Thailand during 1980–2019 and the National Bureau of Statistics of China during 1994–2019. The time series data diagnosis passed all basic assumptions of stationary claimed by the ADF and PP unit root test. The results from CUSUM also suggested that there was no structural break found over the whole period of study (1980–2019). To conclude the cointegration results, this study borrowed the Bound Critical Value from Narayan (2005), which was suitable for a smaller number of observations compared to Pesaran et al. (2001). As a result of the computed *F*-statistic (2.753) not satisfying the necessary bound value, this study showed no long-run association between S-I in Thailand. The cointegration results of Thailand supported the empirical results from Jirayakul and Brahmasrene (2008) and Byrne and Fiess (2009). This is due to the difference in the level of financial market development and economic growth (Tariq et al., 2020) and interest rate arbitrage in Asian countries which causes capital movement in Asian countries. Whereas, this study found a strong correlation in the case of China during the fixed exchange rate regime switching in 1994 and post WTO participation in 2001. The computed *F*-statistic (9.147) strongly indicated the long-run relationship. This capital immobility happens due to the country's capital constraints and some restrictions imposed on the outward investment promotion, as well as the home bias in investment among Chinese investors. In addition, the long-run equation and the Error Correction Model (ECM) were estimated to ensure the existence of a long-run relationship and short-run equilibrium adjustment respectively. We strongly found evidence of a long-run relationship of S-I which implied that the F-H puzzle exists in China. In the same word, we found strong nexus and capital immobility after China's exchange rate regime switching and WTO participation. Our results were in line and provide an extended finding of Narayan (2005) that a long-run relationship of S-I was found during 1994–2019 regardless of the structural break during exchange rate regime switching in 1998 and post-WTO participation in 2001.

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