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Exchange Rate Pass-through (ERPT) into Domestic Prices: Evidence from a Nonlinear Perspective

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Abstract

This paper uses a Markov-switching intercept autoregressive heteroskedasticity model to examine inflation dynamics and the exchange rate pass-through (ERPT) in Malaysia. Two different regimes are identified in the sample period (1990:Q1–2015:Q4). The results confirm a partial ERPT to domestic consumer prices, but the effect of exchange rates on domestic prices is asymmetry. The incomplete ERPT to the consumer price index is consistent with findings in earlier studies. We also confirm a positive but time-varying impact of terms of trade shocks on inflation.

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1. Introduction

Since the late 1970s, the dynamic linkage between exchange rates and prices has received a great deal of attention. Exchange rate movements are known to have a direct impact on price stability, financial stability, and trade balances.[1] The degree of exchange rate pass-through (ERPT), defined as the effect of exchange rate changes on aggregate prices (e.g., import prices, producer prices, and consumer prices), is crucial to the conduct of monetary policy, as it is an important parameter with regard to monitoring and forecasting domestic inflation. Traditional literature is focusing the ERPT to import prices and has stressed the role of market power and price discrimination in international markets (see Choudhri and Hakura, 2001). The accumulated evidence, focusing mainly on data from developed countries, reveals that the pass-through is incomplete—that is, the effect of a unit change in an exchange rate leads to less than a unit change in a given domestic price. The literature using advanced countries as a sample has also documented a decline over time. Several authors have linked the low ERPT to the inflation environment (Gagnon and Ihrig, 2004; Taylor, 2000). Gagnon and Ihrig (2004), for example, tested the hypothesis in 20 developed countries and find that a low and stable inflation rate (reflecting central bank policies) tends to be associated with a low ERPT into consumer prices.[2]

In a seminal paper, Frankel *et al.* (2012) find a partial ERPT and that the degree of ERPT in the 1990s declined significantly in developing countries (e.g., Barhoumi and Jouini, 2008), much more so than in developed countries. Recent notable studies in this line of research include Aleem and Lahiani (2014), Brun-Aguerre *et al.* (2012), Ito and Sato (2008), and Prasertnukul *et al.* (2010). According to these authors, the slow and low pass-through is common not only in developed countries, as claimed by Calvo and Reinhart (2000).[3] A low-inflation regime brought about by credible monetary policies provides greater freedom for a country to pursue an independent monetary policy and to target inflation, which decreases the degree of ERPT. This argument, popularized by Taylor (2000), has received a lot of attention and explains that a low-inflation regime (declining in the volatility of inflation) can reduce the transmission of fluctuations in exchange rates to domestic prices. Similarly, Ben Cheikh and Louhichi (2016), who apply a panel threshold regression, show that countries with higher inflation rates, a barometer of macroeconomic performance, tend to experience a higher degree of ERPT.[4] This means that exchange rate movements in these high-inflation countries have a larger impact on consumer prices and, hence, on short-run inflation. Finally, the literature (e.g., Shioji, 2015) has also pointed out that ERPT to domestic prices in Japan, once considered weak (near zero), has become somewhat stronger. Using Japanese data, Shioji shows a stronger ERPT to domestic consumer prices in recent years.

This paper concerns the degree of ERPT to domestic prices in Malaysia, an emerging market economy characterized by moderate inflation levels. Specifically, the question we explore is: what are the extent and nature of ERPT into the consumer price index (CPI)? In doing so, we extend the existing literature on Malaysia by employing an econometric method to capture the time-varying ERPT to shed additional light on the extent of the pass-through. The endogeneity selections of

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[1] According to Ito and Sato (2008), if domestic prices respond to nominal exchange rate depreciation at a ratio of one to one, then any export competitiveness that accrues from nominal depreciation would be wiped out, because the real exchange rate would remain unchanged. By contrast, when pass-through is low, exchange rate changes do not advance the external adjustments of the economy.

[2] They also find that countries in which either the level or variability of inflation declined substantially tended to have large declines in estimated ERPT.

[3] Calvo and Reinhart (2000) presented evidence that ERPT is larger in emerging economies compared to advanced countries because the latter tend to have higher inflation rates.

[4] Taylor's view has received a lot of support in the literature based on different techniques and procedures. Earlier studies have used dummy variables to provide empirical support for Taylor's view. For example, Edwards (2006) creates a dummy variable that takes the value of one at the time of the adoption of inflation targeting and zero otherwise to show that the adoption of inflation targeting reduced ERPT in both advanced and emerging countries. This finding endorses the view that a credible inflation target provides an anchor for medium- and long-term inflation expectations, thereby making it easier for the business sector to function properly, boosting its growth, and allowing for more effective management of monetary policy.

breakpoints makes the Markov-switching intercept autoregressive heteroskedasticity (MSIAH) model an attractive option for addressing this issue. Briefly, this model allows us to consider possible shifts in the evolution of inflation, which may better represent the actual behavior of inflation; following a probabilistic approach, it is frequently subject to a shift from one state to another and back again (Hamilton, 1990). To the best of our knowledge, this is the first study that applies the Markov-switching method for ERPT to an emerging Asian country.

2. Empirical Model

In modeling ERPT, we allow for time-varying coefficients using the MSIAH approach. It fits well with the idea that the exchange rate declined in the late 1990s following structural reforms was undertaken by the government in the aftermath of the 1997 Asian financial crisis. A special feature of the model is that it allows all the parameters of autoregression, including the intercept term and the variance, to switch between regimes. The model with a state-dependent coefficient can be presented as:[5]

$$\Delta p_t = \begin{cases} \alpha_1 + \delta_1 \Delta p_{t-1} + \theta_1 (y-\bar{y})_{t-1} + \beta_1 \Delta p^*_{t-1} + \sum\limits_{i=0}^{q} \phi_{1i} \Delta er_{t-i} + \sum\limits_{i=0}^{q} \lambda_{1i} \Delta tot_{t-i} + \varepsilon_{1t}, for \quad s_t = 1 \\ \alpha_2 + \delta_2 \Delta p_{t-1} + \theta_2 (y-\bar{y})_{t-1} + \beta_2 \Delta p^*_{t-1} + \sum\limits_{i=0}^{q} \phi_{2i} \Delta er_{t-i} + \sum\limits_{i=0}^{q} \lambda_{2i} \Delta tot_{t-i} + \varepsilon_{2t}, for \quad s_t = 2 \end{cases} \quad (1)$$

where Δ denotes the first-difference operator, p_t is the CPI, $(y-\bar{y})_t$ is the output gap as an indicator for the demand-pull factor, p^*_t is foreign price (proxy by US producer price index), er_t is the nominal exchange rate, tot_t is terms of trade, and $\varepsilon_t \sim N[0, \sigma^2(s_t)]$ at $t = 1, \ldots, T$.[6] All variables in Eq (1) are in logarithm. Note that the price change was calculated as annualized quarterly difference in the logarithm of the CPI while the exchange rate change was calculated as the quarterly difference in the logarithm of the nominal exchange rate. The output gap is constructed as the deviation of the log industrial production index from its Hodrick-Prescott (HP) filtered trend series. The terms of trade are added to the equation to account for a shock that could arise from a change in the energy commodity price of exports on the domestic inflation rate. Theory predicts that the link between the terms of trade and inflation can be positive or negative under a floating exchange rate regime. If the real exchange rate moves in an almost one-for-one fashion with the terms of trade, it puts downward pressure on inflation (Gruen and Dwyer, 1995).

The output gap is considered an important factor in capturing domestic demand shocks in the sample period; s_t is an unobserved random variable denoting the regime, which follows a Markov chain defined by transition probabilities between N states. The transition probabilities govern the movement from one regime (s_t = 1) to another (s_t = 2) and vice versa are estimated. The short-run ERPT in a stable regime is captured by the parameter $\sum_{i=0}^{q} \phi_{1i}$, while long-run pass-through is presented as $\sum_{i=0}^{q} \phi_{1i} / (1-\delta_1)$. Based on this formula, another set of measures is obtained for the ERPT during an unstable regime. We set the lag length (q) to four. Exporters can set prices in their own currencies, referred to in the literature as producer currency pricing, or in currencies of destination markets, or local currency pricing. Currency invoicing choices affect the degree of pass-through effects. The former represents a full ERPT whereas the latter implies a zero pass-through to consumer prices. If the size of the pass-through coefficient is equal to one, then there is an evidence

[5] Several studies (e.g., Shintani *et al.*, 2013) assume these coefficients evolve smoothly. In contrast, we draw on the idea of changing regimes.

[6] Any shocks in energy prices of commodities might have knock-on effects on domestic inflation, and export earnings can change, affecting the exchange rate. We thank an anonymous referee for mentioning to us the impact on inflation of terms of trade shock, and as suggested we include the energy commodity price deflated by the US producer price index as a proxy for the terms of trade, where an exogenous relative foreign price reflects the export commodity basket.

of a complete ERPT. Incomplete (partial) pass-through will occur if the size of the coefficient is less than one.

In the context of Malaysia, a lower pass-through is preferred as it avoids external inflationary pressure on the domestic economy during periods of currency depreciation. Accordingly, the probability matrix consists of transition probabilities p_{ij} from state *i* to state *j*:

$$\Pi = \begin{bmatrix} p_{11} & p_{21} \\ p_{12} & p_{22} \end{bmatrix}, p_{ij} = pr(s_t = j \mid s_{t-1} = i) \tag{2}$$

The smoothed probability $pr(s_t = j \mid r_1, \ldots, r_T)$ of being in state *j* is based on knowledge of the complete series. Hence, the Markov-switching model allows two or more processes to exist with a series of shifts between the states occurring in a probabilistic manner, so that the shift occurs endogenously.

3. Results and Discussion

Quarterly seasonally adjusted frequency data from 1990: Q1 to 2015: Q4 are used in the analysis. Series for the CPI (2010 = 100), bilateral US dollar rates, the nominal effective exchange rate (2010 = 100), the world commodity index (2010 = 100), foreign prices (US producer price index (P^*), 2010 = 100), and the industrial production volume index (IPI, 2010 = 100) come from the International Monetary Fund's *International Financial Statistics (IFS)*. Unit labor costs (2010 = 100), which measure the labor cost per unit of output produced, come from Malaysia Productivity Corporation (MPC) and Oxford Economics. The energy commodity price index (2010 = 100) drawn from World Bank. A brief data description is provided as per Table I.

Table I: Data description.

Variable	Data Source	Definition	Mean	Std dev
Δp	IFS	Annualized quarterly difference in the logarithm of the CPI	2.761	2.562
$(y-\bar{y})$	IFS	Deviation of the log IPI from its HP-filtered trend series	-0.365	5.002
Δer	IFS	Quarterly difference in the logarithm of the bilateral US dollar rates	0.442	4.288
Δp^*	IFS	Quarterly difference in the logarithm of the P^*	0.489	1.947
Δtot	World Bank, IFS	Quarterly difference in the logarithm of the commodity price deflated by the P^*	0.117	8.715

First, we fit the data to the linear Philips curve as a preliminary step. The results in Table II (Column 1) show that all the variables enter with the correct sign (as predicted by theory) except for the foreign price. Most of the variables of interest are not significant at it indicated level, even after allowing for extreme values (2008: Q3, 2008: Q4 and 2005: Q1) identified using a tree-search procedure. Without our going into much detail, note that the size of the pass-through coefficient is relatively small (0.110) and not significant at conventional levels. The small coefficient means that the exchange rate surprises explain only a small proportion of consumer price variation.

To contribute to the current debate on the ERPT, we apply the MSIAH model. All in all, the diagnostic tests show that the regime-dependent model performs better than the standard linear specification in explaining inflation dynamics over the sample period that ended in 2015: Q4. First, the linearity test strongly rejects the null hypothesis of a linear relationship among the variables at the indicated level, even at the upper bound (Davies, 1987). Second, the MSIAH model is immune to the non-normality, autocorrelation, and autoregressive conditional heteroskedasticity (ARCH) effects, suggesting that the MSIAH model fits Malaysian data adequately—again confirming the appropriateness of the Markov-switching model over the linear specification.

Table II: Estimated coefficients for the MSIAH.

	(1)			(2)				
	Linear Model			Nonlinear Model				
				Stable Regime			Unstable Regime	
				$s_t = 1$			$s_t = 2$	
α	1.975***	(0.417)	α_1	0.905***	(0.103)	α_2	3.832***	(0.567)
δ	0.278**	(0.133)	δ_1	0.378***	(0.039)	δ_2	-0.057	(0.159)
θ	0.050	(0.047)	θ_1	0.023	(0.023)	θ_2	-0.011	(0.080)
β	-0.001	(0.162)	β_1	-0.044	(0.062)	β_2	0.020	(0.216)
$\Sigma_{i=0}^{q}\phi_i$	0.110	[0.150]	$\sum_{i=0}^{q}\phi_{1i}$	-0.103**	[0.023]	$\sum_{i=0}^{q}\phi_{2i}$	0.151**	[0.012]
$\Sigma_{i=0}^{q}\lambda_i$	0.010	[0.678]	$\sum_{i=0}^{q}\lambda_{1i}$	-0.009	[0.177]	$\sum_{i=0}^{q}\lambda_{2i}$	0.014**	[0.039]
			σ_1^2	0.522***	(0.067)	σ_2^2	1.663***	(0.192)
			$p\_\{0\|0\}$	0.890***	(0.051)			
			$p\_\{0\|1\}$	0.085*	(0.046)			
Test of Restrictions								
$H_0 : \sum_{i=0}^{q}\phi_{1i} = \sum_{i=0}^{q}\phi_{2i}$				0.221	[0.639]			
$H_0 : \delta_1 = \delta_2$				7.057	[0.008]			
Short- run								
$H_0 : \sum_{i=0}^{q}\phi_i = 1$				584.442	[0.000]		198.402	[0.000]
Long-run								
$H_0 : \sum_{i=0}^{q}\phi_i = 0$				4.903	[0.027]		7.489	[0.006]
$H_0 : \sum_{i=0}^{q}\phi_i = 1$				241.208	[0.000]		269.995	[0.000]
Exclusion restriction for exchange rate				10.682	[0.005]		6.949	[0.031]
Diagnostic Tests								
Log-Likelihood		-198.870		-164.762				
AIC		4.220		3.735				
LR Linearity				54.253	[0.000]			
AR(8)				9.201	[0.326]			
ARCH(8)				0.512	[0.843]			
Norm(2)				1.110	[0.574]			

Notes: ***, ** and * significant at 1, 5 and 10%, respectively. Values in parentheses are robust standard errors. Values in square brackets are the probability for each test. Chi-square (χ^2) statistics using robust standard errors are reported for the hypothesis tests. AIC = Akaike information criterion. Likelihood ratio (LR) statistic here is to test the null hypothesis of a linear model against the alternative hypothesis of the two regimes. AR(n) is the nth order lagrange multiplier (LM) test for autocorrelation; ARCH (m) is the mth order test for autoregressive conditional heteroskedasticity, and Norm (2) is the normality test.

Turning to the nonlinear specification (Column 2), we note that all the coefficients carry the expected sign. Among the control variables, the output gap and foreign price do not appear to have any significant impact on inflation in either regime. This result remained the same when the output gap is replaced by the rate of growth in the gross domestic product (as in Campa and Goldberg, 2005). It should be noted that Mihaljek and Klau (2008) find that the relationship between the output gap and inflation is negative, but the size of the coefficient is relatively small (–0.04). Based on the size of the robust standard error, regime 2 ($s_t = 2$) is associated with periods of economic turbulence in our analysis. Following the literature, regime 1 ($s_t = 2$) is defined as the stable regime and regime 2 as the unstable (high and volatile) regime. We observe that the terms of trade turns out to be significant only in the unstable regime. Specifically, the terms of trade raise inflation during the unstable regime. In a stable regime, the terms of trade put downward pressure on the inflation rate but are not significant at its indicated significance level. The output gap and foreign price, however, enters into the specification with the asymmetry sign but was found to be statistically insignificant.

We found that depreciation of the Malaysian ringgit induced a significant increase in domestic prices following the introduction of capital market liberalization initiatives in the early 1990s. The depreciation of the ringgit in the late 2000s is closely connected with the quantitative and qualitative (QQ) monetary easing gradual exit strategy implemented by the US Federal Reserve. This asset tapering puts pressure on the ringgit, which has been depreciating since then (see Table III for endogenous regime classification). It useful to note that the unstable regime detected in 2013: Q4–2015: Q4 period is tightly linked to the recent decline in energy (fuel) commodity prices. In our view, this finding suggests that the economy is highly dependent on energy commodities. Figure 1 plots the inflation, inflation volatility, changes in the exchange rate, and smoothed regime probabilities during an unstable regime for the entire sample period. The inflation volatility is generated from the exponential generalized autoregressive conditional heteroskedasticity (EGARCH) (1,1) model. As can be seen from the figure, inflation remains in a stable regime during a fixed peg with the US dollar in 1998: Q3–2005: Q2 (see period shaded in gray) but quickly shifts to a high and volatile regime a few quarters before the ringgit is officially unpegged in 2005. Note that a few large changes or swing in the exchange rate are observed, especially during the Asian financial crisis of 1997-98, the recent global financial crisis, and more recently during declines in oil prices, which are also captured by the inflation volatility.

Malaysia has undergone its major structural transformation from a commodity-based economy to an industrial-based economy since last three decades. The agriculture sector is now considering a minor sector of the Malaysian economy, accounting for 8% of GDP (RM 24, 273 million) in 2015: Q3. Meanwhile, the share of output accounted by the manufacturing sector has steadily increased since the transformation of the Malaysian economy. In 2015: Q3, the manufacturing sector is around 22.7% of GDP (RM69, 012 million). Malaysia is highly dependent on food imports for domestic food processing and production. The large swing in the exchange rate has generated some fluctuation in the food inflation (so-called imported inflation) and we can observe several spikes of inflation.

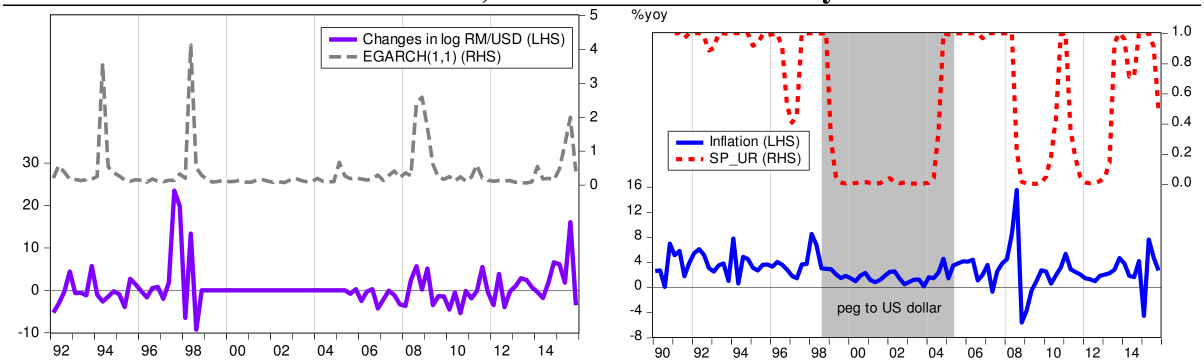
Table III: Expected duration of the regimes (in quarters).

Stable Regime		Unstable Regime	
$s_t = 1$		$s_t = 2$	
1997Q1-1997Q2	2	1991Q1-1996Q4	24
1999Q1-2004Q3	23	1997Q3-1998Q4	6
2008Q3-2010Q3	9	2004Q4-2008Q2	15
2011Q2-2013Q2	9	2010Q4-2011Q1	2
		2013Q3-2015Q4	10
Average duration	10.75		11.40
Annualized inflation rate	1.59%		3.63%

Next, we conduct a Wald test to formally determine whether the pass-through coefficient declined during the stable regime ($H_0 : \Sigma_{i=0}^{q}\phi_{1i} = \Sigma_{i=0}^{q}\phi_{2i}$). Test results failed to reject the null that the pass-through coefficient (-0.103 versus 0.151) is equal across regimes at the indicated level of significance. The small ERPT during the stable inflation environment may reflect that final goods purchases by households, including a large proportion of non-tradable components (e.g., consumption), however, have weakened the pass-through; see An and Wang (2012) and Burstein *et al.* (2007) on this issue. Bank Negara Malaysia (Malaysia's central bank) must factor in the changing behavior of ERPT in its forecasts. Our extent of the pass-through coefficient appears similar to that in other emerging countries as well as advanced countries. Chew *et al.* (2011) and

Peón and Brindis (2014), for instance, estimate pass-through of nearly zero. For example, Chew *et al.* (2011) observe ERPT to be around 1–2%. It should be pointed out that our estimates of the ERPT in the two-regime model are outside the confidence interval range [0.42–0.53] provided by Ben Cheikh and Louhichi (2016) for the low-inflation countries and much lower than in the model that ignores regime shifts.

Figure 1: Smoothed probabilities of being in a high regime, changes in the exchange rate, inflation, and inflation uncertainty



Notes: SP_UR refers to the smoothed regime probability in a high (or unstable) regime. The shaded (gray) regime refers to fixed peg with the US dollar. Inflation is measured as the annualized quarterly difference in the logarithm of the CPI. The inflation volatility is generated from the exponential generalized autoregressive conditional heteroskedasticity (EGARCH) (1,1) model.

Unlike in previous studies, we find evidence to suggest that the ERPT is regime dependent and the impact is asymmetry. There is inflationary effect of currency depreciation during the unstable regime. However, the sign of the coefficient is negative in the stable regime. This indicates that the depreciation tends to lower the domestic inflation rate during the stable regime, vice versa. Following Campa and Goldberg (2005) and others, we formally test for zero ERPT (local currency pricing) and complete ERPT (producer currency pricing). We find that ERPT into domestic prices during an unstable regime is around 0.151 in the short run. Typically, the response of the exchange rate to domestic prices appears to be higher in an unstable regime. A highly significant Wald statistic for restricting the coefficients to one ($H_0: \Sigma_{i=0}^{q}\phi_i = 1$) indicates that the ERPT is incomplete in both regimes. We conclude here that local currency pricing is more prevalent than producer currency pricing in the short run as well as the long run for both regimes. Ito *et al.* (2005) reported that the exchange rate accounts for less than 20% of the variation in the Malaysian CPI, which corroborates our results for an unstable regime.[7] Inflation persistence tends to be higher in an unstable regime, and since exchange rate movements are perceived to be less transitory, local firms tend to respond via price adjustments (stronger pass-through). Using the single equation method, Mihaljek and Klau (2008) estimate the extent of ERPT to CPI for 13 emerging countries. Their findings indicate that exchange rate depreciation has a stronger effect on inflation than appreciation. They find that the ERPT in Malaysia, unlike other countries, has been stable and relatively low pass-through (albeit not incomplete) over the sample period that ended in 2001. Similarly, Choudhri and Hakura (2001) provided evidence of low ERPT to domestic prices in Malaysia as well as other developing countries with single-digit inflation. The empirical findings confirm Taylor's hypothesis: The responsiveness of prices to exchange rate fluctuations depends positively on inflation.

In summary, our findings confirm that a commitment by the central bank to lower or control the inflation level is a crucial factor in determining an ERPT. Therefore, the depreciation of the ringgit will be less costly for Bank Negara Malaysia since its asymmetry impact on inflation is small. For

[7] According to the authors, the negligible response of Malaysia's CPI to the exchange rate shock may partly reflect the fact that it is an oil-producing country.

domestic prices to be responsive to the exchange rate, ERPT needs to be higher (Shioji, 2015). Aron *et al.* (2014) have stressed that the majority of existing studies overlooked destination country cost measures. For a robustness check, we include the rate of change of unit labor costs (ULCs) as a proxy for domestic costs in the model. Lags selection is based on the autometrics tree search procedure. The unreported results reveal that an increase in ULCs has no material impact on general price levels. One explanation is that Malaysia does not have well-developed labor unions, and so we believe that wage-push inflation (wage demands) via union activities is absent, at least in the sample period examined.[8] In the model, we consider both the nominal effective exchange rate and the bilateral US dollar rates (measured in domestic currency per unit of foreign). The model with a bilateral exchange rate and term of trade is chosen because it yields a lower Akaike information criterion (AIC) and log-likelihood value. Importantly, the results obtained from the bilateral rates produce more sensible outcomes in terms of the size and sign of the estimated coefficients. We also further the model with difference indicator for the term of trade and the sample period starts 1992: Q1 which restricted by the availability of the world commodity index. Our conclusion remains robust with the new indicator.

4. Conclusion

This study looks at the ERPT into Malaysia's domestic prices. Based on a two-regime MSIAH that allows ERPT to be time-varying, we show that the degree of ERPT is asymmetry state dependent and extremely low in both regimes. We offer qualified support, rejecting the conventional wisdom that ERPT is always higher in emerging than in developed countries. The survey by Aron *et al.* (2014) makes this point, too, when inflation regimes are properly accounted for in the analysis. Our analysis reveals that the exchange rate has no inflationary impact on inflation when the macroeconomic environment exhibits low and stable variations. In other words, pass-through to consumer prices is negligible in the low regime. Finally, a higher ERPT means that the potential effect of currency depreciation on consumer prices can be large during unstable regime. This means that, despite the large depreciation of the ringgit in recent years, the inflationary effect may not be lasting. As the economy shifts to a stable regime, the few movements in the exchange rate are expected to be transmitted to a lesser degree into domestic prices and monetary policy actions to control inflation can be more independent of exchange rate fluctuation.

Finally, the Markov-switching model cannot be burdened with too many parameters. A restricted (short-lag) structure and missing variables could lead to a biased and misleading measure of ERPT. Another potential caveat is that the results are subject to the problem of endogeneity (also mentioned in Aron *et al.*, 2014). These concerns need to be tempered in light of the limitations of the MSIAH single-equation approach used in this study. In future work, it would be interesting to consider a systematic approach to resolving the misspecification and potential endogeneity of exchange rate issues and develop a deeper understanding of ERPT. Another interesting issue will be the impact of exchange rate change on the export commodity price index.

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[8] These results are available upon request from the authors.

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