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Thai Inflation Dynamics in a Globalized Economy

by

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Abstract: This paper investigates whether the observed changes in Thai inflation dynamics since the 1990s can be attributed to the process of globalization. First, this paper develops a dynamic factor model to extract a global component from underlying inflation rate movements in Thailand and its top trading partners. Based on the empirical findings, the importance of the global factor for Thailand doubled since 2001, emphasizing the growing role of globalization since then. Second, to explore the economic determinants behind the global factor, this paper estimates an unobserved components model for Thai inflation that is consistent with an Open Economy New Keynesian Phillips curve (OE-NKPC). The empirical model incorporates structural breaks to examine how the influences of domestic and global output gaps for Thai inflation changes over time. Based on the findings, long-term inflation expectations declined significantly and became well anchored at an average level of 2.4 percent shortly after the Bank of Thailand adopted an explicit inflation target in 2000. At the same time, short-run inflation movements became increasingly driven by a global rather domestic output gap. Based on an extended OE-NKPC, the global output gap still remains important beyond the direct import price channel during the 2001-2007 period. However, after the global financial crisis, the global output gap only serves to capture the direct effects of world oil price movements on inflation.

Keywords: Inflation; Dynamic Factor Model; Globalization; New Keynesian Phillips Curve; Output Gap; Structural Break; Unobserved Components Model.

JEL Classifications: E3, E5, F4.

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1. Introduction

Inflation dynamics in Thailand has undergone fundamental changes during recent decades. In particular, Thai inflation has become remarkably low and stable since the early 2000s. Furthermore, there has been a marked decline in the degree of inflation rate persistence during recent periods, implying that a temporary shock to the price level dissipates rather quickly (Chantanahom et al., 2004; Khemangkorn et al., 2008). Given that inflation is ultimately a monetary phenomenon, many studies often attribute the improved behavior of Thai inflation to the inflation targeting framework that was adopted in May 2000. Monetary policy however, primarily works through lowering and anchoring long-run inflation trends. Therefore, the monetary policy explanation may not be able to give a full account of changes in Thai inflation dynamics, especially for those that have occurred over the short to medium-run.

This paper investigates whether globalization plays an important role in explaining Thai inflation rate movements since the early 1990s. The focus on the role of globalization stems from the observation that low and stable inflation has become a salient feature of worldwide economies as of late, particularly in advanced countries. According to a number of studies, globalization, defined as the integration of goods, factors, and financial markets, has been suggested to help mute inflationary pressures around the world through a series of favorable external shocks. For example, the integration of low cost economies such as China and India into world trade systems that occurred in the early 2000s have been suggested to help hold down domestic inflation by depressing trade prices and increasing the share of imports in domestic demand (IMF 2006; Kohn, 2006). Greater integration of markets has been suggested to enhance the degree of international competition, which helps restrain markups and producer prices, ultimately lowering inflation (Neiss, 2001; Binici et al. 2012). Given that Thailand is an open economy with trade to GDP levels currently exceeding 100 percent, globalization should have an important bearing for Thai price processes.

One way to quantify the impact of globalization is to extract common movements underlying a series of inflation rates via a factor analysis, to produce a so-called 'global factor' (Ciccarelli and Mojon, 2010; Neely and Rapach, 2011, Manopimoke, 2015). The importance of this global factor is often associated with movements in the global output gap, which measures the amount of economic slack or resource utilization at the global level. Despite the intuitive appeal of the global output gap, the evidence on its link with national inflation rates is mixed. Based on estimated Phillips curve models that are augmented with a global gap variable, Borio and Filardo (2007) finds strong evidence that the global output gap helps explain inflation dynamics in a number of OECD economies. Ihrig et al. (2010), on the other hand, argue that these results are not robust to plausible alternative specifications of the

Phillips curve. Conflicting evidences are also present in earlier work (Ball, 2006; Pain et al., 2008).

However, a main shortcoming of existing studies is their limited focus on advanced economies. With export volumes in Asian countries currently accounting for more than one-third of world trade flows, studying the impact of globalization for an emerging county such as Thailand can be vital towards gaining a more comprehensive analysis of the issue at hand. This paper attempts to fill this gap in the literature by investigating the impact of globalization for Thai inflation dynamics. The study is divided into three main parts. First, a dynamic factor model (DFM) is estimated to study whether there is a global factor driving Thai inflation rate movements, and whether the importance of this global factor has increased over time. Second, to study whether this global factor is related to the global output gap, this paper builds upon the approach of Kim et al. (2014) and develops an unobserved components (UC) model for inflation that is consistent with an open-economy New Keynesian Phillips curve (OE-NKPC). The empirical model incorporates structural breaks to allow the influences of the domestic and global output gaps to vary over time. Last, to identify the channels through which the global output gap is important for Thai inflation, the UC model is augmented to account for the role of external factors, such as the changes in global commodity prices and real exchange rates.

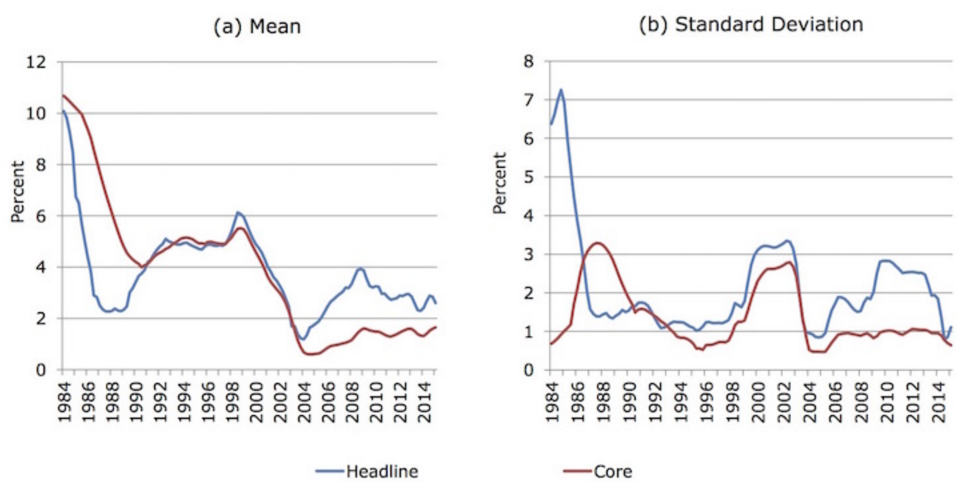
A preview of the main empirical results are as follows. First, the results from the DFM analysis suggest that Thai inflation was increasingly driven by global factors since 2001. A more detailed analysis from the UC OE-NKPC reveals that Thai inflation underwent two distinct structural changes over the 1993-2015 sample. The first regime change in 2001 corresponded to a significant decline in agents' long-term inflation expectations, which occurred shortly after the Bank of Thailand adopted an inflation targeting framework in May 2000. Globalization also appeared to have important repercussions for Thai inflation in the short-run, as evidenced by a growing role for the global output gap at that time, coupled with a decline in the significance of the domestic slack variable. The second structural break occurred after the global financial crisis in 2007, which altered how the global output gap became relevant for Thai inflation dynamics. During 2001-2007, the addition of variables such as import price inflation and oil price movements in the OE-NKPC failed to reduce the role of the global output gap, implying that the importance of the global output gap for Thailand extends beyond the direct import price channel. On the other hand, taking into account world oil price movements in the period thereafter completely removed the statistical significance of the global output gap. This finding implies that after the global financial crisis, the global output gap was only relevant for Thai inflation insofar as it was capturing the effects of world oil price movements.

This paper is organized as follows. Section 2 provides a discussion on how inflation has become more global in nature, and describes the various channels in which globalization may affect inflation while linking them to the experiences of Thailand. Section 3 sets up the DFM to investigate the importance of a global factor for Thai inflation and presents the estimation results. Section 4 lays out the UC model based on the OE-NKPC to examine the empirical relevance of the global output gap in relation to other domestic driving variables for Thai inflation. It also provides a discussion of the empirical findings for the UC model that is augmented with external factors. Section 5 concludes.

2. The Global Dimension of Inflation

Inflation in Thailand has been remarkably low and stable for the past few decades. As shown in Figure 1, the five-year rolling average of Thailand's annual headline and core inflation decelerated sharply since the early 2000s. Furthermore, despite the turmoil from the global financial crisis as well as the recent large swings in global commodity prices, inflation in Thailand has remained remarkably subdued. In past decades, headline inflation has only been as high as 2.6 percent, which is a welcomed contrast to the average of 5.4 percent prior to the year 2000. In large part due to the recent slide in global commodity prices, inflation rates in Thailand has dropped even lower, slipping into negative territory for 15 straight consecutive quarters since January 2015.

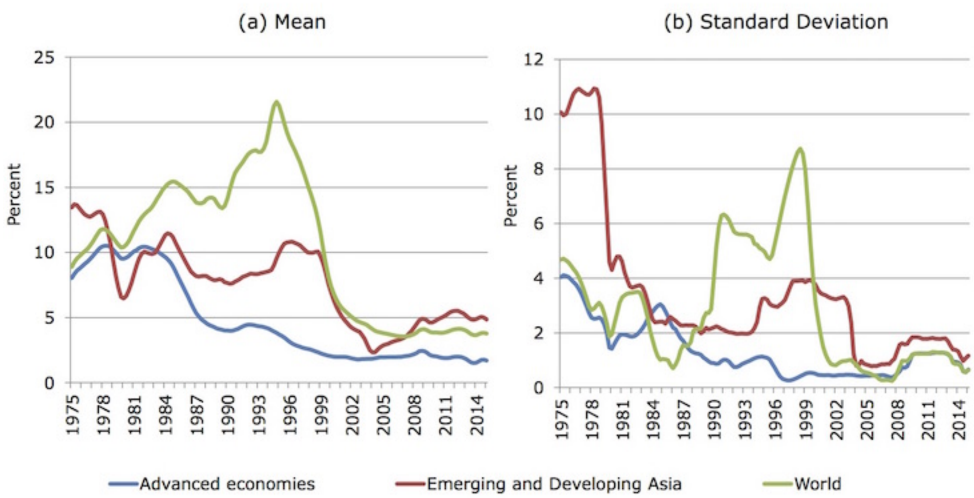
Figure 1: Thai Inflation Mean and Volatility



Note: Inflation is year-on-year changes in the headline consumer price index. The mean and standard deviations are computed using a five-year rolling window. The horizontal axis marks the date at the end of the rolling sample.
Source: Thai Ministry of Commerce, author's calculations.

The inflation experience of Thailand is not country-specific, but echoes the behavior of inflation rates around the world. During recent periods, many advanced economies such as the US and the Euro area have struggled with ultra-low and even negative inflation rates. In fact, low inflation has become a salient feature for most countries during past decades. According to Figure 2, the mean and volatility of inflation rates in advanced economies started to fall in the late 1980s, with emerging market economies following around the year 2000[1]. With inflation rates around the world becoming more low and stable, the degree of co-movement across countries has also increased significantly.

Figure 2: Worldwide Inflation Mean and Volatility



Note: Inflation is year-on-year changes in the aggregated headline consumer price indexes. The mean and standard deviations are computed using a five-year rolling window. The horizontal axis marks the date at the end of the rolling sample.
Source: IMF International Financial Statistics Database, author's calculations.

The relationship between national inflation rates and their underlying driving factors have also changed in recent decades. As documented by IMF (2006, 2013), Pain et al. (2008), and Ball and Mazumder (2011), among others, inflation in a number of advanced economies have become less sensitive to domestic economic conditions. In particular, movements in inflation respond less to changes in domestic slack conditions since the mid 1990s, a phenomenon also known as the flattening of the Phillips curve. Furthermore, a number of studies report a

[1]In Figure 2, world inflation rates are higher than inflation in advanced and Asian emerging countries prior to the year 2000 due to the exceptionally high bouts of inflation in Latin American countries during the debt crisis that struck the region in the 1980s. The average inflation rate in the most densely populated countries in the region, including Mexico, Colombia, Venezuela, Brazil, Bolivia, Uruguay, Peru, Argentina and Chile, nearly touched 160 percent per year in the 1980s and 235 percent in the first half of the 1990s. However, since the late 1990s and early 2000s, inflation rates in these countries have declined dramatically, and in most cases remained low within single digits.

decline in the degree of exchange rate pass-through, particularly in the group of advanced economies (White, 2008). The effect of global commodity price shocks on core inflation rates have also declined dramatically, contributing to the fall in inflation persistence across a number of countries (Cecchetti and Moessner, 2008; Davis, 2012).

Changes in worldwide inflation dynamics coincide with the period of a marked rise in world trade, particularly as emerging countries become more integrated into the world trade system. Trade aggregation began in the 1990s, with the WTO accession of China in 2001 accelerating this process, also known as the 'emerging global factory'. At the same time, Thailand has also become increasingly integrated into global trade systems. The degree of trade openness for Thailand, measured as the country's sum of imports and exports divided by gross domestic product, increased from around 70 percent in the 1990s to over 100 percent in the early 2000s. Given that Thailand has become increasingly exposed to global markets, it is important to understand the channels in which globalization may matter for Thai price processes.

In previous studies, globalization has been suggested to affect inflation through a variety of channels. First, the greater availability of cheap imports from low-cost countries in global markets can directly lower prices through the import price channel. However, the extent in which Thailand may benefit from the tailwinds of globalization through this channel may be rather limited. Although Thailand's trade with lower cost economies have increased over time, the production costs for tradable goods in Thailand did not differ significantly from other low-cost trading partners before global trade accelerated in the early 2000s. Even for advanced economies, the direct impact of lower import prices has been found to be limited as well as short-lasting (see IMF, 2006; Kamin et al., 2006; OECD 2006)[2]. However, note that globalization-fueled growth from emerging countries can also end up driving up world commodity prices, thus ultimately working in the opposite direction for inflation (Adams and Ichino, 1995; Rae and Turner, 2001; Pain et al., 2008)[3].

A second channel through which globalization has been suggested to help keep inflation low is indirectly through enhanced integration of product and factor markets. The entrance of lower cost producers into world trade systems increases the availability of close substitutes abroad, which intensifies competition in domestic markets (Neiss, 2001; Binici et al., 2012).

[2]However, studies have shown that the effects of low cost production on trade prices are likely to be concentrated in particular sectors of the economy. For example, a study by the ECB (2006) show that during 1995-2005, the rising import penetration of low-cost producers in the manufacturing sector has led to a decline in manufacturing import price growth by approximately 2 percent per annum. Feyzioglu and Willard (2006) find that the effect of trade in the US and Japan with China is relatively strong on items such as household furnishings and food.

[3]Apart from its direct effects on prices, world commodity price cycles that have been strengthened by globalization can also explain the enhanced degree of inflation synchronization across countries.

At the same time, a more mobile labor force and the ability for firms to relocate production abroad are other contributing factors that has helped contain producer prices, input prices and markups, which ultimately put downward pressure on inflation. For Thailand, there is some suggestive evidence of the competition effect. As mentioned by Chantanahom et al. (2004), the gap between consumer price inflation (CPI) and producer price inflation (PPI) for Thailand widened significantly since 1999, with PPI rising at a faster rate than CPI. This phenomenon may reflect profit margins of retailers that are being squeezed due to increased competition in Thailand, as the country becomes more open to trade.

Last, greater foreign competition can spur productivity growth through pressures to innovate, as well as to invest in new technologies and production processes. Enhanced integration in trade also enables the spread of the information technology revolution, where advances in communication technology and logistics have helped facilitated the creation of extensive global production chains. Global value chains can provide an opportunity for countries to integrate into the global economy at lower costs (WTO, 2014), and can enhance productivity by allowing countries to specialize in sectors in which they have a comparative advantage. This resulting improvement in productivity in turn lowers the prices of goods relative to the cost of production, thereby keeping inflation low. In addition, as shown by Auer and Mehrotra (2014), intense integration of the manufacturing supply chain among Asian countries seems to have led to more synchronized price movements, as the spillover of shocks from domestic production costs or exchange rates can move through the supply chain with greater ease.

3. A Dynamic Factor Analysis for Thai Inflation

Many studies measure the importance of globalization for inflation by extracting a 'global factor' that captures the common movements underlying international inflation rates. For a group of advanced economies, Ciccarelli and Mojon (2010) find that on average, a single factor can explain nearly 70 percent of the variance in national inflation rates. Based on a rolling principal component analysis, Manopimoke (2015) finds that for a large sample of countries that includes both emerging and advanced economies, the importance of this common factor increased significantly around the year 2000, and during the past few years, accounts for approximately 90 percent of the variability in international inflation rates.

To identify a common factor that is relevant for Thailand, this paper develops a dynamic factor model (DFM) to extract a common component from underlying inflation rate movements in Thailand and its top trading partners. Based on trade statistics, the top trading partners of Thailand are Australia, Hong Kong, Japan, Korea, Malaysia, Philippines, Singapore, the US, Indonesia, the UK, Taiwan, China, and the EU-18 region[4], which accounts for an average of 75 and 79 percent of Thailand's import and export volumes respectively. Due to this large share, the estimated common component from this group of countries will henceforth be referred to as the 'global factor'.

The DFM for inflation is as follows:

$$\pi_{i,t} = \lambda_i^g f_t^g + \lambda_i^r f_t^r + z_{i,t}, \tag{1}$$

$$f_t^g = \gamma_1^g f_{t-1}^g + \gamma_2^g f_{t-2}^c + \eta_t^g, \quad \eta_t^g \sim N(0,1), \tag{2}$$

$$f_t^r = \gamma_1^r f_{t-1}^r + \gamma_2^r f_{t-2}^r + \eta_t^r, \quad \eta_t^r \sim N(0,1), \tag{3}$$

$$z_{i,t} = \gamma_{i,1}^z z_{i,t-1} + \gamma_{i,2}^z z_{i,t-2} + \eta_{i,t}^z, \quad \eta_{i,t}^z \sim N(0, \sigma_i^{z2}), \tag{4}$$

where $\pi_{i,t}$ refers to the inflation rate series of country i, where $i = 1, ..., 14$. Country i represents Thailand and its top 13 trading partners that enters the DFM in no particular order. Based on Eq. (1), the 14 inflation series are decomposed into three components. First, there is a global component f_t^g, which captures the shared movements of all inflation series in the sample. This component may reflect, for example, the effects of global commodity price swings on prices. The second component is an Asia-Pacific regional component f_t^r, which captures movements in inflation that are common only to countries in the Asia-Pacific region. This factor may include the effects of, for example, the underlying regional conditions that led to the build up in inflationary pressures during the Asian financial crisis. Last is a country-specific component $z_{i,t}$, which captures the remaining price pressures in each country that stem from within-country demand and supply shocks.

Next, the coefficients λ_i^g and λ_i^r are factor loadings that reflect the degree in which the variation in $\pi_{i,t}$ can be explained by global and regional components respectively. As is standard in these class of models, the factor loading on the country-specific component is normalized to one. Following Eqs. (2)-(4), the three latent factors follow an autoregressive process (AR) of order 2. The variances of η_t^g and η_t^r are restricted to one and all factors are assumed to be uncorrelated with other factor innovations at all leads and lags for identification purposes. Also, the factor loadings on the regional factor for countries that do not belong to the Asia Pacific region (US, UK, EU-18) are constrained to zero.

[4]Countries in the EU-18 region include Belgium, Germany, Estonia, Ireland, Greece, Spain, France, Italy, Cyprus, Latvia, Luxembourg, Malta, the Netherlands, Austria, Portugal, Slovenia, Slovakia and Finland.

Table 1: **Estimation Results from the Dynamic Factor Model [1993Q1-2015Q1]**

Parameters	AR(2) coefficients of global and regional components
γ_1^g	1.275***(0.108)
γ_2^g	-0.550***(0.102)
γ_1^r	1.355***(0.119)
γ_2^r	-0.475***(0.122)

Estimates of Factor Loadings and Country-Specific Components

Country	λ_t^g	λ_t^r	γ_1^z	γ_2^z	σ_z
Australia	0.336***	-0.049	1.009***	-0.210**	0.609***
	(0.069)	(0.088)	(0.103)	(0.107)	(0.048)
China	0.364*	-0.350*	1.256***	-0.414***	1.718***
	(0.196)	(0.193)	(0.097)	(0.098)	(0.128)
EU-18	0.253***	-	1.028***	-0.075	-0.205***
	(0.029)		(0.154)	(0.158)	(0.021)
Hong Kong	0.430***	0.425***	1.116***	-0.144	0.781***
	(0.096)	(0.102)	(0.110)	(0.111)	(0.064)
Indonesia	0.079	0.069	1.548***	-0.765***	3.824***
	(0.383)	(0.334)	(0.063)	(0.058)	(0.275)
Japan	0.155***	0.109*	0.877***	-0.007	0.547***
	(0.056)	(0.064)	(0.103)	(0.109)	(0.040)
Korea	0.250**	0.544***	0.663***	0.031	0.726***
	(0.097)	(0.088)	(0.141)	(0.140)	(0.071)
Malaysia	0.452***	0.281***	1.088***	-0.410***	0.582***
	(0.084)	(0.071)	(0.106)	(0.100)	(0.049)
Phillipines	0.150	1.108***	1.380***	-0.693 ***	0.774***
	(0.154)	(0.141)	(0.090)	(0.026)	(0.145)
Singapore	0.317***	0.005	1.058***	-0.189	0.673***
	(0.108)	(0.000)	(0.156)	(0.161)	(0.052)
Taiwan	0.399***	0.151*	0.536***	0.132	0.829***
	(0.080)	(0.084)	(0.107)	(0.114)	(0.062)
Thailand	0.690***	0.380***	1.245***	-0.439***	0.620***
	(0.097)	(0.096)	(0.109)	(0.108)	(0.059)
United Kingdom	0.282***	-	1.388***	-0.512***	0.368***
	(0.046)		(0.078)	(0.080)	(0.031)
United States	0.500***	-	0.822***	-0.137	0.369***
	(0.058)		(0.148)	(0.155)	(0.044)

Log-likelihood value: -462.391

Note: ***,**,* denotes statistically significance at the 1, 5 and 10 percent levels respectively. Standard errors are in parentheses.

To estimate the DFM, the inflation series for each country is calculated as the demeaned year-on-year changes in the log CPI index obtained from the International Monetary Fund's International Financial Statistics (IFS) database for the 1993Q1-2015Q1 sample period. The Kalman filter is used to estimate the 30 equations in the DFM, which is comprised of fourteen inflation equations as specified by Eq. (1), one equation for the global factor that follows Eq. (2), one equation for the regional factor as specified by Eq. (3), and fourteen equations for the country-specific factor that follows Eq. (4).

The estimation results are reported in Table 1. As shown, the sum of the AR coefficients for the global and regional components are 0.725 and 0.879 respectively, suggesting that these components are highly persistent. Similarly, the country-specific factor for Thailand displays a high degree of persistence as the sum of its AR coefficients is as high as 0.806. As for the factor loadings, they suggest that the importance of the regional component for Thai inflation is considerable ($\lambda^r = 0.38$) but not as high as the importance of the regional factor for the Philippines ($\lambda^r = 1.11$). However, the loading factor on the global factor for Thailand is the highest among all countries ($\lambda^g = 0.69$). This is not surprising given that the subset of countries chosen for the analysis is tailored to Thailand's trade structure.

Figure 3 plots the estimated factors from the DFM against actual inflation rates. The global component tracks the average value of the 14 inflation series well, particularly since the year 2000 (Figure 3a). For Thailand, the domestic and regional components move closely with actual inflation prior to the year 2000, but these relationships break down in the period thereafter (Figures 3b, 3c). As shown, the global component becomes the prominent driving variable for Thai inflation instead (Figure 3d). This finding implies that since the early 2000s, price processes in Thailand depend increasingly on global economic conditions and less so on shocks that originate from domestic and regional markets.

Given the more prominent role for the global factor in the second part of the sample, a DFM that allows for one structural break in the model parameters may be more appropriate. A DFM with one endogenously determined structural break can be written as follows[5]:

$$\pi_{i,t} = \lambda^g_{i,S_t} f^g_t + \lambda^r_{i,S_t} f^r_t + z_{i,t}, \tag{5}$$

$$f^c_t = \gamma^g_{S_t} f^g_{t-1} + \eta^g_t, \quad \eta^g_t \sim N(0,1), \tag{6}$$

$$f^r_t = \gamma^r_{S_t} f^r_{t-1} + \eta^r_t, \quad \eta^r_t \sim N(0,1), \tag{7}$$

$$z_{i,t} = \gamma^z_{i,S_t} z_{i,t-1} + \eta^z_{i,t}, \quad \eta^z_{i,t} \sim N(0, \sigma^{z2}_{i,S_t}), \tag{8}$$

[5]In contrast to the previous DFM specification with no structural breaks, the dynamic factors in the one structural break model has AR(1) instead of AR(2) dynamics to reduce the number of parameters being estimated. Doing so does not affect the estimation results.

where $S_t = \{1, 2\}$ is a first-order Markov-switching variable with the following matrix of transition probabilities:

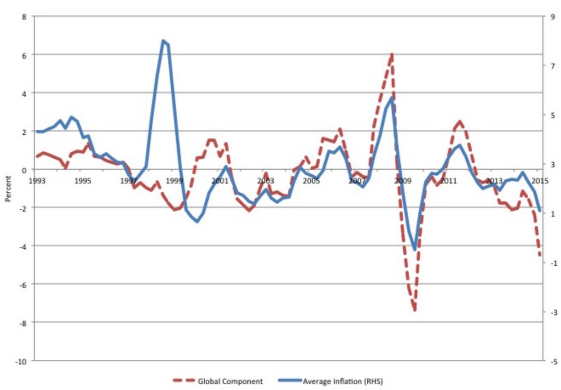
$$P = \begin{bmatrix} p_{11} & 1 - p_{11} \\ 0 & 1 \end{bmatrix}.$$

Note that the $(i, j) - th$ element in the above matrix refers to $Pr[S_t = j | S_{t-1} = i]$.

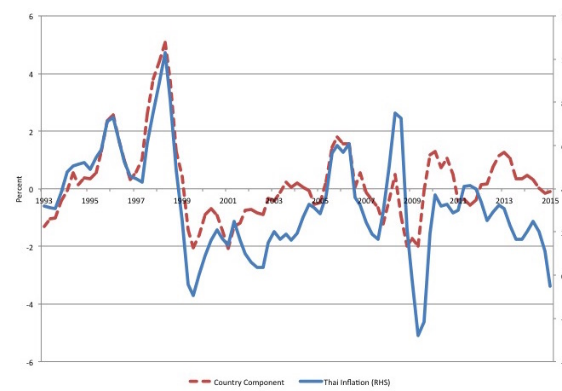
The estimation results from the one-break DFM model is reported in Table 2. Only the parameters that undergo a structural break are reported due to space considerations. As expected, the model finds a structural break in 2001Q1. Analyzing the results further shows that prior to the structural break date, the regional and global factors share approximately equal weight in explaining the overall movements in Thai inflation. However, after the structural break, the regional factor is no longer statistically significant while the factor loading on the global component almost doubles. Compared to the DFM with no structural

Figure 3: Components of the Dynamic Factor Model

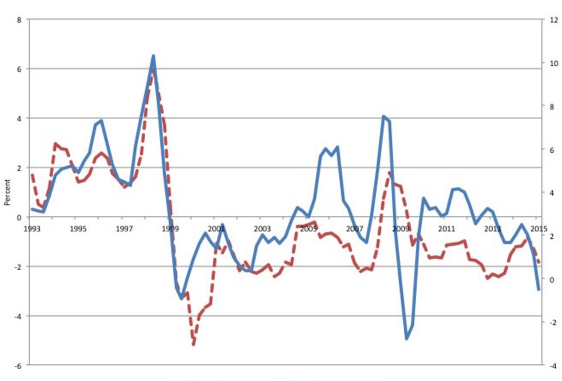
(a) Global Factor and Average Inflation



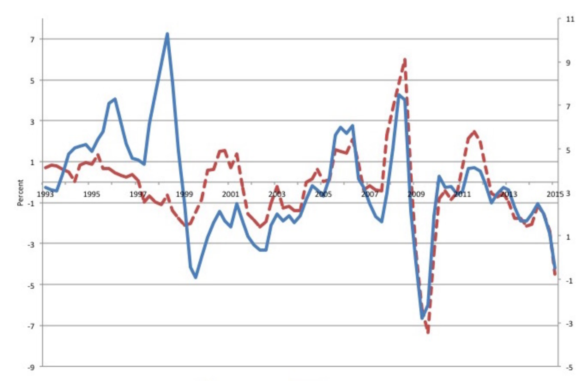
(b) Country Factor and Thai Inflation



(c) Regional Factor and Thai Inflation



(d) Global Factor and Thai Inflation



Sources: IMF International Financial Statistics, author's calculations.

Table 2: **Parameter Estimates from the Dynamic Factor Model with One Structural Break [1993Q1-2015Q1]**

Parameters	Regime 1	Regime 2
λ^g	0.557**(0.223)	1.108***(0.137)
λ^r	0.644***(0.193)	0.148(0.107)
σ^z	0.768***(0.116)	0.598***(0.079)
γ^z	0.874***(0.078)	0.758***(0.108)
Break date	0.976**(0.023) $\rightarrow$ 2001Q1	

Note: ***,**,* denotes statistically significance at the 1, 5 and 10 percent levels respectively. Standard errors are in parentheses.

break, these findings give more concrete evidence that the global factor is an important driving factor for Thai inflation dynamics since the early 2000s, suggesting that the enhanced pace of globalization has had a profound impact on Thai inflation since then.

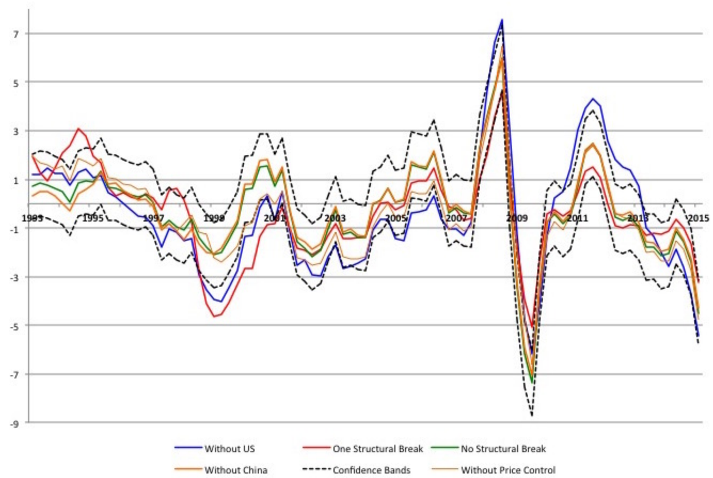
For robustness checks, the global factor belonging to the one-break model is plotted alongside the global factor belonging to the no-break specification, as shown in Figure 4. Additionally, to ensure that estimates of the global factor are not influenced by inflation rate movements in any one country in particular, the plot also contains global factors from the following three models: (1) a DFM that excludes US inflation; (2) a DFM that excludes China inflation; and (3) a DFM that excludes the inflation rate series of Thailand, Indonesia, and the Philippines. While the exclusion of the US and China is self-explanatory, excluding the three Asian countries in the third model is motivated by the fact that these countries rely heavily on price control policies that may potentially distort the DFM analysis. However, judging from the plots, apart from a brief period around the year 2000 which is associated with structural instability, all estimated global factors lie within the 95 percent confidence bands of the global factor from the no-break DFM specification. This finding implies that all global factors are not statistically different.

Finally, to examine the relative importance of the global, regional and country-specific factors in explaining the overall variability in Thai inflation, a variance decomposition according to the one-break DFM estimation results is computed[6]. According to Figure 5, all three components are equally important in explaining the observed variability in Thai inflation prior to 2001. However, in the period thereafter, the importance of this global factor increased significantly, while the significance of country-specific factor falls to below

[6]Under the assumption that the components are orthogonal, the share of inflation variance explained by world, regional and country-specific components can be computed as: $S_i^w = \frac{\frac{\lambda_i^{c2}}{(1-\gamma^{c2})}}{Var(\pi_i)}$, $S_i^r = \frac{\frac{\lambda_i^{r2}}{(1-\gamma^{r2})}}{Var(\pi_i)}$, $S_i^z = \frac{\frac{\sigma_i^{z2}}{(1-\gamma^{z2})}}{Var(\pi_i)}$ where $Var(\pi_i) = \frac{\lambda_i^{c2}}{(1-\gamma^{c2})} + \frac{\lambda_i^{r2}}{(1-\gamma^{r2})} + \frac{\sigma_i^{z2}}{(1-\gamma^{z2})}$. Note that $Var(f^c) = Var(f^r) = 1$.

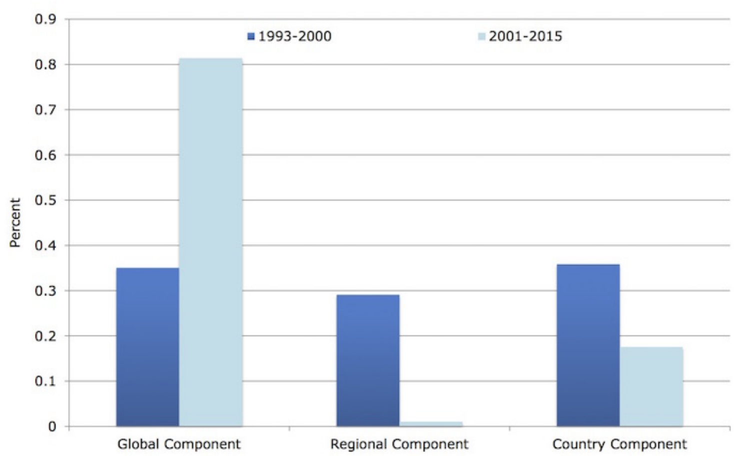
20 percent. The role of the regional component also declined significantly. Thus overall, the empirical analyses in this section provide convincing evidence that the increasing pace of globalization since the early 2000s served to reduce the role of domestic and regional factors in explaining Thai inflation rate movements, while at the same time significantly enhancing the role of global ones.

Figure 4: Global Components from Various Dynamic Factor Models [1993Q1-2015Q1]



Sources: IMF International Financial Statistics, author's calculations.

Figure 5: Variance Decomposition for the DFM with One Structural Break [1993Q1-2015Q1]



Sources: IMF International Financial Statistics, author's calculations.

4. An Unobserved Components Model for the Open Economy New Keynesian Phillips Curve

What are the economic forces behind the sizable influence of the global factor for Thai inflation? Typically, inflation is driven by the country's degree of tightness or slack in resources, captured by the domestic output gap. However, as an economy becomes more internationalized in goods and financial markets, many have conjectured that national inflation rates should become more sensitive to global economic conditions - a notion that is often referred to as the globalization hypothesis. In the Phillips curve framework, the globalization hypothesis translates to inflation in a given country becoming more sensitive to a global rather than a domestic output gap as that country becomes more heavily engaged in international trade[7].

Consider the following open economy New Keynesian Phillips curve (OE-NKPC)[8]:

$$\pi_t = \beta E_t(\pi_{t+1}) + kx_t + k^* x_t^* + \Gamma_t, \tag{9}$$

where π_t is the current inflation rate; β is the subjective discount factor, $E_t(.)$ denotes expectations formed conditional on information up to time t; x_t is the domestic output gap; and x_t^* is the global output gap defined as the difference between actual and potential world output. The coefficients on the output gaps, k and k^* are functions of the deep structural parameters of the economy such as the frequency of price adjustment, the elasticity of substitution between home and foreign goods, and the degree of openness which is inversely related to the home bias in consumption preferences. Γ_t captures shocks to inflation that are not captured by the OE-NKPC, and may include price pressures from terms of trade, the deviation of import prices from the law of one price, and the deviations of the real exchange rate from purchasing power parity[9].

As discussed earlier, the globalization hypothesis implies that as an economy becomes

[7]Note that the domestic output gap already contains some information about global influences on domestic costs and prices as net exports are an integral part of real GDP. Nevertheless, the global output gap can still matter for inflation by capturing additional slack influences from foreign sources not contained in the domestic output gap. For example, the global output gap may capture rising cost pressures from abroad that can raise import prices, the amount of spare capacity overseas that may ultimately reduce workers' bargaining power, or the restraining effects of enhanced competition abroad on domestic producers' markups.

[8]The appeal of the OE-NKPC is that it is derived from a general equilibrium framework based on optimizing behavior of monopolistically competitive firms, giving the model solid microfoundations (see Clarida et al., 2002; Corsetti and Pesenti, 2005).

[9]In an OE-NKPC, the specific form of Γ_t depends on the underlying assumptions of the economy, such as whether the exporting firms engage in local or producer currency pricing. For example, the OE-NKPC derived under producer currency pricing would not involve a term with the real exchange rate, as the law of one price holds and the degree of pass through is complete. A rigorous microfoundation of the OE-NKPC under various assumptions can be found in Clarida et al. (2002), Corsetti and Pesenti (2005), Martínez-Garcia and Wynne (2010), and Zaniboni (2011).

more open, the role of the domestic output gap in Eq. (9) should decline while the global slack measure should gain prominence[10]. Many studies have investigated whether this holds true in the data, but empirical support for the globalization hypothesis has been far from robust. For example, based on different sample periods, selection of countries, and various specifications for the OE-NKPC, Gamber and Hung (2001), Borio and Filardo (2007), and Manopimoke (2015) show that global capacity constraints play an influential role for national inflation dynamics. On the other hand, Tootell (1998), Calza (2009), Ihrig et al. (2007), and Milani (2010) finds little or no predictive power for the global output gap for CPI across a number of countries.

One reason for conflicting results in the literature stem from the difficulty in handling the expectational element $E_t\pi_{t+1}$. How one chooses to measure one-period-ahead inflation expectations is most likely to affect the estimated coefficients on the output gaps. For example, Borio and Filardo (2007) use HP-filtered inflation series as a proxy for the underlying long-run trend rate of inflation. As a result, ample variability and persistence in the inflation series remains, helping them find an important role for the global output gap across a number of developed economies. Ihrig et al. (2007) point out that this approach is prone to model misspecification as it leaves autocorrelation in the residuals. Instead, they proxy the expectational element with lags of inflation, and for a similar data sample, they find no significant role for the global output gap in the OE-NKPC.

To avoid dealing with the inflation expectation element, this paper considers estimation of the closed-form OE-NKPC, which can be obtained by forward iteration of Eq. (9). The resulting specification becomes:

$$\pi_t = \lim_{j\to\infty} \beta^j E_t(\pi_{t+j}) + k\sum_{j=0}^{\infty} \beta^j E_t(x_{t+j}) + k^* \sum_{j=0}^{\infty} \beta^j E_t(x^*_{t+j}) + \tilde{z}_t, \tag{10}$$

where $\tilde{z}_t = \sum_{j=0}^{\infty} E_t(\Gamma_{t+j})$. The first element on the right-hand-side of the above equation represents long-term inflation expectations, but vanishes to zero under the assumption of stationary inflation. However, based on various unit root tests, the null hypothesis of a unit root for Thai inflation cannot be rejected. Therefore, in the spirit of Kim et al. (2014), $\lim_{j\to\infty} \beta^j E_t(\pi_{t+j})$ is interpreted as the Beveridge-Nelson stochastic trend, which can be approximated by a driftless random walk. Note that the remaining three terms on the right-

[10]Theoretically, there are two sides of the camp to this line of argument. Razin and Yuen (2002), Razin and Loungani (2005), and Razin and Binyamini (2007) argue that the opening of the capital account and trade balance reduces the sensitivity of inflation to domestic real activity conditions through channels such as enhanced consumption smoothing and greater consumption diversification. A less popular view is based on the Barro-Gordon framework, where Romer (1993) and Rogoff (2003, 2006) argue that global competition reduces the monopoly power of firms and workings, which increases competition in the markets for goods, services and labor. Increased flexibility in these markets in turn increase the slope of the Phillips curve.

hand-side of Eq. (10) are deviations of actual inflation from a time-varying long-run trend, and captures transitory movements in inflation that together make up an inflation gap.

The above model is an open-economy version of the closed-form NKPC as estimated by Kim et al. (2014). Similar to these authors, the OE-NKPC can be written as an unobserved components (UC) model for inflation, which is henceforth referred to as the baseline specification:

Baseline specification

$$\pi_t = \bar{\pi}_t + k\sum_{j=0}^{\infty} E_{t-1}(x_{t+j}) + k^*\sum_{j=0}^{\infty} E_{t-1}(x^*_{t+j}) + z_t \tag{11}$$

$$\bar{\pi}_t = \bar{\pi}_{t-1} + e_t, \quad e_t \sim i.i.d.N(0, \sigma_e^2) \tag{12}$$

$$z_t = \psi z_{t-1} + \eta_t, , \quad \eta_t \sim i.i.d.N(0, \sigma_\eta^2) \tag{13}$$

$$x_t = \phi_1 x_{t-1} + \phi_2 x_{t-2} + v_t, \quad v_t \sim i.i.d.N(0, \sigma_v^2) \tag{14}$$

$$y^*_t = \tau^*_t + x^*_t, \tag{15}$$

$$\tau^*_t = \delta^*_1 D_{1t} + \delta^*_2 D_{2t} + \delta^*_3 D_{3t} + \tau^*_{t-1} + w^*_t, \quad w^*_t \sim i.i.d.N(0, \sigma_w^{*2}), \tag{16}$$

$$x^*_t = \phi^*_1 x^*_{t-1} + \phi^*_2 x^*_{t-2} + v^*_t, \quad v^*_t \sim i.i.d.N(0, \sigma_v^{*2}). \tag{17}$$

where

$$D_{1t} = \begin{cases} 1, & \text{if } 1993Q1 \leq t < 1997Q3 \\ 0, & \text{otherwise} \end{cases}, \tag{18}$$

$$D_{2t} = \begin{cases} 1, & \text{if } 1997Q3 \leq t < 2007Q4, \\ 0, & \text{otherwise} \end{cases}, \tag{19}$$

$$D_{3t} = \begin{cases} 1, & \text{if } t \geq 2007Q4, \\ 0, & \text{otherwise.} \end{cases}, \tag{20}$$

The above equations are described in turn. First, Eq. (11) is similar to Eq. (10) except that long-horizon inflation expectations are replaced by a driftless random walk; the discount factor β is calibrated to one[11]; and for feasible estimation of the model, the terms $\sum_{j=0}^{\infty} E_t(x_{t+j})$ and $\sum_{j=0}^{\infty} E_t(x_{t+j})$ are replaced with $\sum_{j=0}^{\infty} E_{t-1}(x_{t+j})$ and $\sum_{j=0}^{\infty} E_{t-1}(x_{t+j})$ respectively. In this way, $z_t = k(\sum_{j=0}^{\infty} E_t(x_{t+j}) - \sum_{j=0}^{\infty} E_{t-1}(x_{t+j})) + k^*(\sum_{j=0}^{\infty} E_t(x^*_{t+j}) -$

[11]The discount factor is typically set to 0.99. Calibration of the discount factor to 0.99 did not change the quantitative results from the model.

$\sum_{j=0}^{\infty} E_{t-1}(x^*_{t+j})) + \tilde{z}_t$, and is no longer correlated with the infinite sum terms. However, z_t is now a function of the output gaps, thus the innovation of shocks to inflation are allowed to be correlated with those from the output gap, ie. $Cov(\eta_t, v_t) \neq 0$ and $Cov(\eta_t, v^*_t) \neq 0$.

With the domestic output gap being cyclical in nature, x_t is assumed to follow an AR(2) process, while z_t follows an AR(1) process to capture potential serial correlation in Γ_t. During estimation, x_t is proxied by estimates of the domestic output gap for Thailand. However, no standard proxy exists for the global output gap x^*_t, thus it is treated as a latent state variable with AR(2) dynamics, with movements that are to be extracted from a trend-cycle decomposition for world output y^*_t[12]. According to Eq. (16), trend output τ^*_t is a random walk with drift, where the drift term δ^* incorporates dummy variables as defined by Eqs. (18)-(20) to capture potential slowdowns in output trend growth during the Asian financial crisis, as well as during the most recent global financial crisis.

4.1 Incorporating Structural Breaks

From the DFM analysis in Section 3, a structural change in the relationship between Thai inflation and its global determinants occurred in 2001. To investigate whether the OE-NKPC underwent similar changes, structural breaks are incorporated into the parameters of the baseline UC specification as follows:

$$\pi_t = \bar{\pi}_t + k_{S_t} \sum_{j=0}^{\infty} E_{t-1}(x_{t+j}) + k^*_{S_t} \sum_{j=0}^{\infty} E_{t-1}(x^*_{t+j}) + z_t, \tag{11$'$}$$

$$\bar{\pi}_t = \bar{\pi}_{t-1} + e_t, \quad e_t|S_t \sim i.i.d.N(0, \sigma^2_{e,S_t}), \tag{12$'$}$$

$$z_t = \psi_{S_t} z_{t-1} + \eta_t, \quad \eta_t|S_t \sim i.i.d.N(0, \sigma^2_{\eta,S_t}), \tag{13$'$}$$

where the remaining equations of the baseline UC specification that describes output remain unchanged. Evidently, S_t is a latent state variable that captures regime changes in inflation dynamics.

The number of structural breaks are not imposed upon the model, but are determined via a number of diagnostic tests that are carried out in Appendix A beforehand. The best-fitting

[12] An alternative approach to estimate the global output gap is by applying the HP-filter to y^*_t. While this method is common in the literature, the HP-filtered output gap is a purely statistical measure. Treating x^*_t as a latent state variable in the UC framework is thus preferable as it yields estimates of the output gap that are consistent with the OE-NKPC. Also, note that technically, x^*_t is to be considered as the foreign output gap for Thai inflation. However, Thailand is a small open economy, thus there should be negligible differences between foreign and global output gap measures. Accordingly, x^*_t is referred to as the global output gap.

model turns out to be one that allows for the occurrence of two structural breaks or three regimes, thus $S_t = \{1, 2, 3\}$ and is defined as a first-order Markov-switching variable with the following matrix of transition probabilities:

$$P = \begin{bmatrix} p_{11} & 1-p_{11} & 0 \\ 0 & p_{22} & 1-p_{22} \\ 0 & 0 & 1 \end{bmatrix},$$

where the $(i,j)-th$ element refers to $Pr[S_t = j|S_{t-1} = i]$. Since the regimes in the model are assumed to be non-recurring, p_{33} is the absorbing state and has probability equal to one[13]. Finally, the UC model with two structural breaks can be estimated with the Kim filter once put into state-space form. Readers are referred to Appendix B for a state-space representation of the model.

4.2 Empirical Results

The dataset for estimation of the UC model is quarterly data that spans 1993Q1-2015Q1, which is based on data availability. Inflation is calculated as the changes in the seasonally-adjusted CPI. The domestic output gap x_t is proxied by the Bank of Thailand's (BOT) estimate, which is obtained from a multivariate model for inflation, output and interest rates. The global output y_t^* is the aggregated measure of PPP-adjusted GDP from Thailand's top trading partners, where each country is weighted based on its time-varying trade share with Thailand[14].

Table 3 reports the estimation results from the two structural break UC model. First, the findings suggest that there have been two distinct structural changes that occurred in 2001Q1 and 2007Q1, separating Thai inflation dynamics into three regimes; 1993-2001, 2001-2007, and 2007-2015. Across the three regimes, while there appears to be no changes in the variability of shocks to trend inflation (σ_e), Figure 6 shows that the level of trend inflation declined significantly since the early 2000s and remained remarkably stable thereafter at an average level of 2.4 percent. Compared to estimates of trend inflation from HP-filtered estimates that are plotted alongside, trend inflation from the UC model underwent a sharper decline in 2001 and contains less variability in the subsequent period. Therefore, the UC model results lend more support to the view that the shift in monetary policy towards an

[13] For robustness checks, the baseline UC model was also estimated with Markov-switching parameters while allowing for recurring states, but there was no evidence of recurrence in the regimes.

[14] Other weighting schemes based on the GDP size of trading partners are also considered for robustness checks. The estimation results are qualitatively similar and are available upon request. The set of countries chosen to represent Thailand's top trading partners are the same as for the DFM analysis.

inflation targeting framework in May 2000 successfully lowered as well as anchored long-term inflation expectations in Thailand at a level that is well within the BOT's target range[15].

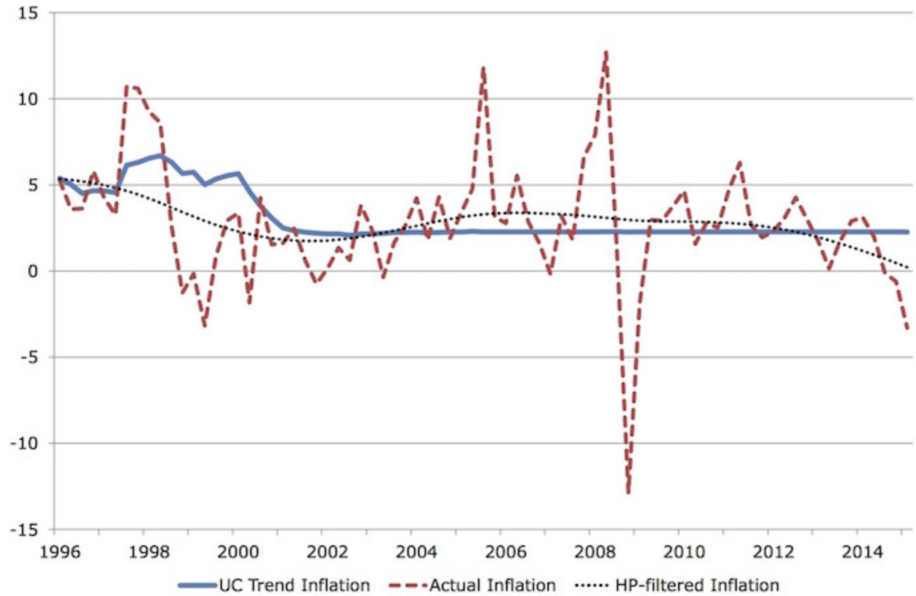
Table 3: **Estimation of UC OE-NKPC with two structural breaks [1993Q1-2015Q1]**

Phillips Curve Parameters	
ψ_1	0.861***(0.094)
ψ_2	-0.643***(0.122)
ψ_3	0.260(0.166)
k_1	0.178*(0.098)
k_2	0.074(0.048)
k_3	-0.050(0.063)
k_1^*	-0.425(0.266)
k_2^*	0.321***(0.104)
k_3^*	0.389*(0.236)
$\sigma_{e,1}$	0.001(0.000)
$\sigma_{e,2}$	0.000(0.000)
$\sigma_{e,3}$	0.000(0.000)
$\sigma_{\eta,1}$	2.553***(0.401)
$\sigma_{\eta,2}$	0.871***(0.146)
$\sigma_{\eta,3}$	3.521***(0.396)
Output Parameters	
ϕ_1	0.765***(0.112)
ϕ_2	0.017(0.105)
ϕ_1^*	1.556***(0.090)
ϕ_2^*	-0.701***(0.069)
σ_v	2.478***(0.200)
σ_v^*	0.408***(0.092)
δ_1^*	0.974***(0.113)
δ_2^*	0.960***(0.038)
δ_3^*	0.836***(0.053)
σ_w^*	0.128(0.096)
ρ_{vw}^*	0.999***(0.002)
Transition Probabilities	
p_{11}	0.970***(0.032)
Break Date	2001Q1
p_{22}	0.957***(0.046)
Break Date	2007Q1
Log-likelihood value:	-283.618

Note: Standard errors are in parentheses. ***,**,* denote statistical significance at the 1, 5, and 10 percent levels respectively.

[15] At first, the BOT targeted core inflation within a range of 0-3.5 percent. Since 2009, the band has been narrowed to 0.5-3 percent, and in 2015, the Bank altered its target to correspond to headline inflation at 2.5 percent with bands of plus and minus 1.5 percent.

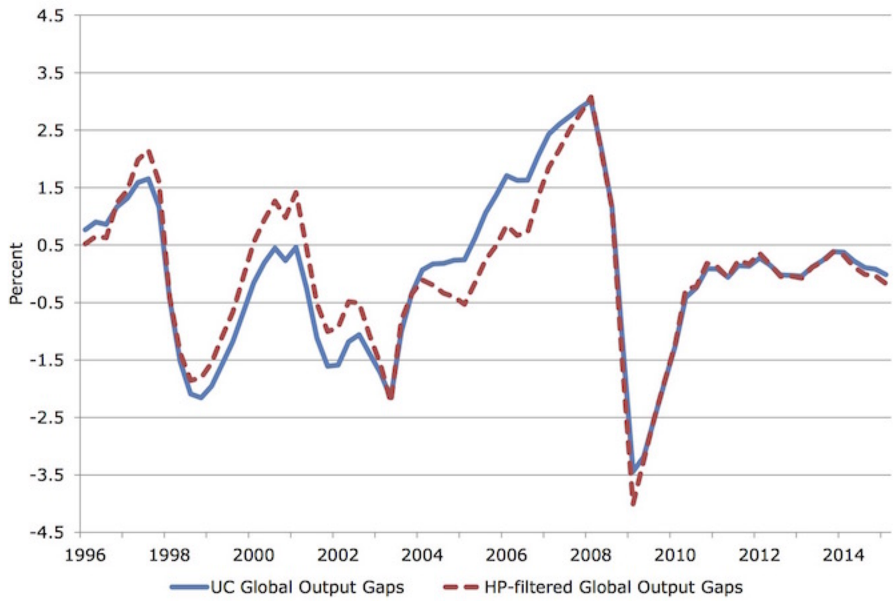
Figure 6: Actual Inflation and Trend Inflation



In line with the finding of a lower and more stable inflation trend, the degree of persistence in z_t also declined over the postwar period, suggesting that inflation returns to target faster after a temporary shock. However, the variability of shocks to z_t while muted in the period after 2001, approximately quadrupled in the post 2007 era. In large part, this burst in volatility captures commodity price swings that occurred during that time. Nevertheless, it is interesting to note that while z_t became more volatile, compared to the pre 2001 era which is also characterized by volatile transitory shocks, the effect of these shocks on inflation have become shorter lived. This is reflected by the lower levels of persistence in z_t, which is most likely a byproduct of better anchored inflation expectations.

Turning to examine the dynamics of output, the sum of the AR(2) coefficients for the domestic output gap is 0.7, suggesting that Thailand's output gap is highly persistent. The estimated global output gap is also persistent, but shocks to the global gap are less volatile when compared to the variability of shocks to the domestic output gap. Figure 7 plots the global output gap from the UC model alongside a global output gap obtained from applying a HP-filter to y_t^*. As shown, the two global gap measures are similar in terms of the peaks and troughs in the world business cycle that they capture, but they also deviate to some extent prior to the global financial crisis. Last, from estimates of δ_t^*, the trend growth rate of world output did not change much after the Asian financial crisis but declined significantly after the recent global financial crisis.

Figure 7: UC and HP-filtered global output gaps



Finally, much of the focus in this paper is on whether the output gap coefficients, k and k^{*}, have been altered as a result of globalization. The sensitivity of inflation to the domestic output gap as captured by the magnitude of k declined since 2001, which is a result that is consistent with the experiences of many advanced economies (IMF, 2006, 2013; Borio and Filardo, 2007). For Thailand, this result is in line with the findings of Chantanahom et al. (2004). These authors use sectoral price data in Thailand to find that the frequency of price adjustments, which is positively related to the coefficient on the domestic output gap in the Phillips curve, declined ever since the Asian financial crisis.

The reduced role for domestic slack conditions in explaining Thai inflation rate movements is replaced by an enhanced role for the global output gap. Since 2001, the coefficient on the global output gap increased significantly, from a statistically insignificant estimate of -0.4 which is of the wrong sign, to a statistically significant estimate of 0.3 in the periods thereafter. Overall, these results are in line with the previous findings from the DFM analysis. The sizable global factor in the DFM that moves closely with Thai inflation since 2001 most likely corresponds to an enhanced role for the global output gap in the OE-NKPC which occurred during the same time.

4.3 The Role of External Factors

The relevance of global output gaps for inflation may be due to the direct effects of import prices on inflation. For example, a tighter global output gap reflecting rapid growth in emerging market economies may be associated with higher commodity prices that, in turn, feed into domestic prices. To investigate this possibility, the z_t process in Eq. (13) of the baseline UC model is modified as follows:

$$z_t = \alpha_{S_t}\Gamma_{t-1} + \eta_t, \tag{21}$$

$$\eta_t = \psi\eta_{t-1} + \epsilon_t, \quad \eta_t|S_t \sim i.i.d.N(0, \sigma^2_{\epsilon,S_t}), \tag{22}$$

where Γ_t represents the following external factors: import prices, changes in oil prices, and changes in non-oil commodity prices[16]. These factors are added one at a time to the baseline UC specification with two structural breaks to investigate whether the global output gap is important only because it is capturing foreign price pressures from the direct import price channel. The z_t process is related to these external factors via the coefficient α, and the remaining shocks to z_t are captured by η_t which is assumed to follow an autoregressive process of order 1.

The estimation results for the augmented UC model are reported in Table 4. Focusing on the first three columns, the inclusion of external factors in the second regime did not reduce the size of the coefficient on the global slack measure, nor removed its statistical significance. This finding implies that during 2001-2007, the global output gap was capturing global influences for inflation beyond the direct import price channel, such as the indirect effects of globalization from enhanced competition, or gains in productivity from technological spillovers.

On the other hand, the inclusion of external factors reduced the coefficients on the global output gap and rendered it statistically insignificant in the post 2007 period. Correspondingly, estimates of α_3 for import price inflation, and changes in oil and non-oil commodity prices become statistically significant in the third regime. In turn, these findings imply that the global output gap was important in the post 2007 period merely because it was capturing the direct effects of import prices on domestic inflation. Investigating further however, import prices are important mainly because of the large influence of world oil price movements. More specifically, once import price inflation is separated into oil and non-oil components, only the coefficient on oil prices remained statistically significant. Similarly, the effect of non-oil commodity prices for inflation appears important only because it was acting as a proxy for changes in world oil prices, since both series are highly correlated after 2007 (Figure 8a). When both oil and non-oil commodity prices are added to the UC specification,

[16]Note that oil prices are in US dollars, so the foreign exchange rate between the Thai baht and the US is added to the UC specification when oil is a regressor in the model. The same applies for non-oil commodity prices. However, the estimated coefficients on the exchange rate are not reported because they are not statistically significant.

only the coefficient on oil commodity prices remained statistically significant. In sum, these findings highlight the importance of world oil price movements in explaining global output gap dynamics since the post 2007 period.

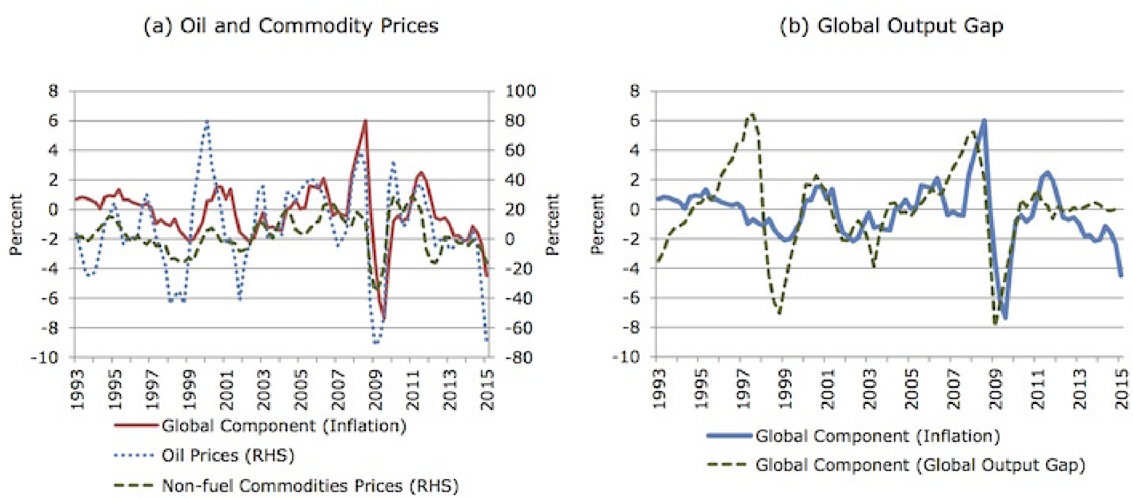
Table 4: **Estimation of the UC OE-NKPC with Two Structural Breaks and External Factors [1993Q1-2015Q1]**

Parameters	Import Inflation	Oil	Non-oil Commodities	Real Exchange Rate
		Phillips Curve Parameters		
ψ_1	0.864***(0.094)	0.857***(0.110)	0.854***(0.090)	0.850***(0.114)
ψ_2	-0.651***(0.122)	-0.626***(0.131)	-0.644***(0.123)	-0.630***(0.116)
ψ_3	0.142(0.169)	0.098(0.191)	0.247(0.161)	0.270(0.179)
k_1	0.170*(0.095)	0.183*(0.099)	0.179*(0.093)	0.152*(0.088)
k_2	0.084(0.057)	0.090(0.056)	0.074*(0.042)	0.111*(0.062)
k_3	0.019(0.045)	0.023(0.041)	0.033(0.048)	-0.040(0.064)
k_1^*	-0.412(0.280)	-0.345(0.256)	-0.491(0.316)	-0.679***(0.314)
k_2^*	0.335***(0.110)	0.215**(0.089)	0.398**(0.178)	0.224**(0.091)
k_3^*	0.098(0.181)	0.172(0.113)	0.144(0.172)	0.369(0.254)
α_1	0.001(0.026)	-0.042(0.046)	0.127(0.177)	-0.140(0.086)
α_2	-0.017(0.027)	0.041*(0.023)	-0.059(0.090)	-0.193**(0.086)
α_3	0.288***(0.047)	0.190***(0.021)	0.379***(0.073)	0.162(0.259)
σ_e	0.000(0.000)	0.000(0.073)	0.000(0.000)	0.000(0.000)
$\sigma_{\eta,1}$	2.519***(0.392)	2.560***(0.403)	2.549***(0.396)	2.434***(0.395)
$\sigma_{\eta,2}$	0.876***(0.155)	0.853***(0.147)	0.846***(0.141)	0.820***(0.130)
$\sigma_{\eta,3}$	2.496***(0.283)	2.024***(0.231)	2.708***(0.306)	3.530***(0.398)
		Output Parameters		
ϕ_1	0.764***(0.112)	0.770***(0.113)	0.773***(0.111)	0.741***(0.115)
ϕ_2	0.019(0.106)	0.012(0.108)	0.008(0.105)	0.047(0.107)
ϕ_1^*	1.580***(0.092)	1.455***(0.119)	1.579***(0.093)	1.541***(0.090)
ϕ_2^*	-0.729***(0.074)	-0.599***(0.100)	-0.732***(0.083)	-0.701***(0.073)
σ_v	2.478***(0.200)	2.477***(0.200)	2.477***(0.200)	2.483***(0.202)
σ_v^*	0.393***(0.092)	0.483***(0.118)	0.407***(0.082)	0.467***(0.090)
δ_1^*	0.962***(0.113)	0.949***(0.111)	0.966***(0.109)	0.953***(0.107)
δ_2^*	0.962***(0.038)	0.975***(0.034)	0.968***(0.037)	0.967***(0.033)
δ_3^*	0.827***(0.052)	0.815***(0.045)	0.824***(0.051)	0.836***(0.048)
σ_w^*	0.145(0.097)	0.046(0.122)	0.130(0.085)	0.064(0.089)
ρ_{vw}^*	0.999***(0.000)	0.999***(0.000)	0.999***(0.000)	0.999***(0.000)
		Transition Probabilities		
p_{11}	0.975***(0.029)	0.975***(0.029)	0.972***(0.0309)	0.970***(0.034)
Break Date	2002Q4	2002Q4	2001Q3	2001Q1
p_{22}	0.964***(0.043)	0.964***(0.042)	0.956***(0.047)	0.964***(0.042)
Break Date	2009Q3	2009Q3	2007Q3	2007Q4
Log-likelihood value:	-269.858	-259.900	-272.737	-279.797

Note: Standard errors are in parentheses. ***,**,* denote statistical significance at the 1, 5, and 10 percent levels respectively.

To explain why the indirect effects of globalization for inflation disappears and why only oil price movements dominate short-run inflation rate movements in the post 2007 period is beyond the scope of this paper. However, some suggestive evidence can be offered. First, further investigations reveal that the importance of oil price movements for Thai inflation is unrelated to domestic oil consumption. For Thailand, the share of energy components in the CPI has already been gradually increasing since the mid 1990s, and in fact, there has been a slowdown in fuel consumption ever since the global financial crisis. Second, Figure 8a, which contains a plot of the estimated global factor for inflation from the previous DFM analysis alongside world oil prices, suggests that a large portion of international inflation rate movements are also closely aligned with world oil price dynamics since 2007. In other words, the importance of oil prices for Thai inflation in the post-crisis period appears to be a phenomenon that is not only specific to Thailand, but is pertinent to other countries as well[17].

Figure 8: Global Inflation and its Determinants



Note: Global inflation and the global output gap is extracted from a dynamic factor model for inflation and output gaps respectively. The change in Dubai oil prices is the log year on year change in the Dubai oil price series.
Sources: IMF International Financial Statistics, author's calculations.

Second, the indirect effects of globalization that are no longer relevant for Thai inflation

[17]The change in the relationship between inflation and world oil prices may be related to the evidence of a structural change in oil price dynamics after the global financial crisis (Arezki et al, 2015; Baffes et al., 2015). In contrast to the pre 2007 period where changes in the prices of oil and other commodities were mainly driven by global demand for resources, particularly from emerging countries such as China and India (Killian, 2009), supply-side factors appeared to have played a more important role in driving oil price changes in the more recent period. This stems from the rapid growth in the production of unconventional oil such as shale oil from the United States as well as the decline of OPEC's share of global oil supply.

beyond the direct import price channel since 2007 also appears to be a global phenomenon. As shown in Figure 8b, the global factor for inflation from the DFM analysis is highly correlated with a global factor for output gaps in Thailand and its top trading partners, but only during the 2001-2007 period. This implies that in the post-crisis period, inflation may no longer be affected by the gains from globalization that indirectly enhance productivity or intensify competition. In a way, this finding may reflect the 'Great Trade Collapse', which characterizes the recent slowdown in global trade and the structural change in the trade-income relationship among a number of countries (Ferrantino and Taglioni, 2014; Boz et al., 2014; Constantinescu et al., 2015).

Last, the augmented UC model can also be used to examine evidence of exchange rate pass-through in Thailand. In doing so, the real exchange rate is added to the UC specification as an external factor, and the estimation results are reported in the last column of Table 4. Based on the estimated α_3 coefficients, the degree of exchange rate pass-through for Thailand is in general low or negligible. These findings are in line with the findings of Buddhari and Chensavasdijai (2003), whom explain that low pass-through in Thailand is typically a result of lower inflation expectations as well as the prevalence of administered price measures. However, to explain the complete lack of pass-through in the first the third regimes, this paper refers to an explanation given by McCarthy (1999), Goldfajn and Werlang (2000) and Styrin et al. (2012). These authors show that there is typically low pass-through in economies where exchange rate variability is high, or when a country enters a low inflation environment. Judging from the economic conditions in the pre 2001 and post 2007 periods, this argument fits with the finding of no pass-through during these periods. In the first regime, exchange rate volatility was relatively high due to the abandonment of the exchange rate peg after the 1997 financial crisis. During the third regime, low inflation was a prominent feature of Thailand's economy due to sliding world oil prices after the global financial crisis.

5. Conclusion

This paper explores the effects of globalization for the behavior of Thai inflation during 1993-2015 based on a dynamic factor model and an unobserved components model for inflation that is consistent with an open-economy New Keynesian Phillips curve. The models carefully account for structural breaks in Thai inflation dynamics which have been found to occur in 2001 and 2007.

The empirical findings strongly suggest that globalization has altered Thai inflation dynamics, particularly in the short-run. The relationship between Thai inflation and the domestic output gap declined dramatically in 2001, coinciding with a significant increase in the role of a global output gap. After the global financial crisis in 2007, the effect of globalization for Thai inflation underwent another structural change. In particular, the sensitivity of inflation to global slack pressures that extend beyond the direct import price channel have been muted, while movements in world oil prices appeared to play a more important role in determining Thailand's overall price movements.

In sum, the findings suggest that since 2001, external developments play a key role in driving inflation developments in Thailand. A direct implication of this result is that monetary authorities must now pay more attention to external developments and respond to a wider range of shocks. More fundamentally, greater sensitivity of inflation to global factors begs the question as to whether the Bank of Thailand still has the ability to control inflation within national borders. Based on the empirical findings, the Bank of Thailand's ability to control inflation remains unabated despite greater globalization. Since the implementation of an inflation targeting framework in May 2000, Thailand has in fact achieved improved inflation outcomes with long-term inflation expectations being well-anchored at an average level of 2.4 percent. Therefore, maintaining good central bank communication is certainly key for inflation control especially in today's global environment where inflation has become more prone to volatile external shocks than ever before.

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Appendix A: Determining the Number of Structural Breaks

To determine the number of structural breaks in Thai inflation dynamics, the baseline UC model is first estimated with no structural breaks. Then, a diagnostic test is carried out to examine whether the model is misspecified, and if so, the model is estimated again with incremental number of structural breaks. This process continues until the appropriate number of structural breaks is determined. As is standard for Markov-switching models, diagnostic tests are based on a Ljung-Box test that is applied to the standardized residuals and the squares of the standardized residuals of the inflation series. The underlying idea of the test is that serial correlation in the standardized residuals may reflect model misspecification, whereas serial correlation in their squares may signal the presence of remaining ARCH effects.

The estimation results associated with the baseline UC model with no structural breaks are in Table A1, and its corresponding serial correlation test results are reported in column 2 of Table A2. As shown, there is remaining serial correlation in the squares of the standardized residuals at the 10 percent level, signifying remaining ARCH effects. Therefore, the no-break model is misspecified.

Table A1: Estimation results from the UC OE-NKPC with no structural breaks [1993Q1-2015Q1]

Phillips Curve Parameters	
ψ	0.444***(0.109)
k	0.021(0.047)
k^*	0.162(0.205)
σ_e	0.201(0.220)
σ_η	3.113***(0.258)
Output Parameters	
ϕ_1	0.810***(0.115)
ϕ_2	-0.039(0.115)
σ_v	2.473***(0.199)
ϕ_1^*	1.398***(0.090)
ϕ_2^*	-0.606***(0.090)
σ_v^*	0.505***(0.041)
Log-likelihood value: -292.440	

Note: Standard errors are in parentheses.
,,* denote statistical significance at the 1, 5, and 10 percent levels respectively.

Table A2: Tests of model misspecification

Standardized residuals			
Lag	No-break	One-break	Two-break
1	0.561	0.833	0.325
2	0.167	0.450	0.248
3	0.277	0.617	0.413
4	0.405	0.764	0.541
5	0.385	0.521	0.184
6	0.462	0.611	0.268
7	0.464	0.644	0.367
8	0.566	0.668	0.428
Square of standardized residuals			
1	0.083	0.424	0.244
2	0.063	0.345	0.485
3	0.133	0.494	0.586
4	0.230	0.662	0.687
5	0.320	0.762	0.785
6	0.424	0.830	0.825
7	0.528	0.891	0.862
8	0.573	0.865	0.872

Note: Reported are p-values associated with the Ljung-Box tests under the null of no serial correlation.

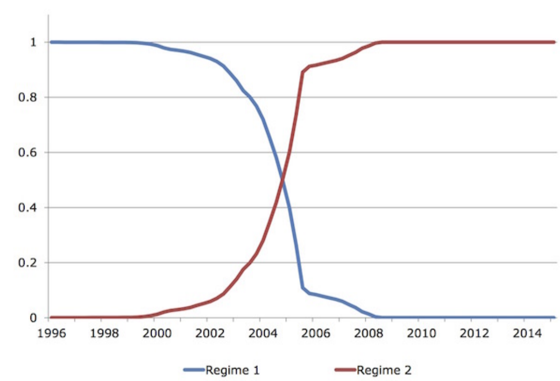
Table A3 reports the estimation results associated with the one-break UC model. According to the Ljung-box test statistics in column 3 of Table A2, the one-structural break model appears well-specified. However, by examining the smoothed probabilities of the UC model that is plotted in Figure A1, the transitional period between regimes occur over a long span of approximately 8 years, suggesting that the UC model may require an additional structural break. Finally, by estimating a UC model with two structural breaks, the model offers a good fit to the data. Not only does it yield sharp transitions between regimes, but the serial correlation tests in the last column of Table A2 also confirms that the model is well-specified. The estimation results are reported and discussed in the main body of the paper.

Table A3: Estimation of the UC OE-NKPC with one structural break [1993Q1-2015Q1]

Parameters	Regime 1	Regime 2
	Phillips Curve Parameters	
ψ	0.534***(0.178)	0.303*(0.182)
k	0.101(0.071)	-0.029(0.068)
k^*	-0.071(0.261)	0.404(0.335)
σ_e	0.405(0.329)	0.000(0.000)
σ_η	2.374***(0.340)	3.445***(0.414)
	Output Parameters	
ϕ_1	0.792***(0.113)	
ϕ_2	-0.016(0.110)	
ϕ_1^*	1.418***(0.089)	
ϕ_2^*	-0.636***(0.090)	
σ_v	2.474***(0.200)	
σ_v^*	0.505***(0.041)	
p_{11}	0.974***(0.026) $\rightarrow$ Break date: 2002Q2	
Log-likelihood value: -289.967		

Note: Standard errors are in parentheses. ***,**,* denote statistical significance at the 1, 5, and 10 percent levels respectively.

Figure A1: Smoothed Probabilities for the One-Break UC Model



Note: The figure shows smoothed probabilities associated with the one-structural UC-NKPC model. Smoothed probabilities are different from filtered probabilities in the sense that they are estimated based on incorporating information up until the end of the sample period.

Appendix B: A State-Space Representation of the UC OE-NKPC Model

The corresponding state-space representation for the UC OE-NKPC model with two structural breaks can be written as:

Measurement equation

$$
\begin{bmatrix} \pi_t \\ x_t \\ y_t^* \end{bmatrix} = \begin{bmatrix} 1 & 1 & 0 & 0 & 0 & 0 & 0 \\ 0 & 0 & 1 & 0 & 0 & 0 & 0 \\ 0 & 0 & 0 & 0 & 1 & 0 & 1 \end{bmatrix} \begin{bmatrix} \bar{\pi}_t \\ z_t \\ x_t \\ x_{t-1} \\ x_t^* \\ x_{t-1}^* \\ \tau_t^* \end{bmatrix} + \begin{bmatrix} k_{S_t} \sum_{j=0}^{\infty} E_{t-1}(x_{t+j}) + k_{S_t}^* \sum_{j=0}^{\infty} E_{t-1}(x_{t+j}^*) \\ 0 \\ 0 \end{bmatrix}
$$

Transition equation

$$
\begin{bmatrix} \bar{\pi}_t \\ z_t \\ x_t \\ x_{t-1} \\ x_t^* \\ x_{t-1}^* \\ \tau_t^* \end{bmatrix} = \begin{bmatrix} 0 \\ 0 \\ 0 \\ 0 \\ 0 \\ 0 \\ \delta_1^* D_{1t} + \delta_2^* D_{2t} + \delta_3^* D_{3t} \end{bmatrix} + \begin{bmatrix} 1 & 0 & 0 & 0 & 0 & 0 & 0 \\ 0 & \psi_{S_t} & 0 & 0 & 0 & 0 & 0 \\ 0 & 0 & \phi_1 & \phi_2 & 0 & 0 & 0 \\ 0 & 0 & 1 & 0 & 0 & 0 & 0 \\ 0 & 0 & 0 & 0 & \phi_1^* & \phi_2^* & 0 \\ 0 & 0 & 0 & 0 & 1 & 0 & 0 \\ 0 & 0 & 0 & 0 & 0 & 0 & 1 \end{bmatrix} \begin{bmatrix} \tilde{\pi}_{t-1} \\ z_{t-1} \\ x_{t-1} \\ x_{t-2} \\ x_{t-1}^* \\ x_{t-2}^* \\ \tau_{t-1}^* \end{bmatrix}
$$

$$
+ \begin{bmatrix} 1 & 0 & 0 & 0 & 0 \\ 0 & 1 & 0 & 0 & 0 \\ 0 & 0 & 1 & 0 & 0 \\ 0 & 0 & 0 & 0 & 0 \\ 0 & 0 & 0 & 1 & 0 \\ 0 & 0 & 0 & 0 & 0 \\ 0 & 0 & 0 & 0 & 1 \end{bmatrix} \begin{bmatrix} e_t \\ \eta_t \\ v_t \\ v_t^* \\ w_t^* \end{bmatrix}
$$

$$\begin{bmatrix} e_t \\ \eta_t \\ v_t \\ v_t^* \\ w_t^* \end{bmatrix} \sim i.i.d.N\left(\begin{bmatrix} 0 \\ 0 \\ 0 \\ 0 \\ 0 \end{bmatrix}, \begin{bmatrix} \sigma^2_{e,S_t} & 0 & 0 & 0 & 0 \\ 0 & \sigma^2_{\eta,S_t} & \sigma_{\eta,S_t}\sigma_v & \sigma_{\eta,S_t}\sigma_{v^*} & 0 \\ 0 & \sigma_{\eta,S_t}\sigma_v & \sigma^2_v & 0 & 0 \\ 0 & \sigma_{\eta,S_t}\sigma_{v^*} & 0 & \sigma^2_{v^*} & \sigma_{v^*}\sigma_{w^*} \\ 0 & 0 & 0 & \sigma_{v^*}\sigma_{w^*} & \sigma^2_{w^*} \end{bmatrix}\right),$$

$$D_{1t} = \begin{cases} 1, & \text{if } 1993Q1 \leq t < 1997Q3 \\ 0, & \text{otherwise} \end{cases}$$

$$D_{2t} = \begin{cases} 1, & \text{if } 1997Q3 \leq t < 2007Q4, \\ 0, & \text{otherwise} \end{cases}$$

$$D_{3t} = \begin{cases} 1, & \text{if } t \geq 2007Q4, \\ 0, & \text{otherwise.} \end{cases}$$

Based on the assumption that the domestic output gap follows an AR(2) process, the infinite sum term $\sum_{j=0}^{\infty} E_{t-1}(x_{t+j})$ in the inflation equation can be computed as:

$$\sum_{j=0}^{\infty} E_{t-1}(x_{t+j}) = e_1' F (I_2 - F)^{-1} \tilde{X}_{t-1}$$

where $e_1 = \begin{bmatrix} 1 \\ 0 \end{bmatrix}$, $F = \begin{bmatrix} \phi_1 & \phi_2 \\ 1 & 0 \end{bmatrix}$ and $\tilde{X}_{t-1} = \begin{bmatrix} x_{t-1} \\ x_{t-2} \end{bmatrix}$.

Similarly, the expression for $\sum_{j=0}^{\infty} E_{t-1}(x^*_{t+j})$ can be written as:

$$\sum_{j=0}^{\infty} E_{t-1}(x^*_{t+j}) = e_1' F^* (I_2 - F^*)^{-1} \tilde{X}^*_{t-1}$$

where $F^* = \begin{bmatrix} \phi_1^* & \phi_2^* \\ 1 & 0 \end{bmatrix}$ and $\tilde{X}^*_{t-1} = \begin{bmatrix} x^*_{t-1} \\ x_{t-2} \end{bmatrix}$.