

Thai Inflation Dynamics in a Globalized Economy

Pym Manopimoke, Wanicha Direkudomsak
Bank of Thailand

November 2015*

Abstract: This paper investigates whether the observed changes in Thai inflation dynamics since the 1990s can be attributed to the process of globalization. First, we develop a dynamic factor model to extract a global factor that is responsible for inflation rate movements in Thailand and its top trading partners. Empirical findings suggest that the importance of this global factor for Thailand doubled since 2001, emphasizing the growing role of globalization since then. Second, we develop an unobserved components model for Thai inflation with structural breaks that is consistent with an Open Economy New Keynesian Phillips Curve to explore the economic determinants behind the global factor. Based on the estimation results, long-term inflation expectations in Thailand fell and became well-anchored at the Bank of Thailand's newly adopted inflation target since 2001. During the same time, Thai inflation also became increasingly dependent on the global output gap while becoming less sensitive to domestic resource capacity constraints. By separately studying the factors that may have had direct and indirect effects on the global output gap, we find that Thai inflation underwent another structural change in 2007. While the global output gap in the 2001-2007 period is influenced by many global factors, its movements since 2007 only reflect the direct effects of world oil prices on inflation.

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

JEL Classifications: E3, E5, F4.

*The authors are grateful to Piti Disyatat, Roong Poshyananda Mallikamas, Surach Tanboon, Jariya Preamsin and Tosapol Apaitan for their support and insightful comments. Special thanks also go to our colleagues in the Monetary Policy Group at the Bank of Thailand for useful advice and their assistance in providing relevant data. The views expressed in this paper do not necessarily reflect the views of the Bank of Thailand, and any errors in this paper are our own.

1 Introduction

Inflation dynamics in Thailand has undergone fundamental changes during recent decades. In particular, Thai inflation has remained 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 inflation in Thailand to the adoption of inflation targeting in 2001, which has been suggested to help keep inflation low and stable through well-anchored long-run inflation expectations.

Monetary policy however, primarily works through lowering long-run inflation trends. Therefore, the monetary policy explanation may not be able to fully account for changes in inflation that may occur over the short to medium-term. Furthermore, the observed shift towards low and stable inflation in Thailand is in fact a worldwide phenomena, thus it is highly likely that the recent changes in Thai inflation may stem from other factors aside from an improved monetary policy framework. For example, by the early 2000s, the level and variability of international inflation rates has undergone a significant decline, especially in advanced economies. A number of studies also report a fall in the degree of inflation persistence as well as exchange rate pass through (White, 2008). Based on factor analyses, Ciccarelli and Mojon (2010), Neely and Rapach (2011), and Manopimoke (2015) find a sizable global factor responsible for movements in international inflation rates, which explains why international inflation rates have become increasingly synchronized in recent years.

While we do not discount the fact that monetary policy has an important bearing for Thai inflation, especially in the long-run, this paper aims to investigate the extent in which changes in Thai inflation can also be explained by global factors. Based on a growing literature, 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 countries such as China and India into world trade systems 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). Globalization fueled growth may also strengthen global commodity price cycles, which ultimately lead to greater comovements of international inflation rates. However, a serious shortcoming in this literature is that investigations are mostly limited to the experience of advanced economies. With export

volumes in emerging Asian countries currently accounting for more than one-third of world trade flows, the impact of globalization on inflation in developing countries such as Thailand is a vital piece of evidence that is needed to gain a deeper understanding on the link between globalization and inflation.

The impact of globalization on inflation is often measured by the empirical relevance of the global output gap, defined as the deviation of world demand from world supply. Thus far, there is inconclusive evidence as to whether the global output gap matters for national price processes (see Ball, 2006; Borio and Filardo, 2007; Ihrig et al. 2007; Pain et al., 2008). This is mainly due to difficulties in the measurement of the global output gap, which is an unobserved variable. In this paper, we introduce a new methodology to measure the global output gap and examine its importance for Thai inflation. In doing so, we first develop a dynamic factor model to extract a common factor from international inflation rates which represents global inflation, and investigate its empirical relevance for Thai inflation over time. Then, to study whether the global factor is related to economic variables such as the global output gap, we build upon the approach of Kim et al. (2014) and develop an unobserved components (UC) model for inflation that is consistent with an open-economy New Keynesian Phillips curve (OE-NKPC). An important advantage of this approach is that it allows us to estimate the global output gap as a latent state variable, and provides us with estimates of the gap that is consistent with the OE-NKPC. In addition, the UC framework gives a trend-cycle decomposition of inflation that allows us to focus on the role of the global output gap for Thai inflation in the short-run. Finally, to identify the channels in which the global output gap is important for Thai inflation, we also augment the UC model to separately account for the role of external factors, such as changes in commodity price dynamics and exchange rates.

A preview of our main empirical results are as follows. First, Thai inflation dynamics experienced two distinct structural changes over the 1993-2015 sample. The first regime shift occurred in 2001, while the second break took place during the onset of the global financial crisis in 2007. Second, we observe a significant shift in agents' long-term inflation expectations during the first regime change, which most likely stemmed from the Bank of Thailand's implementation of an inflation targeting framework in May 2000. Third, apart from changes in long-run trend inflation, we also find important changes to short-run inflation in 2001. In particular, the importance of domestic spare capacity constraints, as measured by the domestic output gap, declined while the global output gap became a prominent driving variable for Thai inflation since then. This finding implies that since the early 2000s, globalization has had important implications for short-run inflation dynamics Thailand. Last, we find that since 2007, the dominant driver behind the global output gap

became changes in world oil price movements, suggesting that the channels in which external forces transmit to Thai inflation since the global financial crisis has largely been through the direct import price channel.

This paper is organized as follows. Section 2 provides a description of recent changes in worldwide inflation, and describes the various channels in which globalization may affect inflation, particularly in Thailand. Section 3 sets up the dynamic factor model 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 for Thai inflation and provides a discussion of the empirical findings. Section 5 concludes and provides the monetary implications of our results.

2 The Global Dimension of Inflation

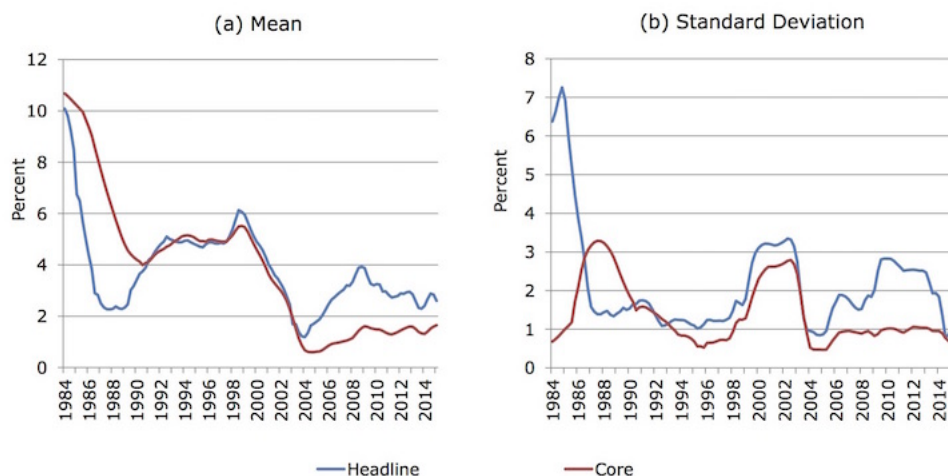
Inflation dynamics in Thailand has been remarkably low and stable since the early 2000s. As shown in Figure 1, the five year rolling average of Thailand's annual headline and core inflation have decelerated sharply since the early 2000s. Furthermore, despite the turmoil from the global financial crisis as well as the recent large swings in global commodity prices, Thailand's inflation rates have remained remarkably subdued. Headline inflation in Thailand has only been as high as 2.6 percent in the past decade, which is a welcomed contrast to the average of 5.4 percent prior to the year 2000. During recent periods, Thailand's inflation rates have fallen even further, with headline CPI inflation currently standing at -0.77 percent and core inflation at 0.95 percent, in large part due to the slide in global commodity prices.

The inflation experience of Thailand is not country-specific, but echoes the behavior of inflation rates around the world. According to Figure 2, the mean and volatility of inflation rates in advanced economies started to fall in the late 1980s, coinciding with the period of the Great Moderation. Emerging market economies then followed around the year 2000¹. With inflation rates around the world becoming more stable, the degree of co-movement across countries has also increased significantly.

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, current inflation rates in a number of

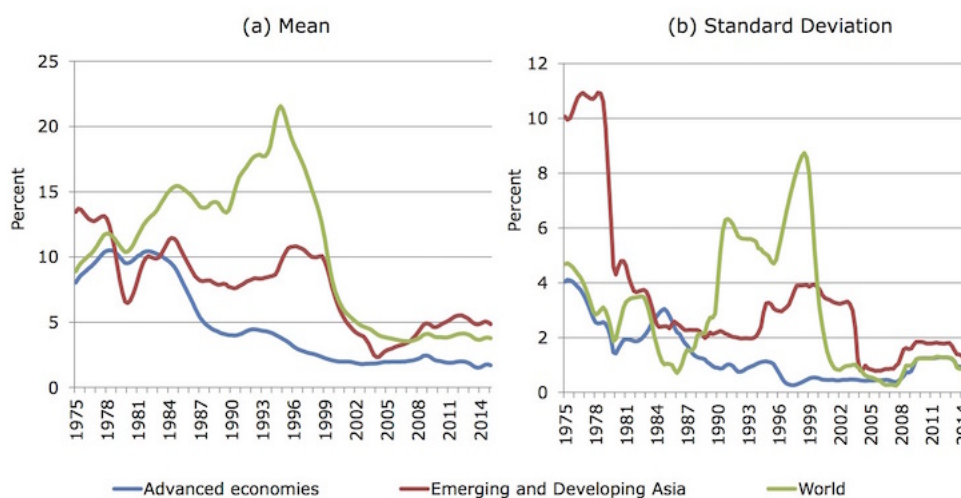
¹In Figure 2, world inflation rates are higher than inflation in advanced and Asian emerging countries prior to the year 2000 due to the exceptionally high bouts of inflation in Latin American countries during the debt crisis that struck the region in the 1980s. The average inflation rate in the most 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 dropped dramatically, and in most cases remained low in the single digits.

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, authors' calculations.

Figure 2: Worldwide Inflation Mean and Volatility

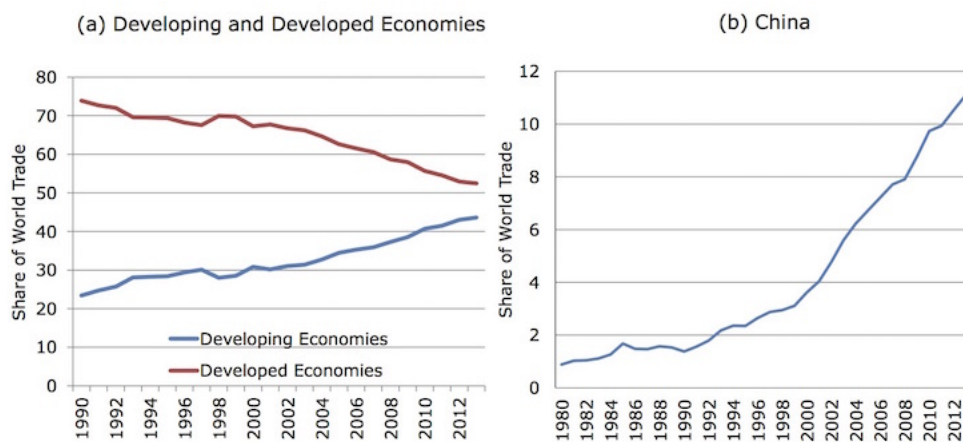


Note: Inflation is year-on-year changes in the aggregated headline consumer price indexes. The mean and standard deviations are computed using a five-year rolling window. The horizontal axis marks the date at the end of the rolling sample.
 Source: IMF International Financial Statistics Database, authors'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

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 empirical studies report a 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 dramatically declined, contributing to the fall in inflation persistence across a number of countries in past decades (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’. As depicted in Figure 3a, the emerging market’s share of world trade, defined as the sum of exports and imports from emerging markets to total world trade, accelerated in the year 2000. At the same time, the share of world trade from advanced economies has been on the decline, signifying enhanced integration of world trade systems. As shown in Figure 3b, China’s integration into international markets in 2000 has been a key driving force behind the acceleration in global trade, as evident by the steep rise in the country’s share of world trade.

Figure 3: Share of World Trade

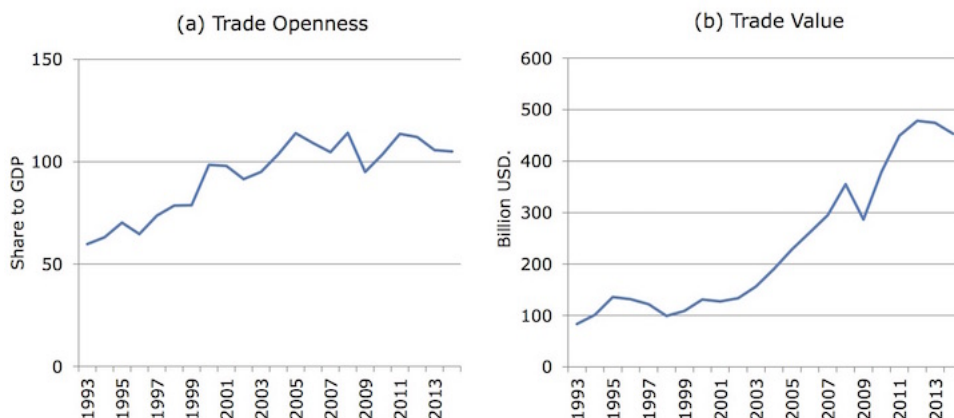


Note: Trade numbers are measured in US dollars at current prices and current exchange rates in millions.
Source: UNCTAD

Around the same time, Thailand has also become increasingly integrated into global trade systems. Figure 4a shows that the degree of trade openness for Thailand, measured as the country’s sum of imports and exports divided by gross domestic product, continued to rise since the 1990s. In Figure 4b, the trade value for Thailand’s trade accelerated sharply in

the early 2000s as well².

Figure 4: Trade Statistics for Thailand



Source: UNCTAD and IMF International Financial Statistics.

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 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)³.

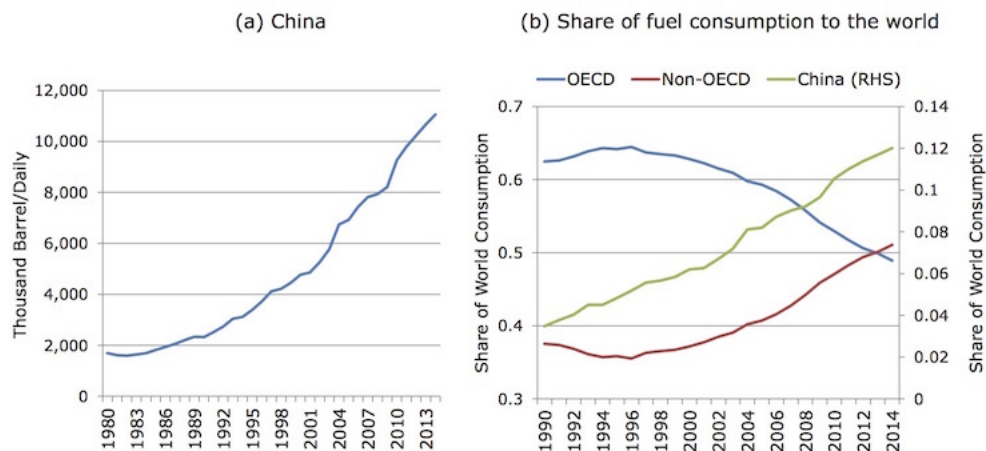
Note however, that globalization can provide offsetting effects to prices via the import price channel. Based on various estimation methodologies, a number of studies suggest that globalization-fueled growth from emerging countries turned out to be a significant driving factor in driving up world commodity prices (Adams and Ichino, 1995; Rae and Turner, 2001; Pain et al., 2008). A quick glance at Figure 5 reveals that fuel consumption in developing

²In Figure 5a, the acceleration in openness appears to start earlier in 1997 instead of the early 2000s due to the drop in GDP that occurred during the wake of the Asian Financial crisis.

³However, 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.

countries, especially China, had accelerated since 2000. By 2005, oil consumption from non-OECD economies accounted for approximately 40 percent of total global oil consumption with one fifth of this being due to China. Finally, apart from its direct effects on prices, note that world commodity price cycles that have been strengthened by globalization can also explain the greater degree of price synchronizations across countries.

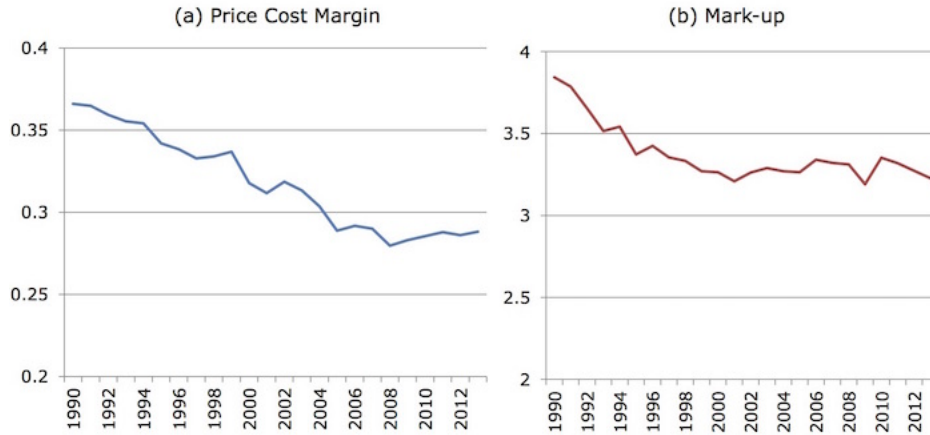
Figure 5: Fuel Consumption



Source: British Petroleum.

A second channel in 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). 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. Ever since Thailand became more open to trade in the 1990s, proxies of competition such as the price cost margin and mark-up implies that domestic competition has increased (see Figure 6), and may be responsible for the overall fall in Thailand's inflation rate which occurred during the same time.

Figure 6: Proxies of Competition in Thailand



Note: The price cost margin (PCM) is a widely used proxy for monopolistic mark-up. We calculate it as the difference between GDP and wage compensation, divided by total output. The mark-up measure is calculated as the ratio between GDP and the compensation of employees.

Source: NESDB, authors' calculations.

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 helping to keep 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 easily move through the supply chain.

3 Dynamic Factor Analysis for Thai Inflation

One way to examine the importance of globalization for worldwide inflation is to investigate whether there component that can capture common movements driving international inflation rates. Based on various statistical models, a number of authors find an important common component that drives the overall movements of inflation rates across countries. For a group of advanced economies, Ciccarelli and Mojon (2010) find that on average, a sin-

gle common factor can explain nearly 70 percent of the variance in national inflation rates. Based on a time-varying analysis, Manopimoke (2015) finds that for a large sample which represents worldwide inflation, the importance for the common factor reached as high as 90 percent during recent years.

To identify a common factor that is relevant for Thailand, we develop a dynamic factor model (DFM) which extracts a common component from the inflation rates of Thailand and its top trading partners. The top trading partners of Thailand include: Australia, Hong Kong, Japan, Korea, Malaysia, Philippines, Singapore, the US, Indonesia, UK, Taiwan, China, and the EU-18 region⁴, which accounts for an average of 75 and 79 percent of Thailand’s imports and exports respectively during the past decade. Due to this large share, we will henceforth refer to this estimated common component as a ‘global factor’.

The DFM for inflation is as follows:

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

where $\pi_{i,t}$ is the inflation rate series i , where $i = 14$.

The DFM decomposes the 14 inflation series into the following three components. First, there is a global component f_t^g , which captures movements in inflation rates that are shared across all countries 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 movements for each of the 14 inflation series that stem from within-country demand and supply shocks. The coefficients λ_i^g and λ_i^r are factor loadings for the global and regional factors respectively, and reflect the degree to which variation in $\pi_{i,t}$ can be explained by those corresponding factors. Note that during estimation, 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.

In specifying the dynamics of the three latent factors, each factor is assumed to follow an autoregressive process (AR) of order 2:

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

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

⁴Countries 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.

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

where for identification purposes, 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. Note that these set of assumptions are standard for these class of models.

To estimate the DFM, we use 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. We utilize the Kalman filter to estimate a total of 30 equations, which is comprised of 14 inflation equations as specified in Eq (1), 1 equation for the global factor that follows Eq (2), 1 equation for the regional factor as specified by Eq (3), and 14 equations for the country-specific factor as represented by 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 estimates of 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 loading factor for the Philippines ($\lambda^r = 1.11$). However, the loading factor on the global factor for Thailand is the highest in the group of countries ($\lambda^g = 0.69$). This is not surprising given that the subset of countries for the analysis is chosen specifically to reflect Thailand's trade structure.

In Figure 7 we plot the estimated factors from the DFM against actual inflation rates. As shown in Figure 7a, the global component tracks the average value of the 14 inflation series well, particularly since the year 2000. According to Figures 7b-c, the domestic and regional components moves closely prior to the year 2000, and can explain the majority of movements in Thai inflation. However, these relationships break down in 2000, as the global factor becomes the prominent driving variable for Thai inflation instead. The missing link between Thai inflation and the regional component after the year 2000, and the rise in importance of the global factor for Thai inflation suggests that since this time, Asia Pacific has become more integrated with the global economy. In turn, domestic inflation rates in Thailand become increasingly driven by worldwide economic conditions.

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

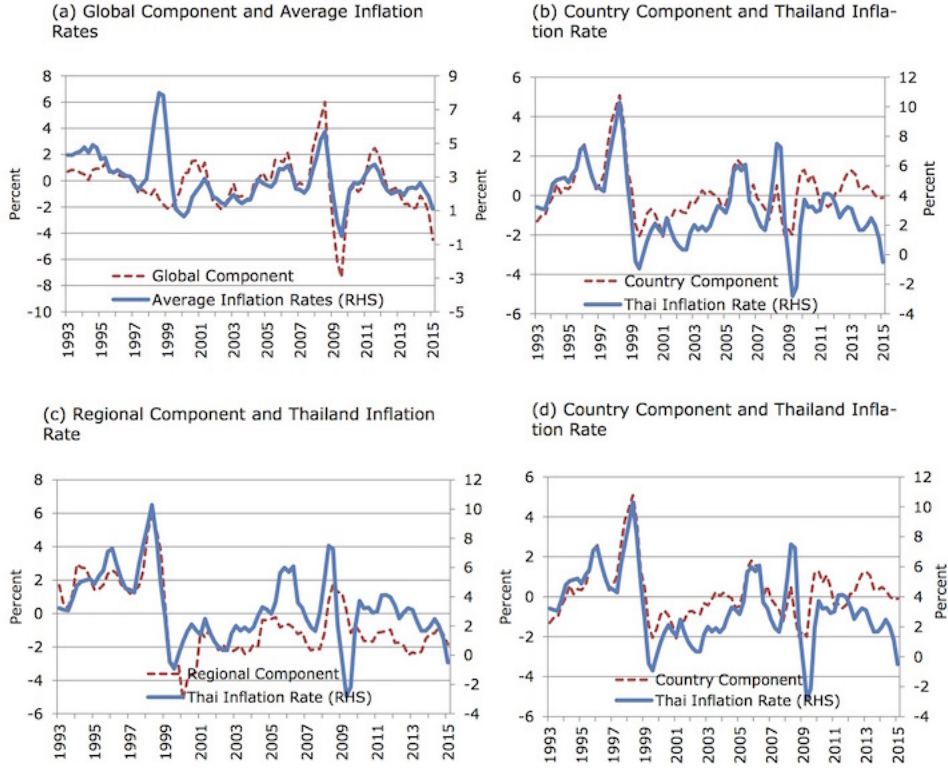
Parameters	AR(2) coefficients of global and regional components
γ_1^g	1.275***(0.108)
γ_2^g	-0.550***(0.102)
γ_1^r	1.355***(0.119)
γ_2^r	-0.475***(0.122)

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

Log-likelihood value: -462.391

Figure 7: Components of the Dynamic Factor Model



Sources: IMF International Financial Statistics, authors' calculations.

Given the shift towards a more prominent role for the global factor since the year 2000, a DFM that allows for one structural break in its factor loadings may be more appropriate. We therefore consider the following DFM that allows for one endogenously determined structural break in the factor loading coefficients:

$$\pi_{i,t} = \lambda_{i,S_t}^g f_t^g + \lambda_{i,S_t}^r f_t^r + z_{i,t}, \quad (1')$$

$$f_t^c = \gamma_{S_t}^g f_{t-1}^g + \eta_t^g, \quad \eta_t^g \sim N(0, 1), \quad (2')$$

$$f_t^r = \gamma_{S_t}^r f_{t-1}^r + \eta_t^r, \quad \eta_t^r \sim N(0, 1), \quad (3')$$

$$z_{i,t} = \gamma_{i,S_t}^z z_{i,t-1} + \eta_{i,t}^z, \quad \eta_{i,t}^z \sim N(0, \sigma_{i,S_t}^{z2}), \quad (4')$$

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

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

with the (i, j) - *th* element referring to $Pr[S_t = j | S_{t-1} = i]$. Note that in contrast to the

previous no-break DFM specification, in which we also refer to as the baseline model, we assume that the dynamic factors in the one structural break model has AR(1) instead of AR(2) dynamics to reduce the number of parameters to be estimated. Doing so does not affect the estimation results, because as reported in Table 1, the majority of the AR(2) coefficients in the DFM is not statistically significant.

The estimation results from the one-break DFM model is shown in Table 2. As expected, the model estimates a structural break in 2001Q1, where we see a distinct shift in the factor loadings on the regional and global component. A quick glance at the results show that prior to the break date, the regional and global factor share approximately equal weight in explaining the overall movements in Thai inflation rates. However, after the break, the regional factor is no longer statistically significant while the factor loading on the global component almost doubles. Compared to the no-structural break DFM, these findings give more concrete evidence that the global factor is an important driving factor for Thai inflation dynamics, most likely due to the acceleration in globalization that occurred during this time.

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

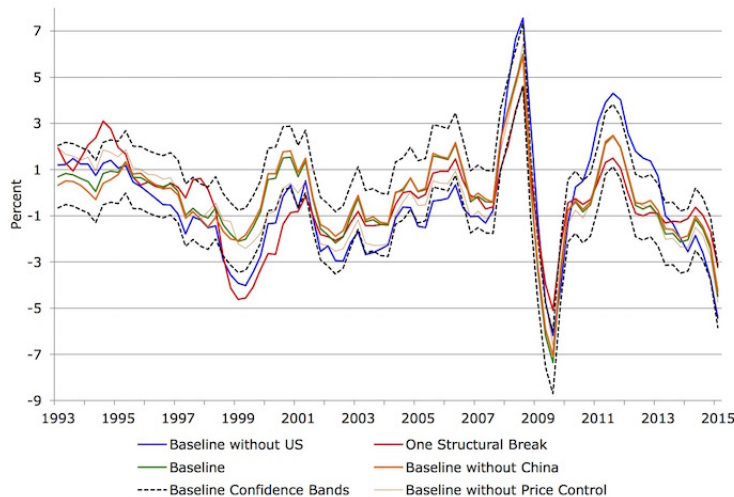
Parameters	Regime 1	Regime 2
λ^g	0.557**(0.223)	1.108***(0.137)
λ^r	0.644***(0.193)	0.148(0.107)
σ^z	0.768***(0.116)	0.598***(0.079)
γ^z	0.874***(0.078)	0.758***(0.108)
Break date	0.976**(0.023) \rightarrow 2001Q1	

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

We plot the global factor that belongs to the one structural break model in Figure 8 alongside the 95 percent confidence bands associated with the global factor from the baseline specification. As shown, the global factor from the one-break model are well contained within the confidence bands, implying that the estimated global factors from the two models are not statistically different. For robustness checks, we also plot estimated global factors from three additional models. First, previous studies have shown that the US economy leads the rest of the world and that US shocks are transmitted to other countries to a great extent (Canova and Marrinan 1998; Artis et al. 2007; Dees et al. 2007). To check whether the global component may be dominated by US inflation movements, we reestimate the DFM without the US price series. Second, for the same reasoning, we repeat the same analysis

but we exclude the price series of China. Last, we exclude the price series of Thailand, Indonesia and the Philippines in the analysis as these countries implement relatively heavy price control policies, which as a result may distort our analysis. Based on these robustness tests, with the exclusion of a brief period in the early 2000 period where we have identified a structural change, all estimated global factors from the robustness checks lies within the 95 percent confidence bands of the baseline DFM specification.

Figure 8: Global Components from Various Dynamic Factor Models [1993Q1-2015Q1]

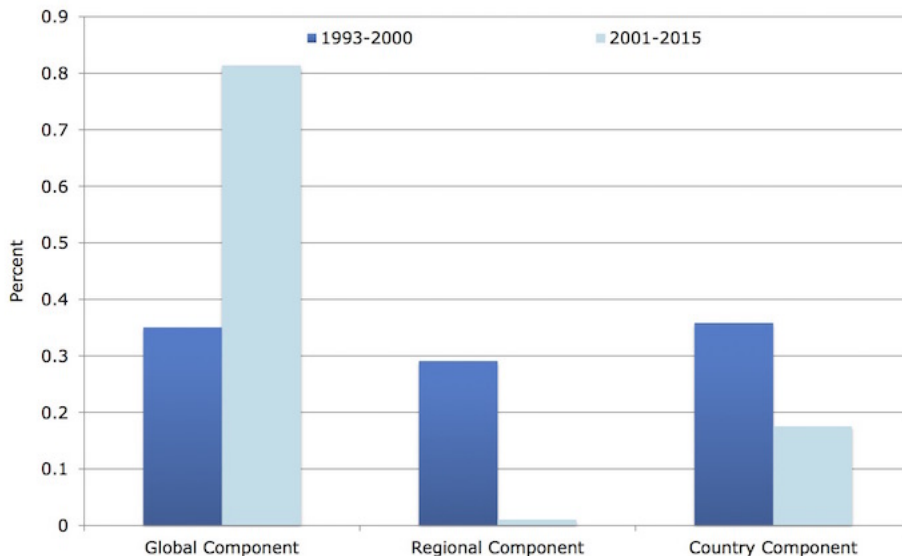


Sources: IMF International Financial Statistics, authors' calculations.

Finally, we perform a variance decomposition to measure the relative importance of the global, regional and country-specific factors in explaining the overall variability in Thai inflation rates⁵. The variance decomposition results are based on the one structural break DFM, and the findings are reported in Figure 9. According to the results, we observe that in the pre 2000 period, all three components play an approximately equal role in explaining inflation variability in Thailand, but since 2001, we see that the importance of the global factor towards explaining inflation volatility has increased significantly, while the significance of country-specific factor falls below 20 percent. Furthermore, the role of the regional component in the post 2001 period is near zero, which suggests that movements in Thai inflation since then had expanded beyond the effects of shocks originating from regional markets, owing to the increased pace of globalization occurring during that time.

⁵Under the assumption that the components are orthogonal, it is straight forward to decompose inflation variance into three parts. The share of inflation variance explained by the three components in each regime can be computed as: $S_i^w = \frac{\lambda_i^{c2}}{Var(\pi_i)}$, $S_i^r = \frac{\lambda_i^{r2}}{Var(\pi_i)}$, $S_i^z = \frac{\sigma_i^{z2}}{Var(\pi_i)}$ 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 $Var(f^c) = Var(f^r) = 1$.

Figure 9: Variance Decomposition [1993Q1-2015Q1]



Note: Estimation results are based on the DFM with one structural break.
Sources: IMF International Financial Statistics, authors' calculations.

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

In the previous section, we find that there is a sizable global factor underlying Thai inflation dynamics. Existing studies typically capture global influences through a global output gap, defined as the difference between world demand world supply. More specifically, the global output gap is the difference between actual output and potential output at the global level, and is a measure of tightness or slack in the use of global resources. To examine the relationship between Thai inflation and the global output gap, we consider estimation of an open economy New Keynesian Phillips curve (OE-NKPC). The appeal of the OE-NKPC is that it is derived from a general equilibrium framework based on optimizing behavior of monopolistically competitive firms (see Clarida, Gali and Gertler, 2002; Corsetti and Pesenti, 2005), giving the model solid microfoundations.

The OE-NKPC can be written as:

$$\pi_t = \beta E_t(\pi_{t+1}) + kx_t + k^*x_t^* + \Gamma_t \quad (5)$$

where π_t is the current inflation rate; β is the subjective discount factor, $E_t(\cdot)$ denotes expectations formed conditional on information up to time t ; x_t is the domestic output gap which is a function of firm's marginal costs; and x_t^* is the global output gap. The

coefficients on the output gaps, k and k^* are functions of the deep structural parameters of the model 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 may capture remaining shocks to inflation that are not captured by the output gap which may represent international competitiveness measures such as the terms of trade, the deviation from the law of one price of import prices, and the deviations from purchasing power parity of the real exchange rate.⁶

The distinguishing factor of the above open-economy from its closed-economy counterpart is the presence of the global output gap variable x_t^* . In an open economy, firms can export their goods to a foreign country. These cross-border pricing decisions introduces additional dependence between current inflation and marginal costs of the exporting firms in the other country, and as a consequence, on a measure of the global output gap⁷. In studying the role of globalization for inflation in the Phillips curve framework, it has been conjectured that as an economy becomes more internationalized in goods and financial markets, national inflation rates should become more sensitive to global factors rather than traditional domestic determinants. In other words, globalization should reduce the importance of the domestic output gap in Eq. (5) while increasing the prominence of the global slack measure⁸.

While the globalization hypothesis has gained much interest among policymakers and academics alike, empirical support for this proposition has been far from robust. For example, based on different sample periods, selection of countries, and various specifications for the open economy NKPC, Gamber and Hung (2001), Borio and Filardo (2007), and Manopimoke

⁶In the open economy NKPC, Γ_t the specific form of the NKPC depends on the underlying assumptions of the model, such as whether the exporting firms engage in local or producer currency pricing. For example, the 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 open economy NKPC under different assumptions can be found in the works of Clarida, Gali and Gertler (2002), Corsetti and Pesenti (2005), Steinsson (2005), Martínez-García and Wynne (2010), and Zaniboni (2011), among others.

⁷Note that the domestic output gap does already contain some information about global influences on Thailand costs and prices, as net exports are measured in Thailand's real GDP. Therefore, strong demand from abroad for goods and services is captured in the measure of Thailand's domestic output gap. Nevertheless, the global output gap can capture additional global influences such as the rising cost pressures in foreign economies that may put upward pressure on import prices, the amount of spare capacity overseas that may result in the weak bargaining power of domestic workers through more integrated labor markets, or the restraint on markups for domestic producers that result from enhanced competition in a more globalized economy.

⁸Theoretically, 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.

(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 consumer price inflation in a number of countries.

One reason for the conflicting results in the literature regarding the empirical relevance of the global output gap is the difficulty in the appropriate handling of the expectational element $E_t\pi_{t+1}$. In particular, how one proxies for one-period-ahead inflation expectations is most likely to influence the statistical significance of the output gap driving variables. For example, Borio and Filardo (2007) use HP-filtered inflation series as a proxy for the underlying long-run trend rate of inflation, which leaves ample variability and persistence in the remaining short-run movements of inflation to be explained by the domestic and global output gaps. However, as Ihrig et al. (2007) point out, this approach leaves autocorrelation in the residuals, which is a form of model misspecification. By proxying the expectational element with lagged values of actual inflation, these authors in turn leaves very little persistence in the inflation gap to be explained by the output gaps. As a result, they find no significant role for the global output gap, in contrast to Borio and Filardo (2007).

To avoid dealing with the inflation expectation element in Eq. (5), we follow the approach of Kim et al. (2014). This approach involves forward iteration of the OE-NKPC, yielding the following closed-form specification:

$$\pi_t = \lim_{j \rightarrow \infty} \beta^j E_t(\pi_{t+j}) + k \sum_{j=0}^{\infty} \beta^j E_t(x_{t+j}) + k^* \sum_{j=0}^{\infty} \beta^j E_t(x_{t+j}^*) + \tilde{z}_t, \quad (6)$$

where $\tilde{z}_t = \sum_{j=0}^{\infty} E_t(\Gamma_{t+j})$, and may be serially correlated. The first element on the right-hand-side of the above equation represents long-term inflation expectations and vanishes under the assumption of stationary inflation which fluctuates about a zero long-run trend. However, based on various unit root tests for inflation, we cannot reject the null hypothesis of no unit root for Thailand inflation. As emphasized in the literature, a unit root in inflation implies a time-varying inflation trend, and accounting for this time variation is critical towards estimation of the NKPC (see Cogley and Sbordone, 2008; Kim et al., 2014). Therefore, we follow Kim et al. (2014) and approximate $\lim_{j \rightarrow \infty} \beta^j E_t(\pi_{t+j})$ by a driftless random walk, and interpret this term as a Beveridge-Nelson stochastic trend (Beveridge and Nelson, 1981). Note that the remaining three terms on the right-hand-side of Eq (6) are components of a inflation gap, defined as deviations of actual inflation from a time-varying long-run trend, and captures short-run movements in inflation at business cycle frequencies.

Kim et al. (2014) estimates a closed-form NKPC for US inflation as an unobserved components (UC) model which effectively decomposes inflation into trend and cycle compo-

nents⁹. In this paper, we build an empirical model for Eq. (6) in the spirit of Kim et al. (2014), but we extend their framework to an open economy case. This yields the following baseline UC model for inflation:

Baseline specification

$$\pi_t = \bar{\pi}_t + k \sum_{j=0}^{\infty} E_{t-1}(x_{t+j}) + k^* \sum_{j=0}^{\infty} E_{t-1}(x_{t+j}^*) + z_t, \quad (7)$$

$$\bar{\pi}_t = \bar{\pi}_{t-1} + e_t, \quad (8)$$

$$z_t = \psi z_{t-1} + \eta_t, \quad (9)$$

$$x_t = \phi_1 x_{t-1} + \phi_2 x_{t-2} + v_t, \quad (10)$$

In the above specification, trend inflation follows a driftless random walk and the discount factor is calibrated to one¹⁰. To capture the cyclical nature of the output gaps, x_t and x_t^* are allowed to have AR(2) dynamics. z_t captures the influence of Γ_t and may be potentially serially correlated, and therefore we allow it to follow an AR(1) process. Finally, for feasible estimation of the model, it can be observed that in Eq. (7), we can replace the terms $\sum_{j=0}^{\infty} E_t(x_{t+j})$ and $\sum_{j=0}^{\infty} E_t(x_{t+j}^*)$ in Eq. (6) with $\sum_{j=0}^{\infty} E_{t-1}(x_{t+j})$ and $\sum_{j=0}^{\infty} E_{t-1}(x_{t+j}^*)$, giving $z_t = k(\sum_{j=0}^{\infty} E_t(x_{t+j}) - \sum_{j=0}^{\infty} E_{t-1}(x_{t+j})) + k^*(\sum_{j=0}^{\infty} E_t(x_{t+j}^*) - \sum_{j=0}^{\infty} E_{t-1}(x_{t+j}^*)) + \tilde{z}_t$, so that the infinite sum terms are no longer correlated with z_t . However, z_t is potentially correlated with x_t and x_t^* thus we allow $Cov(\eta_t, v_t) \neq 0$ and $Cov(\eta_t, v_t^*) \neq 0$.

The data set used to estimate the model above is based on data availability. We use quarterly data that spans 1993Q1-2015Q1, and inflation is the quarterly changes in the seasonally-adjusted consumer price index (CPI). For the domestic output gap, x_t is proxied by the Bank of Thailand's (BOT) measure of the domestic gap which is obtained from a multivariate model for inflation, output and interest rates¹¹. Since we do not have a measure of the global output gap x_t^* , we treat x_t^* as an latent state variable and extend the UC model for inflation with the following UC model for foreign output:

$$y_t^* = \tau_t^* + x_t^*, \quad (11)$$

⁹Typically, trend-cycle decompositions for inflation are statistical in nature, such as the univariate UC-SV model of Stock and Watson (2007). The decomposition approach of Kim et al. (2014) has more economic structure and content, yet maintains the flexibility of the statistical approach of Stock and Watson.

¹⁰The 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.

¹¹Results are qualitatively the same using HP gap. We also experimented with unit labor cost measures but labor data for Thailand does not extend past 2000.

$$\tau_t^* = \delta_1^* D_{1t} + \delta_2^* D_{2t} + \delta_3^* D_{3t} + \tau_{t-1}^* + w_t^*, \quad w_t^* \sim i.i.d.N(0, \sigma_w^{*2}), \quad (12)$$

$$x_t^* = \phi_1^* x_{t-1}^* + \phi_2^* x_{t-2}^* + v_t^*, \quad v_t^* \sim i.i.d.N(0, \sigma_v^{*2}), \quad (13)$$

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

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

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

such that the global output gap can be estimated based on observed output data of y_t^* within the UC-NKPC framework¹². In the above specification, the series y_t^* is an aggregated real output series of Thailand's top trading partners, constructed by weighing each country's PPP-adjusted GDP by its trade share with Thailand¹³. The set of countries selected as Thailand's trading partners is the same as those that were selected for estimation of the DFM in the previous section.

According to the specification above, y_t^* is decomposed into a trend component τ_t^* and a global output gap component x_t^* with movements that are assumed to follow a random walk with drift, and an AR(2) process respectively. For the trend output component, δ^* is the non-constant trend output growth rate. By incorporating the dummy variables, D_{1t} , D_{2t} and D_{3t} , δ^* is allowed to undergo two known structural breaks in 1997Q3 and 2007Q4 to capture the slowdown in RGDP growth during the Asian financial crisis as well as during the most recent global financial crisis respectively. Note that by doing so, it is not imposing a structural break, but merely allowing one to happen. Readers are referred to Appendix A for the state-space representation of the model as described in Eqs (7)-(13).

¹²Note that technically, estimates of x_t^* are considered as a measure of the foreign output gap that is relevant for Thailand. However, with Thailand being a small open economy, there should be negligible differences between foreign and global output gap measures. Therefore, we refer to x_t^* as the global output gap. Also, in obtaining an estimate of the global output gap, we could alternatively apply the HP-filter to y_t^* . However, it is well known that the HP-filter can produce filtered series that can be subject to end point problems. In addition, the HP-filtered gap is a purely statistical measure of the output gap, while we prefer a measure that has more economic content. By estimating y_t^* within the UC framework, we obtain estimates for the global output gap that are consistent with an OE-NKPC.

¹³We also tried weighing the countries by GDP share, but the results were robust to this alternative specification. Results are available upon request.

4.1 Incorporating Structural Breaks

In bringing the baseline UC specification to the data, the preliminary DFM analysis in Section 3 strongly suggests that there may have been structural changes in the relationship between inflation and its driving variables. To allow for this possibility, we incorporate structural breaks into the baseline specification as follows:

Structural Break Specification

$$\pi_t = \bar{\pi}_t + k_{S_t} \sum_{j=0}^{\infty} E_{t-1}(x_{t+j}) + k_{S_t}^* \sum_{j=0}^{\infty} E_{t-1}(x_{t+j}^*) + z_t, \quad (7')$$

$$\bar{\pi}_t = \bar{\pi}_{t-1} + e_t, \quad e_t | S_t \sim i.i.d.N(0, \sigma_{e,S_t}^2), \quad (8')$$

$$z_t = \psi_{S_t} z_{t-1} + \eta_t, \quad \eta_t | S_t \sim i.i.d.N(0, \sigma_{\eta,S_t}^2), \quad (9')$$

$$x_t = \phi_1 x_{t-1} + \phi_2 x_{t-2} + v_t, \quad v_t \sim i.i.d.N(0, \sigma_v^2), \quad (10')$$

where the equations for the UC model for foreign output follows Eqs. (11)-(13) exactly.

In deciding upon the number of structural breaks S_t , since we cannot directly apply structural break tests to the UC specification, we estimate the model with incremental number of structural breaks, and apply serial correlation tests to examine whether there is any model mis-specification. In particular, we start with a model with no structural breaks, and compute the p-values of the Ljung-Box test statistics for both the standardized residuals and square of standardized residuals of the inflation series. The underlying idea is that if we are able to reject the null of no serial correlation in the standardized residuals, the model may be misspecified as there is remaining serial correlation in the error terms left to be explained. As for the square of the standardized residuals, should there remain any serial correlation, there may be remaining ARCH effects in the inflation series that needs to be explained. If the no structural break model is mis-specified, we incorporate one structural break in the model, and repeat the serial correlation tests. The process is repeated until we find the best-fitting model. Due to space considerations, we report the results of this exercise in Appendix B.

Based on the estimation results, a two structural break model is most appropriate for Thai inflation dynamics. Therefore, $S_t = 1, 2$ 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 $Pr[S_t = j | S_{t-1} = i]$.

4.2 Empirical Results

The results from the two structural break UC model is reported in Table 3. In the first column, we report the estimation results based on using the HP-filtered global output gap to proxy for x_t^* , which is the common output gap proxy in the existing literature. We compare these results against the main findings reported in the second column, which treats x_t^* as a latent state variable in our UC model with two structural breaks that is consistent with the OE-NKPC.

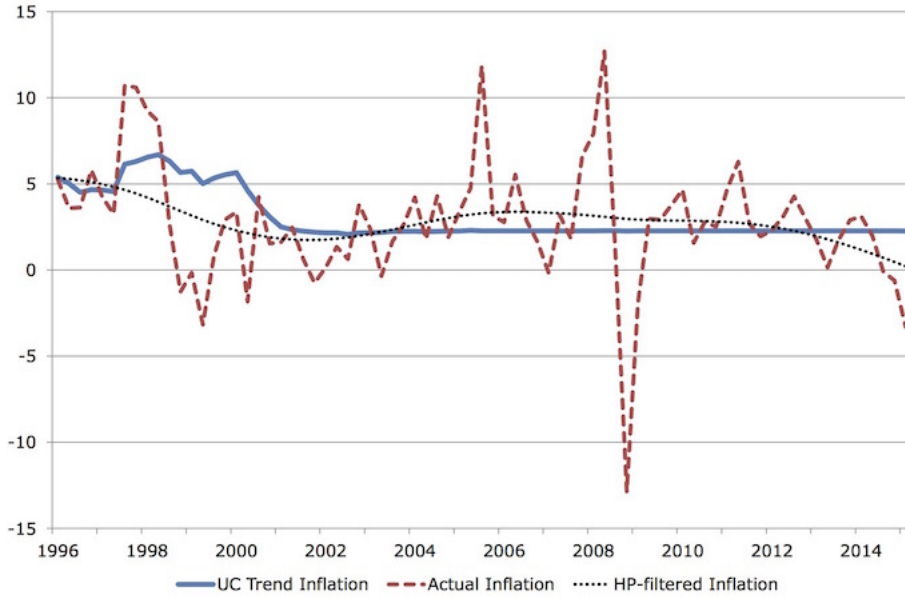
The key findings are as follows. First, the estimation results suggest that there have been two distinct structural changes that occurred in 2001 and 2007, which divide Thai inflation dynamics into three regimes; 1993-2001, 2001-2007, and 2007-2015. We examine the parameter estimates, one at a time, across the three regimes to examine the changes that are responsible for these regime shifts. First, although there appears to be no changes in the variability of shocks to trend inflation (σ_e), estimates of trend inflation from the UC model in Figure 10 shows that there have been a significant shift in the level of trend inflation in 2001. More specifically, while the movements in trend inflation has varied somewhat prior to 2001, since then, long-term inflation expectations has been well-anchored at an average level of 2.4 percent¹⁴. Note that in contrast to HP-filtered estimates of headline and core CPI inflation, trend inflation estimates from the UC model shows a more distinct drop in the early 2000 period, and is less volatile in the period thereafter. Accordingly, it paints a better picture in support of the view that the shift in monetary policy towards an inflation targeting regime that occurred in May 2000 helped lowered and anchored long-term inflation expectations.

¹⁴The Bank of Thailand’s implementation of an inflation targeting regime started in May 2000, where the Bank targeted core inflation with a target 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.

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

Parameters	UC with HP-filtered gap	UC model with estimated gap
Phillips Curve Parameters		
ψ_1	0.851***(0.098)	0.861***(0.094)
ψ_2	-0.653***(0.130)	-0.643***(0.122)
ψ_3	0.281(0.178)	0.260(0.166)
k_1	0.163*(0.092)	0.178*(0.098)
k_2	0.144*(0.083)	0.074(0.048)
k_3	-0.037(0.067)	-0.050(0.063)
k_1^*	-0.421(0.316)	-0.425(0.266)
k_2^*	0.397***(0.131)	0.321***(0.104)
k_3^*	0.434(0.333)	0.389*(0.236)
$\sigma_{e,1}$	0.001(0.221)	0.001(0.000)
$\sigma_{e,2}$	0.000(0.110)	0.000(0.000)
$\sigma_{e,3}$	0.000(0.211)	0.000(0.000)
$\sigma_{\eta,1}$	2.558***(0.403)	2.553***(0.401)
$\sigma_{\eta,2}$	0.927***(0.159)	0.871***(0.146)
$\sigma_{\eta,3}$	3.526***(0.394)	3.521***(0.396)
Output Parameters		
ϕ_1	0.784***(0.112)	0.765***(0.112)
ϕ_2	-0.005(0.108)	0.017(0.105)
ϕ_1^*	1.404***(0.078)	1.556***(0.090)
ϕ_2^*	-0.615***(0.063)	-0.701***(0.069)
σ_v	2.475***(0.120)	2.478***(0.200)
σ_v^*	0.505***(0.041)	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}	0.973***(0.030)	0.970***(0.032)
Break Date	2002Q1	2001Q1
p_{22}	0.943***(0.055)	0.957***(0.046)
Break Date	2006Q3	2007Q1
Log-likelihood value:	-280.818	-283.618

Figure 10: Actual Inflation and Trend Inflation



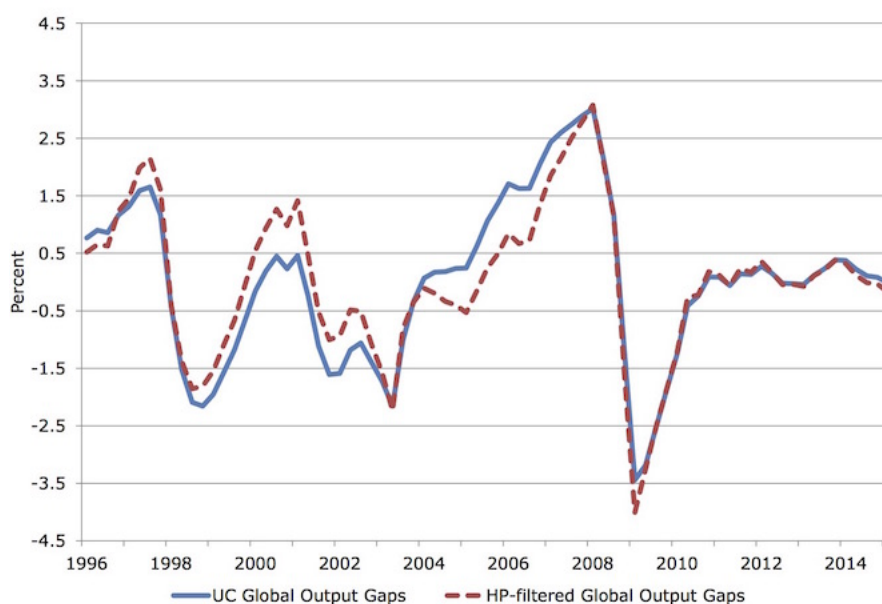
The marked shift in trend inflation in 2001 corresponds to a period of reduced inflation gap persistence, as measured by the magnitude of the ψ parameter. In 2007, Thai inflation gap persistence is reduced even further, implying that headline inflation reverts faster to its trend following temporary shocks to inflation. In part, these results provide further support for the shift towards improved monetary policy since early 2000, as better anchored inflation expectations can help reduce inflation persistence. As for the variability of shocks to the inflation gap z_t , estimates of σ_η indicate that while the volatility of shocks decline after the first break, it approximately quadruples after the second break. The large swings in σ_η most likely captures volatile movements in import prices, which appear to have important bearings for Thai inflation. However, as indicated by the reduced persistence of z_t , although these import price swings can be volatile, the effects of that they have on inflation are becoming shorter lived over time.

Our attention is focused particularly on the estimates of the output gap coefficients, k and k^* , and their evolution over time. In line with the experiences of advanced economies, we find that the sensitivity of inflation to the domestic output gap declines. In the UC model that employs the HP-filtered global gap, the decline does not come until 2007. However, the bivariate UC model with an unobserved global gap suggests that the decline happened earlier since 2001. The latter result is in line with the findings of Chantanahom et al. (2004) where they 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, declined after the Asian financial crisis.

Interestingly, the reduced sensitivity of inflation to domestic slack conditions occurred at a time when the link between inflation and the global output gap increased significantly, from an estimate of -0.4 which is insignificant and of an incorrect sign in the pre 2001 period, to a statistically significant estimate of 0.3 in the period thereafter. After 2007, the coefficient in front of the global output gap still remains large, but is not statistically significant for the model based on the HP-filtered global gap. Overall, these findings highlight the influence that globalization may have for domestic inflation dynamics. As Thailand becomes increasingly open and integrated with its trading partners, prices in the country become increasingly sensitive to global slack measures, and less so with domestic economic conditions.

Next, we examine the dynamics of the output gap parameters. The sum of the AR(2) coefficients of the domestic output gap is 0.7, suggesting that Thailand's output gap is a highly persistent process. Similarly, the HP-filtered and UC global gap measures are approximately as persistent, but shocks to the global gaps are less volatile when compared to the variability of shocks to Thailand's output gap. According to Figure 11, the UC global gap is able to capture the same peaks and troughs as the HP-filtered global gap. In general, both series comove closely, especially after the global financial crisis.

Figure 11: UC and HP-filtered global output gaps



Last, we find that the permanent shocks to the trend component of the aggregated real output series of Thailand's top trading partners is not statistically significant from zero. However, its trend output growth rate is significant and has been on the decline since the beginning of the sample. After the Asian financial crisis, potential output growth of the global output series declined only slightly, from 3.9 to 3.8 percent, while a significant decline

occurs after the global financial crisis where trend output growth rate drops to a level of 2.24 percent.

4.3 The Role of External Factors and the Global Output Gap

In this section, we augment the UC-NKPC with two structural breaks to explicitly account for the influences of external factors in the z_t process. Accounting for such factors may be important if the global output gap is not a sufficient summary statistic for capturing the effects of global influences on inflation. Furthermore, we might find that the global output gap is merely important in the previous specification because it is capturing the hidden influence of these omitted variables.

To examine the role of external factors, we consider adding the change in term of trade shocks, import prices, oil prices, non-oil commodity prices, and the real exchange rate, one-at-a-time and also as combinations, to the UC specification in Eqs. (7')-(10') and (11)-(13). However, we modify the z_t process in Eq. (9) to account for the influence of external factors Γ_t explicitly as follows:

$$z_t = \alpha_{S_t} \Gamma_{t-1} + \eta_t, \quad (14)$$

$$\eta_t = \psi \eta_{t-1} + \epsilon_t, \quad \eta_t | S_t \sim i.i.d. N(0, \sigma_{\epsilon, S_t}^2), \quad (15)$$

where α captures the importance of the first lag of Γ_t on domestic inflation. η_t is an AR(1) process to capture any remaining serially correlated factors that may matter for inflation.

The estimation results are reported in Table 4. Overall, the inclusion of Γ_{t-1} in the UC specification does not significantly alter model parameter estimates, apart from the coefficient on the global output gap in the post 2007 period. As shown, after the inclusion of external factors, all k^* estimates in the third regime are now reduced and statistically significant, implying that in the two-break baseline specification, they were large in magnitude merely due to the effects of external factors through the direct import price channel. This result is in line with the finding that the influence of external factors as captured by α_3 are statistically significant only in the third regime.

Focusing on the results in the first column, we find that the direct effect of import prices on Thai inflation is negligible in the pre 2007 period, but is as high as 0.3 thereafter. Upon further investigation, we find that the influence of oil prices is largely responsible for this finding, largely due to three main reasons. First, the coefficient on world oil prices is also statistically significant in the post 2007 period, as shown in the second column. Second, by including non-oil imports as an external factor, we find that non-oil import prices has no

bearing on domestic inflation, even in the post 2007 period¹⁵. Third, while the influence of non-oil commodity prices on inflation is statistically significant in the post 2007 regime as reported in the third column, once we include both oil and non-oil commodity prices in the UC specification as external factors, non-oil commodity price movements are no longer relevant for Thai inflation dynamics. Therefore, the significant effect on non-oil commodity prices in the post crisis period may simply be a spurious reflection of movements in oil prices¹⁶.

Finally, for real exchange rate effects, we find that the degree of pass-through is of the right sign and statistically significant only in the second regime. This result is consistent with an extensive study by Buddhari and Chensavadijai (2003), where the authors show that the degree of exchange rate pass through in Thailand is typically low, which may be due to lower inflation expectations as well as the prevalence of administered price measures. We also observe that during the first and third regimes in which the exchange rate pass-through effect is statistically insignificant, these were periods of high exchange rate volatility, and relatively low inflation respectively¹⁷. This result is in line with the findings of a number of empirical studies. McCarthy (1999), Goldfajn and Werlang (2000) and Styryn et al. (2012) find that the degree of exchange rate pass-through is typically lower when exchange rate variability is high, or when the country is in a low inflation environment. For the latter reason, this is because in a low inflation environment, price increases are more noticeable to consumers, causing demand to react more strongly to the price increase. As a result, producers become more reluctant to pass-through production costs to prices.

¹⁵Note that the import price series are in Thai Baht, but oil prices are in USD, so we also include the foreign exchange rate between the Thai baht and US in the UC specification with oil. The coefficients on the exchange rate are not significantly different from zero so we do not report the results due to space considerations. The same applies for the non-oil commodity price results. Results available upon request.

¹⁶We also allowed the disaggregated components of non-oil commodities (food, metals, beverages, agricultural products) to enter the UC specification one at a time, but none of their coefficients were statistically significant. Also, the coefficient in front of Thailand's changes in terms of trade was not statistically significant so we did not report the results due to space considerations.

¹⁷The sharp depreciation of the Thai currency following the abandonment of the exchange rate peg after the 1997 financial crisis induced high exchange rate volatility during that time. Inflation has also drifted lower since the GFC especially during recent periods due to low oil prices.

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.

4.4 Further Discussion of Results

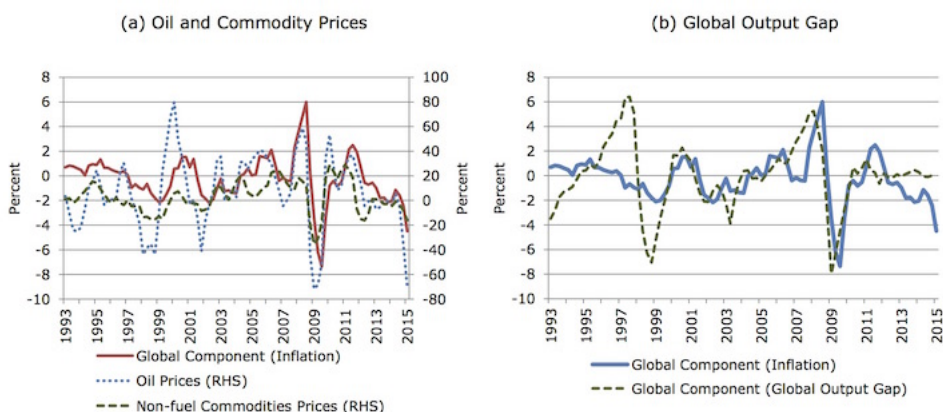
The findings thus far point towards the importance of a global output gap for Thai inflation dynamics since 2001, albeit the significance of the global gap is mainly determined by the direct impact of world oil prices since 2007. In contrast, the global output gap in the 2001-2007 period is influenced by both the direct impact of world oil prices, and to a larger extent other indirect effects of globalization on inflation resulting from enhanced competition, participation in global value chains, and technology spillovers. To explain why oil price movements dominate short-run inflation dynamics since 2007 while the indirect effects of globalization beyond the direct import price channel disappears require a structural model which is outside the scope of this paper. However, in this section we offer some suggestive evidence that the rising influence of oil in the post-crisis period alongside the diminished role of the indirect price channel is not a phenomenon specific to Thailand, but is in fact a global phenomenon.

Regarding the impact of oil, we carried out some investigations and found that our findings are unrelated to increases in domestic oil consumption. This is because the share of energy components in the CPI has already been gradually increasing since the mid 1990s, and there has actually been a slowdown in fuel consumption since 2007. Furthermore, as shown in Figure 12a, it is clear that global inflation, proxied by the estimated global factor from the DFM analysis in Section 2, comoved closely with world oil prices only since 2007, with a degree of correlation that increased from 61 to 78 percent. Such a change could correspond with the growing evidence in the existing literature that there has been structural changes 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 during the past 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.

Similarly, the diminished role for the global output gap beyond the direct effects of oil since 2007 may be a global phenomenon. In Figure 12b, the global factor from the DFM analysis is highly correlated with the global output gap from 2000 until after the crisis, but is uncorrelated thereafter. This implies that in the post crisis period, prices may no longer be affected by the gains from enhanced productivity and competition. This phenomenon can related to the recent slow down in global trade, which is also known as the 'Great Trade Collapse' (Ferrantino and Taglioni, 2014). Boz et al. (2014) finds that half of the slowdown in global trade can be attributed to cyclical factors stemming from weak global demand, while

the remaining half is due to structural factors. As for structural factors, Constantinescu et al. (2015) explains that in the post-Great Recession period there has been a marked change in trade-income relationship. More specifically, they explain that global trade is growing more slowly not only because world income growth is lower, but also because trade has become less responsive to income growth. In addition, participation in global value chains (GVCs)¹⁸ has been growing at a slower pace since the global financial crisis. As shown in Figure 13, GVCs have been increasing for both advanced and emerging economies up until the mid 2000s period, but declined or flattened since then. The slower pace of global supply chain is an important determinant to reduce trade activity and lower the global elasticity of trade because it also reflects the less border crossing of intermediate goods.

Figure 12: Global Inflation and its Determinants

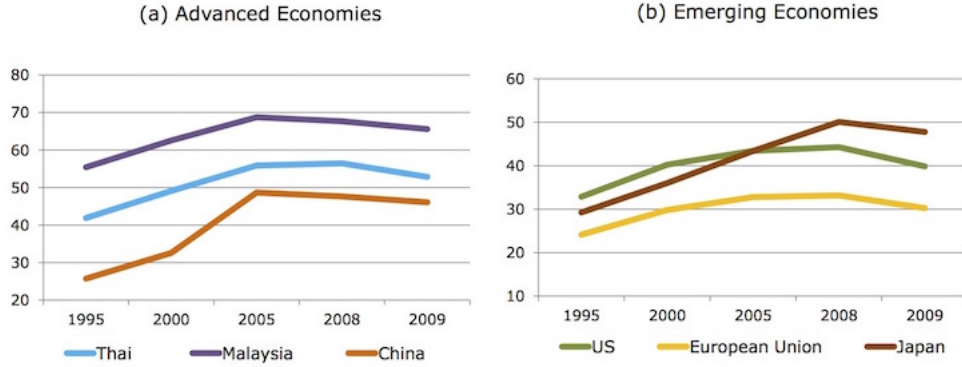


Note: Global inflation and the global output gap is extracted from a dynamic factor model for inflation and output gaps respectively. The change in Dubai oil prices is the log year on year change in the Dubai oil price series.

Sources: IMF International Financial Statistics, authors' calculations.

¹⁸GVC reflects the extent to which a country is a user of foreign inputs and a supplier of intermediate inputs that are used in other countries' exports.

Figure 13: Global Value Chain Participation



Sources: OECD

5 Conclusion

Globalization accelerated since the year 2000, and while it is generally accepted that globalization impacts inflation dynamics, there are disagreements about the channels as well as the magnitude of the effect. In this paper, we explore the effects of globalization for Thai inflation during the 1993-2015 period. First we develop a dynamic factor model to study the influence of domestic versus global factors in driving Thai inflation over time. Then, we extend the closed-economy New Keynesian Philips Curve (NKPC) of Kim et al. (2014) to an open-economy framework within the context of an unobserved components (UC) model. Careful testing of structural breaks within the UC model are carried out to identify structural changes, and we find that the dynamics of Thai inflation has undergone two distinct shifts in 2001 and 2007.

The empirical results suggest that in 2001, the short-run relationship between inflation and the domestic output gap has weakened, also known as a flattening of the Phillips curve. While this phenomenon may be attributed to improved monetary policy, the increase in sensitivity of inflation to global slack conditions that occurred during the same time favors the globalization hypothesis in the spirit of Borio and Filardo (2007). With a weakened relationship between inflation and domestic economic conditions, the implications for monetary policy are twofold. First, a given monetary expansion can be associated with lower inflation (Frankel, 2009), but at the same time, a flatter Phillips curve also implies that it is more costly to fight against increases in inflation. In today's low inflation environment, the flattened Phillips curve gives central bankers more room to pursue expansionary agendas. Nevertheless, it must be cautioned that policymakers do not discount the inflationary risks of expansionary monetary policy. In a more integrated economy, as evidenced by the increased

dependence between inflation and global slack conditions, easy monetary policy can create a global inflation bias if countries fail to internalize aggregate inflationary consequences of those policies at the global level.

Since the global financial crisis in 2007, Thai inflation underwent another structural change, mainly for two reasons. First, the sensitivity of inflation to global gap conditions that extend beyond the direct import price channel have been muted, most likely due to structural changes in trade relationships, less integrated global value chains, and the slowdown in global trade. Second, global oil prices play a greater role in determining the overall price movements than ever before. Based on the existing literature, oil price dynamics may have undergone a structural break after the global financial crisis. For example, Arezki et al. (2015) reveal that supply-side factors appeared to have played a more important role in driving the 50 percent drop in the price of oil between in the past years, in contrast to the commodities super cycle in the 2000s that was mostly driven by the strong growth in emerging market economies such as China and India.

Overall, we find that external developments play a key role in driving inflation developments in Thailand, particularly since the early 2000s. However, we find that monetary policy in Thailand has been particularly effective in keeping Thai inflation low and stable. In particular, we find that the implementation of an inflation targeting framework in May 2000 has successfully lowered and stabilized trend inflation since 2001, suggesting that long-term inflation expectations in Thailand has remained well anchored since then. Therefore, in today's uncertain economy with future risks of deflation, the ability of the central bank to successfully anchor long-term inflation expectations will become all the more important in shielding the economy against volatile external shocks.

References

- Adams, F.G., and Y. Ichino, 1995. Commodity prices and inflation: A forward-looking price model. *Journal of Policy Modeling*, Vol. 17, 397-426.
- Arezki R., Breuer P., Husain A.M., Haksar V., Helbling T., Medas P. and Sommer M., 2015. Global implication of lower oil prices. IMF Staff Discussion Note, No. 15/15.
- Artis, M.J., Galvao, A.B., and M. Marcellino, 2007. The transmission mechanism in a changing world. *Journal of Applied Econometrics*, 22(1), 9-61.
- Auer, R., and A. Mehrotra, 2014. Trade linkages and the globalization of inflation in Asia and the Pacific. BIS Working Papers, No. 446.
- Baffes, J. M., Kose, A., Ohnsorge, F., and M. Stocker, 2015. The great plunge in oil prices: causes, consequences, and policy responses. World Bank Policy Research Notes, March 2015.
- Ball, L.M., 2006. Has globalization changed inflation? NBER Working Paper, No. 12687.
- Ball, L.M. and S. Mazumder, 2011. Inflation dynamics and the great recession. NBER Working Paper, No. 17044.
- Binici, M., Y-W., Cheung, and K. Lai, 2012. Trade openness, market competition and inflation: Some sectoral evidence from OECD countries. *International Journal of Finance and Economics*, 17(4), 321-36.
- Borio, C., and A. Filardo, 2007. Globalisation and inflation: New cross-country evidence on the global determinants of domestic inflation. BIS Working Paper, No. 227
- Boz E., Bussiere, M., and C. Marsili, 2014. Recent slowdown in global trade: Cyclical or structural? Centre for Economic Policy Research.
- Buddhari, A. and V. Chensavadijai, 2003. Inflation dynamics and its implications for monetary policy. Bank of Thailand Symposium, 2003.
- Calza, A., 2009. Globalization, domestic inflation and global output gaps: Evidence from the Euro area. *International Finance*, 12(3), 301-320.
- Canova, F., and J. Marrinan, 1998. Sources and propagation of international output cycles: Common shocks or transmission. *Journal of International Economics*, 46, 133-166.
- Cecchetti S. and R. Moessner. 2008. Commodity Prices and Inflation Dynamics. BIS Quarterly Review, December.
- Chantanahom, P., Poon Patpibul, C., and P. Vongsinsirikul, 2004. Exploring inflation in Thailand through sectoral price setting behavior and underlying trend. Bank of Thailand Symposium, 2004.
- Ciccarelli, M., Mojon, B., 2010. Global inflation. *The Review of Economics and Statistics*, 92(3), 524-535.
- Clarida, R., Gali, J., and Gertler, M., 2002. A simple framework for international mon-

etary policy analysis. *Journal of Monetary Economics*, 49(5), 879-904.

Cogley, T. and A. M. Sbordone, 2008. Trend inflation, indexation, and inflation persistence in the new Keynesian Phillips curve. *American Economic Review*, 98(5), 2101-26.

Constantinescu, C., Mattoo, A., and M. Ruta, 2015. The global trade slowdown: Cyclical or structural. Policy Research Working Paper No. 7158, World Bank

Corsetti, G., and Pesenti, P., 2005. International dimensions of optimal monetary policy. *Journal of Monetary Economics*, 52(2), 281-305.

Davis, J.S., 2012. The effect of commodity price shocks on underlying inflation: The role of central bank credibility. Federal Reserve Bank of Dallas Globalization and Monetary Policy Institute Working Paper No. 134.

Dees S., F. di Mauro, M. H. Pesaran, and L. V. Smith, 2007. Exploring the international linkages of the euro area: A global VAR analysis. *Journal of Applied Econometrics*, 22(1)

European Central Bank, 2006. Effects of the rising trade integration of low cost countries on Euro Area import prices, *European Central Bank Monthly Bulletin*, August, 56-57.

Ferrantino, M. and D. Taglioni, 2014. Global value chains in the current trade slowdown. World Bank Economic Premise 137.

Feyzioglu, T., and L. Willard, 2006. Does inflation in China affect the United States and Japan? IMF Working Paper, No. 06/36.

Frankel, J., 2009. Comment on the simple geometry of transmission and stabilization in closed and open economies. NBER International Seminar on Macroeconomics 2007, NBER chapters, 119-129

Gamber, E. N., and J. H. Hung, 2001. Has the rise in globalization reduced U.S. inflation in the 1990s?, *Economic Inquiry*, 39(1), 5873.

Goldfajn, I., and Werlang, S.R.C., 2000. The pass through from depreciation to inflation: A panel study. Banco Central do Brasil Working Paper Series No. 5

Ihrig, J., Kamin, S.B., Lindner, D., and J. Marquez, 2007. Some simple tests of the globalization and inflation hypothesis. *International Finance*, 13(3), 343-375.

International Monetary Fund. 2006. How has globalization changed inflation? World economic outlook, April, Chapter 3

International Monetary Fund. 2013. The dog that didn't bark: Has inflation been muzzled or was it just sleeping? World Economic Outlook, April.

Kamin, S., Marazzi, M., and J.W. Schindler, 2006. The impact of Chinese exports on global import prices. *Review of International Economics*, Vol. 14, 179-201.

Khemangkorn, V., Mallikamas, R.P., and P. Sutthasri, 2008. Inflation dynamics and implications on monetary policy. Bank of Thailand Symposium 2008.

Kilian, L., 2009. Not all oil price shocks are alike: Disentangling demand and supply

shocks in the crude oil market. *American Economic Review* 99(3), 1053-69.

Kim, C. J., Manopimoke, P., and C.R. Nelson, 2014. Trend inflation and the nature of structural breaks in the New Keynesian Phillips curve. *Journal of Money, Credit and Banking*, 46(2-3), 253-266.

Kohn, D.L., 2006. The effect of globalization on inflation and their implications for monetary policy. Speech at the Federal Reserve Bank of Boston's 51st Economic Conference, Chatham, Massachusetts.

Manopimoke, P., 2015. Globalization and international inflation dynamics: The Role of the global output gap. PIER Discussion Paper, No. 8.

McCarthy, J., 1999. Pass through of exchange rates and import prices do domestic inflation in some industrialized economies. BIS Working Papers, No. 79.

Milani, F., 2010. Global slack and domestic inflation rates: A structural investigation for the G-7 Countries, *Journal of Macroeconomics*, 32(4), 968-981.

Neely, C.J. and D.E. Rapach, 2011. International comovements in inflation rates and country characteristics. *Journal of International Money and Finance*, Vol. 30(7), 1471-90.

Neiss, Katherine S., 2001. The markup and inflation: Evidence in OECD countries, *Canadian Journal of Economics*, 34, 570-587.

OECD., 2006. OECD Economic Outlook, Vol. 2006/1. OECD, Paris.

Pain, N., Koske, I., and M. Sollie, 2008. Globalisation and OECD consumer price inflation. *OECD Journal: Economic Studies*, Issue 1, 1-32.

Rae, D., and D. Turner, 2001. A small global forecasting model. OECD Economics Department Working Papers, No. 286

Razin, A. and A. Binyamini, 2007. Flattened inflation-output tradeoff and enhanced anti-inflation policy: Outcome of globalization? NBER working paper, No. 13280

Razin, A. and P Loungani, 2005. Globalization and equilibrium inflation-output tradeoffs. NBER working paper, No. 11641

Razin, A. and C-W. Yuen, 2002. The new Keynesian Phillips curve: Closed economy versus open economy. *Economics Letters*, Elsevier, vol. 75(1), March

Rogoff, K.S. 2003. Globalization and global disinflation. In *Monetary Policy and Uncertainty: Adapting to a Changing Economy*. Federal Reserve Bank of Kansas City.

Rogoff, K.S. 2006. Impact of globalization on monetary policy. *Proceedings in the Economic Policy Symposium*, Jackson Hole, Federal Reserve Bank of Kansas City, 265-305.

Romer, D., 1993. Openness and inflation: Theory and evidence. *The Quarterly Journal of Economics* 108(4), 869-903.

Styrin K. and O., Zamulin, 2012. Exchange rate pass-through, monetary policy, and variability of exchange rates. Center for Economic and Financial Research Working Papers.

Steinsson, J., 2005. The dynamic behavior of the real exchange rate in sticky-price models. *American Economic Review*, 98(1), 519-533.

Stock, J.H., and M. Watson, 2007. Why has U.S. Inflation become harder to forecast? *Journal of Money, Credit and Banking*, 39, 3-33.

Tootell, G. M., 1998. Globalization and US inflation. *New England Economic Review*, 98, 21-33.

White, W., 2008. Globalisation and the determinants of domestic inflation. BIS Working Papers 250, Bank for International Settlements.

World Trade Organization, 2014. The rise of global value chains. *World Trade Report*, 2014.

Zaniboni, N., 2011. Globalization and the Phillips Curve. Georgetown University.

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

The corresponding State-Space representation for the UC model in Eqs. (7)-(13) can be written as:

Measurement equation

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

Transition equation

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

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

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

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

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

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

Note that 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 $\tilde{X}_{t-1}^* = \begin{bmatrix} x_{t-1}^* \\ x_{t-2}^* \end{bmatrix}$.

Once put into State-Space form, the UC model can be estimated with the Kalman filter.

Appendix B

Here we present the estimation results from the no-structural break and one-break models. Based on tests of serial correlation on the models' standardized residuals, we explain why the two structural UC-NKPC model is the appropriate choice for Thai inflation dynamics.

First, we present the estimation results from the no-break model, as shown in Table B1.

Table B1: Estimation of 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)
<i>Log-likelihood value: -292.440</i>	

Note: Standard errors are in parentheses.

**,*,* denote statistical significance at the 1, 5, and 10 percent levels respectively.

Based on tests of serial correlation, we see that the no-break model is misspecified. As shown in column 2 of Table B2, there remains serial correlation in the square of the standardized residuals at the 10 percent level, signifying remaining ARCH effects.

Table B2: Tests of model misspecification

Standardized residuals				
Lag	No-break	One-break	Two-break (HP-filtered global gap)	Two-break (UC global gap)
1	0.561	0.833	0.268	0.325
2	0.167	0.450	0.246	0.248
3	0.277	0.617	0.420	0.413
4	0.405	0.764	0.557	0.541
5	0.385	0.521	0.175	0.184
6	0.462	0.611	0.262	0.268
7	0.464	0.644	0.353	0.367
8	0.566	0.668	0.430	0.428
Square of standardized residuals				
1	0.083	0.424	0.200	0.244
2	0.063	0.345	0.434	0.485
3	0.133	0.494	0.485	0.586
4	0.230	0.662	0.563	0.687
5	0.320	0.762	0.700	0.785
6	0.424	0.830	0.661	0.825
7	0.528	0.891	0.697	0.862
8	0.573	0.865	0.661	0.872

Note: P-values of Ljung Box tests under the null of no serial correlation.

In Table B3, we report the estimation results of the one structural break model.

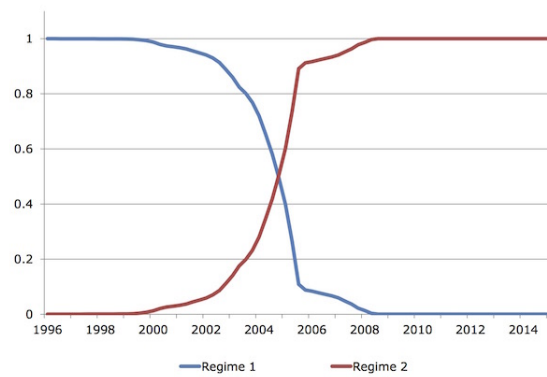
Table B3: Estimation of the UC OE-NKPC with one structural break [1993Q1-2015Q1]

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) → Break date: 2002Q2	
<i>Log-likelihood value: -289.967</i>		

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

According to the Ljung-box test statistics in column 3 of Table B2, the one-structural break model appears well specified. However, by examining the smoothed probabilities of the UC model in Figure B1, the transitional period between the two regimes occurs over the span of approximately 8 years, suggesting that we may be able to find a third regime during this transitional period. We estimate the two break model which offers sharper transitions between regimes, and report the estimation results in the main body of the paper. The serial correlation tests associated with the two structural break model (columns 4 and 5) shows that the model is not misspecified.

Figure C1: Smoothed Probabilities



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.