Analysis on Exchange Rate Pass-Through with Evidence from Northeast Asia

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Abstract : This paper theoretically and emprically investigates the relationship between exchange rate and price level. By building a theoretical model, where the central bank tries to offset the fluctuation in exchange rate market, we show that the exchange rate pass-through might be positive , negative or unobvious.With the data of three major economies in Northeast Asia, i.e., China, Japan and ROK, we show that the evidences do support our theory sequentially by means of Restricted and Markov- Switching Vector Autoregressive models.

Keywords: Pass-through, Cointegration, VECM, Markov-Switching VAR

JEL Classification: C12, F41

I. Introduction

Since the collapse of the Bretten Woods system, exchange rates have become floating in many countries. Since then, the fluctuation of exchange rates has generated lots of interests in the effects of exchange rate movements on national price levels. Henceforth, diverges on whether the effect of exchange rates pass-through exists come into being. Many literatures dispute on the existence and the strength of exchange rate pass-through to domestic goods price. For example, Feenstra (1989) uses U.S. import unit values from Japan as a measure of price for cars, trucks and motorcycles and shows that the effects of exchange rate pass-through on the above goods, especially on motorcycles are obvious. However, Papell (1993) shows that the exchange rates have relatively small influence on national price levels with data of G7 countries. Recently, a large body of theoretical research shows that the degree of exchange rate pass-through has stark implications for the conduct of monetary policy (e.g., Smets andWouters, 2002; Corsetti and Pesenti, 2005; Adolfson, 2002; Sutherland, 2005; and Monacelli, 2005), the choice of exchange-rate regime (Engel 2002; Devereux and Engel 2003), and the international transmission of shocks (see Betts and Devereux 2001). A parallel empirical literature has therefore developed to try to accurately measure exchange rate pass-through and to assess its stability across time. Those literatures including Kichian (2001)Campa and Goldberg (2002), Gagnon and Ihrig (2004), and Bailliu and Fujii(2004), suggest that exchange rate pass-through has declined in recent years in industrialized countries. But almost all the economists don't thnk that the pass-through effect might be negative, in which case, the appreciation of the currency in one country would give rise to inflation of the national price level since the imported goods price is supposed to be higher, measured by the local currency[2]. However, recent evidence, for example from China, shows that

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[2] Froot and Klemperer(1989) once built a model to show that the foreign producers might strategically decrease the imported goods prices to maintain their market shares. It eventually turns out tomake the exchange rate

the appreciation of the currency might lead to inflation instead of deflation. It seems very hard to explain such phenomenon by existing theories, which is an important motivation to this paper.

In this paper, we investigate the relationship between exchange rate and price level with the data of China, Japan and South Korea. Among them, exchange rate pass-through in Japan has comparatively widely been studied, since it is a member of G7 countries, while the effect in the other two countries, especially that in China has drawn few attention so far. As it is well known, after July 2005, the Chinese central bank has reformed its policy on exchange rate market. Henceforth, the Chinese currency, RMB becomes much more floating than before and the fluctuation of RMB is worthy of noticing. The Northeast Asia area is a very active and important economy, which has attracted many economists' interests. However, to our knowledge, there is few literature discussing on the exchange rate pass-through in this area, although currencies here are playing very important role in the international market and three countries possess huge foreign reserves.

The contributions of this paper are both theoretical and empirical. At first, we build a theoretical model where the central bank tries to offset the fluctuation of exchange rate market by monetary policy[3]. When the exhange rate depreciates, according to the traditional theory, the price of imports goods will increase measured by local currency. However, in our point of view, the central bank will tend to offset the depreciation by purchasing local currency with Forex Reserves, which, in turn, makes the money supply decrease. The tightening monetary policy will lead the price to go down instead of going up. The two effects are conflicting with each other. The pass-through effect thus becomes more flexible. It can be positive, negative or unobvious in different countries since the control power of the central banks differs from country to country. That means the appreciation of local currency might induce inflation instead of deflation in the country, which case, as we show, happens in China now.

Based on Dornbusch's (1976) rational expectation version of the Fleming-Mundall model and forward-looking price adjustment, we derive an open economic model where the reduced form is a non-linear combination of the structural coefficients. According to the model we construct, we start up our empirical work. As for the methodology, we use both Error Correction and Markov-switching Vector Autoregressive models. As Papell (1997) once pointed out that the levels of time sereis are generally having unit roots, which is shown in our paper too. Furthermore, as Papell (1993) once showed, we also verify the cointegration between exchange rate and its fundmentals. That means the error term might be correlated with the linear combination of the dependant and the independants. Hence,we use Vector Error Correction (1) model to estimate the exchange rate pass-through of the Northeast Asian countries.

Moreover, as Clements and Hendry(1999) once pointed out that structural breaks might result

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pass-through effect negaitve. As we discuss below, we think the negative effect is due to the strategy of the central bank instead of the price producers since it is intuitively very hard to imagine that the price can be endogenously decided.

[3] As we will see, such arrangement will make the money supply connected with the past exchange rate in our concrete model. Then we can get error correction VAR model with the parameter of money one period forward.

in that model. To make our theory robust, we also use Markov-switching (2)- VAR(1) to estimates the same data. While there have already been a lot of literatures focusing on modelling non-linear feature of time series, the attemptation to use regime-switching variables to estimate the parameters are comparatively rare, although it is well known that it is really necessary to assume the existence of invisible regime-switching variables. By the results of this paper, we can clearly see this necessity of this arrangement.It turns out that based on our evidence, the results are consistent with VEC model when the intercept and the coefficient of exchange rate are affected by regime shifts according to the regime switching nature while other variable coefficients are as tradition steady.

With the data from China, Japan and ROK, we find that the evidences do support our theory. Using traditional and regime switching empirical analysis methods, i.e., vector error correction and MS-VAR estimations, we find that although all three economies have a similar geographical condition[4], the exchange rate pass-through is positive in ROK, but negative in China, while , as many literatures argued, the effect is not obvious in Japan. As far as data is concerned, our data are all monthly instead of quarterly which is very popular in most of other literatures. The reason is straightforward. Usually, economists use real GDP, which are announced every quarter instead of every month, as a parameter to study exchange rate pass- through, while we don't in this paper. By the neoclassical macroeconomic theory, we find that real output can be denoted as a function of interest rate, price level and money supply all of which are available in monthly level and hence can act as instrumental variable for real GDP.[5]

This paper is organized as follows. Section 2 builds a theoretical model in which the central bank will try to stablizet the fluctuation of exchange rate market, which theoretically explain why the pass-through effect might be negative or positive. Section 3 describes the data scope and estimation methodologies in this paper. And Section 4 illustrates the estimation results to support our theory. Section 5 concludes and shows further research direction.

II. The Model

In this section, we derive a macroeconomic model in an open economy. We will study both the goods and the capital market. And the model built in this section will be foundation for our empirical work. To simplify, we use the assumption of perfect capital mobility[6]. Hence uncovered interest rate parity(UIRP) holds, which means that the domestic interest rate equals the foreign interest rate plus the expected rate of currency depreciation

[4] We note that the economic core of China lies in the east areas which are very close to Japan and Korea.

[5] Actually, as we will see in the next section, the domestic interest rate does not necessary appear in the estimation equation. According to uncovered interest rate parity, it depends on foreign interest rate and real and expected exchange rate.

[6] Although it seems not the case in China, it is since a lot of hot money does flow in China in the recent years secrectly.

$$r_t = r_t^* + (\hat{e}_{t+1} - e_t) + \varepsilon_{1t} \tag{1}$$

Based on the work of Dornbusch (1976), the goods market depends on parameters such as the price difference between export and import countries, the export country real output, the domestic monetary policy, etc., which can be described as follows:

$$y_t = \theta_1(e_t + p_t^* - p_t) + \theta_2 y_t^* - \theta_3 r_t + \varepsilon_{2t} \tag{2}$$

Moreover, following neoclassical macroeconomic model, the real output is a function of real money supply and interest rate with error term. Hence the monetary demand equation can be rearranged as follows:

$$y_t = m_t - p_t + \alpha_1 r_t + \varepsilon_{3t} \tag{3}$$

where p_t is the price level, m_t is the domestic money supply, y_t is the output and r_t is the interest rate for period *t*. In this paper, all variables are in logarithms except for interest rates. Besides, $\theta_1, \theta_2, \theta_3$ and α_1 are all time-independent constants. In the frame of this paper, all variables with asterisks denote the counterparts in the United States.[7] ε s are error terms. By equation (3), we can similarly get:

$$y_t^* = m_t^* - p_t^* + \beta_1 r_t^* + \varepsilon_{3t}^* \tag{4}$$

Following traditional arrangement, we still assume that the American price level p^* and interest rate r^* are first order autoregressive process[8].

$$p^*_t = \alpha_2 p^*_{t-1} + \varepsilon_{4t}.$$

$$r^*_t = \alpha_3 r^*_{t-1} + \varepsilon_{5t} \tag{5}$$

The standard open economy price adjustment specifications, Dornbusch (1976)， Mussa (1981) and Eaton and Turnovsky (1983), are not consistent with a unit root in the price level. In Mussa (1981), the rate of inflation is specified as equal to the expected change rate of the

7 Since exchange rates in this paper are denoted by the amount of local currency per unit U.S. dollar, we regard American variables as foreign ones in traditional literature.

8 In the regime of this paper, the United States is supposed to be a big and open economy, which cannot be affected by other countries' policy.

equilibrium and the actual price levels. In order to impose a unit root, we modify Mussa's specification to eliminate the second term. According to Papell (1997), with the long purchasing power parity[9], so that the equilibrium price level equal to the exchange rae plus the foreign price level, the equation becomes:

$$p_t - p_{t-1} = (\hat{e}_t + \widehat{p}^*_t) - (e_{t-1} + p^*_{t-1}) + \varepsilon_{6t} \tag{6}$$

As Parsley and Popper(1998) once pointed out, the central bank always tries to offset fluctuation caused by the prices and exchange rate shock [10]. Hence we assume that the monetary policy $m_t = m(g_t)$ and that the control variable $g_t = G(e_{t-1})$, where $\frac{\partial m}{\partial g} > 0,\ \frac{\partial g}{\partial e} < 0$.

Hence the money supply becomes

$$m_t = \alpha_4 m_{t-1} + \alpha_5 e_{t-1} + \varepsilon_{7t}\,. \tag{7}$$

Here the coefficient of exchange rate is negative to denote the strategy of the central bank. Such arrangement differs from many tradition literatures, Papell(1993) for example, which suppose that the domestic money supply is also a first order autoregressive process like the foreign variables.Substitute equations (1) ,(2), (4),(5) and (7) into (3) and collect terms,

$$\hat{e}_{t+1} = \frac{\alpha_1 + \theta_1 + \theta_3 - \alpha_5 / \alpha_4}{\alpha_1 + \theta_3} e_t + \frac{1-\theta_1}{\alpha_1 + \theta_3} p_t + \frac{1}{\alpha_4(\alpha_1 + \theta_3)} m_{t+1} + \frac{\theta_1 - \theta_2}{\alpha_1 + \theta_3} p_t^* - \frac{\theta_2}{\alpha_1 + \theta_3} m_t^* - \frac{\theta_2 + \theta_3}{\alpha_1 + \theta_3} r_t^* + v_{1t}$$

which can be denoted in a simple way as follows.

$$\hat{e}_{t+1} = \delta_1 e_t + \delta_2 p_t + \delta_3 m_{t+1} + \delta_4 p_t^* + \delta_5 m_t^* + \delta_6 r_t^* + v_{1t} \tag{8}$$

Based on the above derivation, we can get the expression for the price level by substituting (5),(7) and (8) into (6) [11]

[9] We notice that thePPP is rejected using price indices by Isard (1977) with unit values, by Richardson (1978)in second difference form and for a variety of country pairs. More recently, Froot, Kim, and Rogoff (1995) document deviations from the PPP in commodity data span-ning up to seven centuries. Nevertheless, we still treat the international market in a traditional way for the sake of simplicity.

[10] In that paper, they suppose that the central bank offesets a fixed proportion of the fluctuation casued by the exchange rate. Here we assume the central bank has a more flexible target, which is closer to the reality.

[11] By equations (1) (3) and (4), the output level which is announced quarterly can be expressed as the function of money supply and foreign interest rate which are announced monthly.

$$p_{t+1} = (\delta_1 - 1)e_t + (\delta_2 + 1)p_t + \delta_3 m_{t+1} + (\delta_4 + \alpha_2 - 1)p_t^* + \delta_5 m_t^* + \delta_6 r_t^* + v_{2t}$$

which can also be expressed as follows.

$$p_{t+1} = \lambda_1 e_t + \lambda_2 p_t + \lambda_3 m_{t+1} + \lambda_4 p_t^* + \lambda_5 m_t^* + \lambda_6 r_t^* + v_{2t} \quad (9)$$

If m_t is independent of the exchange rate e_{t-1}, it is straightforward to see that the coefficient of exchange rate in (9) is positive. However, if equation (7) holds, $\frac{\partial m}{\partial g} > 0$, $\frac{\partial g}{\partial p} < 0$, λ_1 might be negative, since $dm(g_t) * de_{t-1} \leq 0$. i.e. the central bank strategy might decrease the effect of exchange rate on price level or even totally change the property.

The method of undermined coefficients transforms the system of stochastic difference equations into a system of deterministic difference equations, which does not generally have a unique solution without further specification. The first two eigenvalues of the system, x_1 and x_2 are the ones of the matrix,

$$\begin{bmatrix} \delta_1, & \delta_2 \\ \delta_1\text{-}1 & \delta_1 + 1 \end{bmatrix}$$

It is straightforward that $x_1 = \delta_1 + \delta_2$ and $x_2 = 1$. The further discussion around the uniqueness of the solution is described in more detail by Papell (1997). The reduced form of cointegrating VAR can be expressed as:

$$\vec{z}_t = \vec{c} + A\vec{z}_{t-1} + B(L)\vec{\omega}_t \quad (10)$$

where $\vec{c}$ is the error correction intercept vector, $\vec{z}_t = (e_t, p_t, m_{t+1}, p_t^*, m_t^*, r_t^*)'$, L is the one-period lag operator and in order to make the zero lag coefficient matrix the identity matrix, the combination of v_t , i.e., $\vec{\omega}_t = (\omega_{1t}, \omega_{2t}, \omega_{3t}, \omega_{4t}, \omega_{5t}, \omega_{6t})'$.

Correspondingly, we can get the Markov-switching vector atuoregressive process model which can be expressed as:

$$\overrightarrow{zm}_t = \vec{c}(s_t) + AM(s_t)\overrightarrow{zm}_{t-1} + BM(L) * \overrightarrow{\omega m}_t \quad (11)$$

where *s* is the regime varible, the intercept vector *c(s)* and variable coefficient matrix *AM* are regime-switching. For details of this model, we will discuss in the next section.

III Data and Methodology

The data for all variables are monthly series. As section 2 showed us, the domestic money supply is one-stage forward compared to other series. The Chinese data are from July 2005 to June 2008 covering the periods since the exchange rate policy reform of the Chinese central bank. The data of Japan and Republic of Korea is from Januanry 1981 to June 2008 covering the periods of major exchange rates fluctuation after the collapse of the Bretton Woods system. In figure 1, we can see the trends of exchange rates of the three countries.

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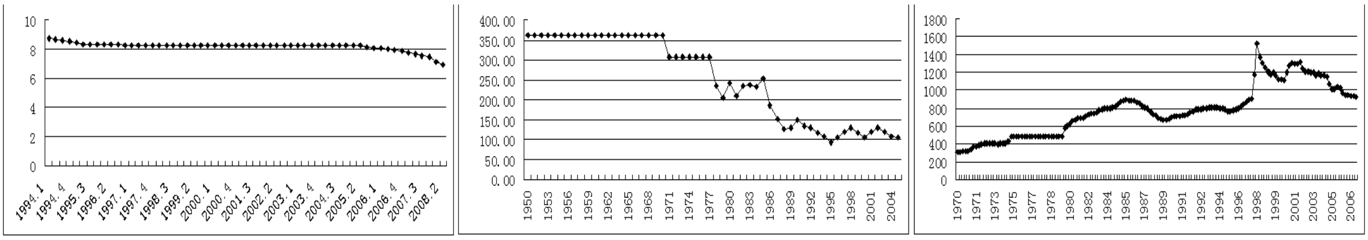


In table 1, we report Augmented-Dickey-Fuller unit root tests for the variables: exchange rate, domestic and foreign (American) prices, money supplies, which enter the reduced form. The ADF tests provide no evidence against the null hypothesis of unit roots since no rejection of the null hypothesis happens.

Table 1 Augmented- Dickey-Fuller Tests For Unit Roots

（MacKinnon (1996) one-sided p-values in parentheses）

$$\Delta z_t = a_1 + a_2 t + a_3 z_{t-1} + \sum_{i=1}^{k} \Delta z_{t-i}$$ [12]

	e	cpi	M1	*Test critical values*		
				1%	5%	10%
China	1.082818	-1.294032	-0.702726	-4.243644	-3.544284	-3.204699
	(0.9999)	（0.8730）	(0.9650)			
Japan	-2.199681	-2.748649	-0.457224	-3.985857	-3.423377	-3.134639
	(0.4877)	（0.2178）	（0.9992）			
ROK	-2.391642	-1.307016	-0.530455	-3.985857	-3.423377	-3.134639
	(0.3832)	（0.8847）	（0.9819）			
US	-3.137010*	-1.243852	-1.71118	-3.985857	-3.423377	-3.134639
	(0.0995)	(0.8993)	（0.7444）			

[12] Following the procedure in Campbell and Perron (1991) we choose the appropriate lag k for the ADF and EG tests. Start with some upper bound on k, say kmax, chosen a priori. Estimate an autoregression of order kmax. If the last included lag is significant (using the standard normal asymptotic distribution), select k=kmax. If not reduce k by one until the coefficient on the last included lag is significant. If none are significant, select k=0. Here we get kmax=2.

*The first column of the United States is interest rate which is a variable in our model as foreign interest rate.

Since all variables have unit roots, it is possible that some of them might have cointegrating relations. Table 2 shows that exchange rates of three Northeast Asia countries do have conintegration relations to their fundamentals. Since the hypotheses of None conintegration are all rejected. Here we use Johansen and Juselius (1990) conintegration tests by trace and Maximum eigenvalue statistics. As we see, in three countries, "none cointegrating relations" trace and maximum eigenvalue statistics are higher than the corresponding critical values although the cointegrating relations which is revealed by trace statistics might differ from those revealed by maximum eigenvalue.

Table 2 Cointegration Tests

$$e_t = a_{20} + a_{21}t + a_{22}cpi_t + a_{23}m_t + a_{24}cpius_t + a_{25}usm_t + a_{26}usr_t + u_t$$

Without trends

		None	*At most 1*	*At most 2*	*At most 3*	*At most 4*	*At most 5*
China	Trace Statistics	203.48*	106.44*	66.96*	36.45*	20.76*	7.28*
	Max-Eigen Statistics	97.03*	39.48*	30.51*	15.68*	13.48	7.28*
Japan	Trace Statistics	189.63*	118.29*	72.74*	39.72*	10.73	3.57
	Max-Eigen Statistics	71.34*	45.54*	33.02*	28.99*	7.16	3.57
ROK	Trace Statistics	115.66*	60.81	31.11	17.42	5.67	4.62
	Max-Eigen Statistics	54.84*	29.71	13.68	11.76	4.77	0.89
Critical vale for Trace Statistics at 0.05 level		95.75	69.81	47.86	20.80	15.49	3.84
Critical vale for Max-Eigen Statistics at 0.05 level		40.07	33.88	27.58	21.13	14.26	3.84

* denotes rejection of the hypothesis at the 0.05 level

**MacKinnon-Haug-Michelis (1999) p-values

With trends

		None	*At most 1*	*At most 2*	*At most 3*	*At most 4*	*At most 5*
China	Trace Statistics	248.88*	130.79*	91.25*	54.96*	24.73	11.23

	Max-Eigen Statistics	118.09*	39.54*	36.29*	30.23*	13.51	11.23
Japan	Trace Statistics	206.66*	135.30*	89.38*	56.36*	26.91*	5.43
	Max-Eigen Statistics	71.35*	45.92*	33.02*	29.45*	21.48*	5.43
ROK	Trace Statistics	147.30*	89.94*	56.31	27.48	13.91	4.62
	Max-Eigen Statistics	57.35*	33.63*	28.84	13.56	9.29	4.62
Critical vale for Trace Statistics at 0.05 level		117.71	88.80	63.88	42.91	25.87	12.52
Critical vale for Max-Eigen Statistics at 0.05 level		44.50	38.33	32.12	25.82	19.39	12.52

* denotes rejection of the hypothesis at the 0.05 level

**MacKinnon-Haug-Michelis (1999) p-values

As is well known, the cointegration between exchange rate and its fundamentals requires error correction term. Suppose that the level of a n-dimensional vector $\vec{z}_t$ can be represented as a nonstationary pth-order vector autoregression: $\vec{z}_t = \vec{c} + A_1 \vec{z}_{t-1} \ldots - A_p \vec{z}_{t-p} + \vec{\omega}_t$ which can be rearranged as:

$$\Delta\vec{z}_t = \vec{c} + \sum_{i=2}^{p} A_i \Delta\vec{z}_{t-1} \ldots + \sum_{i=p}^{p} A_i \Delta\vec{z}_{t-p+1} - A(1)\vec{z}_{t-1} + \vec{\omega}_t \quad (12)$$

where $A(L) \triangleq (I_n - A_1 L \ldots - A_p L^p)$ and $\vec{z}_t$ has k cointegrating relations. Hence equation (12) can be written as

$$\Delta\vec{z}_t = \vec{c} + \sum_{i=2}^{p} A_i \Delta\vec{z}_{t-1} \ldots + \sum_{i=p}^{p} A_i \Delta\vec{z}_{t-p+1} - B\vec{y}_{t-1} + \vec{\omega}_t \quad (13)$$

wehre $\vec{y}_{t-1}$ is an h×1 vector, *B* is an n×h matrix and $-B\vec{y}_{t-1}$ is stationary and also the error correction term of the cointegration system. Taking expectation of both sides of equation (13) results in

$$(I_n - \sum_{i=2}^{p} A_i ... - \sum_{i=p}^{p} A_i) E(\Delta \vec{z}_t) = \vec{c} - BE(\vec{y}_{t-1}) \tag{14}$$

As usual, we assume that the roots *x* of $|I_n - \sum_{i=2}^{p} A_i x ... - \sum_{i=p}^{p} A_i x^{p-1}| = 0$ are all outside the unit circle. Therefore, $I_n - \sum_{i=2}^{p} A_i ... - \sum_{i=p}^{p} A_i$ is nonsingular. In order to represent a system in which there is no drift in any of the variables, viz., $E(\Delta \vec{z}_t) = 0$, the following restriction has to be imposed: $\vec{c} = BE(\vec{y}_{t-1})$. With this restriction, we can build our VAR models with variables in levels.

Another method used in this paper is Markov-switching Vector Autoregressive process. By allowing for changes in regime of the process generating the time series, the MS-VAR model has been proposed as an alternative to the constant-parameter, linear time-series models. The general idea behind this class of regime-switching models is that the parameters of a *K*-dimensional vector time series process $\{\overrightarrow{zm_t}\}$ depend on an unobservable regime variable $s_t \in \{1, 2, ...M\}$, which represents the probability of being in a particular state of the reality.

$$p(\overrightarrow{zm_t} \mid \overrightarrow{ZM}_t, Y_t, s_t) = \begin{cases} f(\overrightarrow{zm_t} \mid \overrightarrow{ZM}_t, Y_t, \gamma_1) & if \quad s_t = 1 \\ f(\overrightarrow{zm_t} \mid \overrightarrow{ZM}_t, Y_t, \gamma_2) & if \quad s_t = 2 \\ \ldots\ldots & \\ f(\overrightarrow{zm_t} \mid \overrightarrow{ZM}_t, Y_t, \gamma_M) & if \quad s_t = M \end{cases} \tag{15}$$

where $\overrightarrow{ZM}_t = \{\overrightarrow{zm_{t-i}}\}_{i=0}^{\infty}$ denotes the history of $\overrightarrow{zm_t}$, Y_t are exogenous variables, and γ_m is the parameter vector associated with regime *m*.

The statistical model requires the formulation of a mechanism that governs the evolution of the stochastic and unobservable regimes on which the above density function depends. In Markov-switching models, the regime-generating process is an ergodic Markov chain with a finite number of states defined by the transition probabilities: $p_{ij} = \Pr(s_{t+1} = j \mid s_t = i)$. And the ergodic M-state Markov process with an irreducible transition matrix

$$P = \begin{bmatrix} p_{11}, & \cdots & p_{1M} \\ \vdots & & \\ p_{M1}, & \cdots & p_{MM} \end{bmatrix}$$

The prediction of MS-VAR processes uses the Markov property of the joint process $\{(\overrightarrow{zm}_t, s_t)\}$ of the regime variable s_t and the stacked vector of the observed variables $\overrightarrow{zm} = (\overrightarrow{zm}_t, \overrightarrow{zm}_{t-1} ... \overrightarrow{zm}_{t-p+1})$, i.e.the relevant information concrning the evolutionof the system output in the future. Conditional on the history of regimes, $S_t = \{s_{t-j}\}_{j=0}^{\infty}$, the density of $\overrightarrow{zm}$ entails the Markov property:

$$p(\overrightarrow{zm}_t \mid \overrightarrow{ZM}_{t-1}, S_t) = p(\overrightarrow{zm}_t \mid \overrightarrow{ZM}_{t-1}, s_t)$$

Since it only depends on the distribution of the error term $\vec{\omega}_t$, which is independent of $\overrightarrow{ZM}_t$, However, the marginal process $\overrightarrow{zm}$ generally is not Markovian.[13]

$$p(\overrightarrow{zm}_t \mid \overrightarrow{ZM}_{t-1}, S_t) = \sum_{m=1}^{M} p(\overrightarrow{zm}_t \mid \overrightarrow{zm}_{t-1}, s_t = m) p(s_t = m \mid \overrightarrow{ZM}_{t-1}) \neq p(\overrightarrow{zm}_t \mid \overrightarrow{zm}_{t-1})$$

The explanation is straightforward. Former observations of $\overrightarrow{zm}_t$ reveal information on the unobservable regime variable which then affects the predictive probability density of $\overrightarrow{zm}_t$.

$$p(s_t \mid \overrightarrow{ZM}_{t-1}) \neq p(s_t \mid \overrightarrow{zm}_{t-1})$$

Only if the regimes are not autocorrelated, that is, $p(s_t \mid s_{t-1}) = p(s_t)$, the Markov property of $\overrightarrow{zm}_t$ would be re-established as $p(\overrightarrow{zm}_t \mid \overrightarrow{ZM}_{t-1}, s_t) = p(\overrightarrow{zm}_t \mid \overrightarrow{ZM}_{t-1})$. In this case the regime variable s_t is said to be unpredictable.

As we saw in section 2, the most general form of a Markov-switching vector autoregressive

[13] This suggests the existence of finite –order mixed VARMA representation of these processes.(Krolzig,1997).

process is given by (10). Here the parameter shift functions $\vec{c}(s_t), AM(s_t)$ and $\overrightarrow{\omega m_t}$ describe the dependence of the parameters on the realized regime s_t. Concretely, in this paper, we find that the most efficient arrangement is to let the intercept and the coefficient of exhcnage rate in (9) to be affected by the regime shift, while all variables are in the first order differences.

$$\Delta p_{t+1} = c(s_t) + \lambda_1(s_t)\Delta e_t + \lambda_2 \Delta p_t + \lambda_3 \Delta m_{t+1} + \lambda_4 \Delta p_t^* + \lambda_5 \Delta m_t^* + \lambda_6 \Delta r_t^* + v_t \quad (16)$$

IV Estimation Results

According to section 3, we know that there is cointegration between exchange rate and its fundamentals. Hence we use Vector Error Correction as the benchmark to investigate the effect of exchange rate pass-through. From table 3 we can see that the exchange rate's change in China will lead the next period price level to move towards the opposite direction, the change in Japan has no obvious effect on the next period price level, while the change in South Korea will lead the next period price level to move towards the same direction.

Table3 Vector Error Correction Estimation Results (t-statistics in brackets)

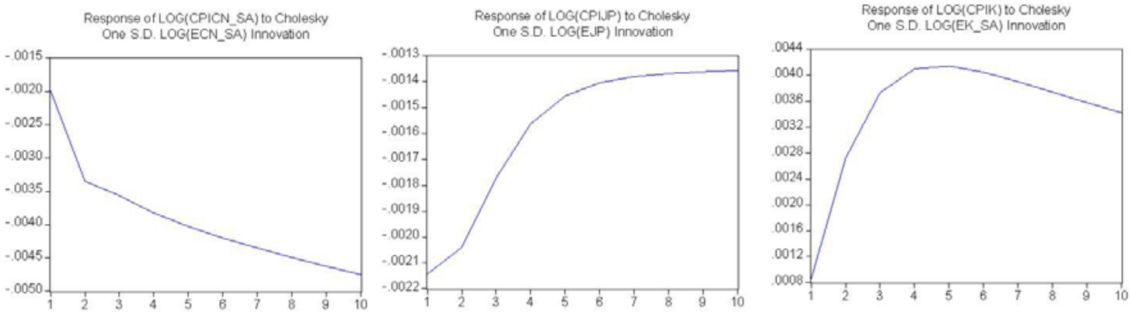
$$p_{t+1} = \lambda_1 e_t + \lambda_2 p_t + \lambda_3 m_{t+1} + \lambda_4 p_t^* + \lambda_5 m_t^* + \lambda_6 r_t^* + v_t$$

	e	cpi	M1	cpius	M1us	rus
China	-0.52869	-0.05291	0.00716	0.02727	0.04873	0.00101
	[-2.34009]	[-0.22872]	[0.05296]	[0.10513]	[0.20602]	[0.18142]
Japan	0.02651	0.2584	-0.00256	0.56248	-0.03200	-3.89E-05
	[1.23961]	[3.10134]	[-0.15693]	[2.35052]	[-0.28083]	[-0.04872]
ROK	0.05997	0.430875	0.00882	-0.00014	0.10155	0.28214
	[3.73382]	[8.83327]	[0.53245]	[-0.23505]	[1.24209]	[1.48655]

Furthermore we also simulate impulse of exchange rate to check the response of the price level for the three countries. As figure 2 shows, in China the price level is going down after the shock of the depreciation of RMB. And the trends seems to get more and more obvious. In Japan, the price level responds in the same direction after the same shock. But the trends is getting more and more unobvious. In South Korea, the price level responds towards a backward way and the trends remains steady afterwards.

Figure 2 Impulse and Rsponse to Cholesky Innovation

China Japan ROK



To make sure that our theory can be supported by empirical evidence, we use MS-VAR model. Comparing many kinds, we eventually choose MS (2)-VAR (1) model to estimate the effects of exchange rate pass-through. Table 4 shows that we can get the similar results as VEC model, viz. the exchange rate's change in China will lead the next period price level to move towards the opposite direction, the change in Japan has no obvious effect on the next period price level, while the change in South Korea will lead the next period price level to move towards the same direction.

Table 4 Markov-switching Vector Autoregressive Estimation Results

< MS (2)-VAR (1) Model >

$$\Delta p_{t+1} = c(s_t) + \lambda_1(s_t)\Delta e_t + \lambda_2 \Delta p_t + \lambda_3 \Delta m_{t+1} + \lambda_4 \Delta p_t^* + \lambda_5 \Delta m_t^* + \lambda_6 \Delta r_t^* + v_t$$

Non Switching Parameters (p-values in parentheses)

	Cpi-1	M1	Cpius-1	M1us-1	Rus-1
China	-0.0830	0.2161	0.2493	0.1561	0.0083
	(0.17)	(0.00)	(0.00)	(0.12)	(0.00)
Japan	0.2326	-0.0510	0.3077	0.0182	-0.0005
	(0.00)	(0.00)	(0.01)	(0.72)	(0.18)
ROK	0.3709	-0.0209	0.2636	-0.0253	0.0003
	(0.00)	(0.04)	(0.01)	(0.51)	(0.44)

Switching Parameters (p-values in parentheses)

	State 1			*State 2*		
	Standard Deviation	Parameters for the Intercept	Parameters for Exchange Rate	Standard Deviation	Parameters for The Intercept	Parameters for Exchange Rate
China	0.0006	-0.0067	-1.6738	0.0005	-0.000	-0.5029
	(0.00)	(0.00)	(0.00)	(0.01)	(0.09)	(0.00)
Japan	0.0090	0.0094	0.0283	0.0045	-0.0004	-0.0041
	(0.00)	(0.00)	(0.61)	(0.00)	(0.45)	(0.71)

ROK	0.0088 (0.00)	0.0060 (0.00)	0.0500 (0.03)	0.0031 (0.00)	0.0019 (0.00)	0.0507 (0.00)

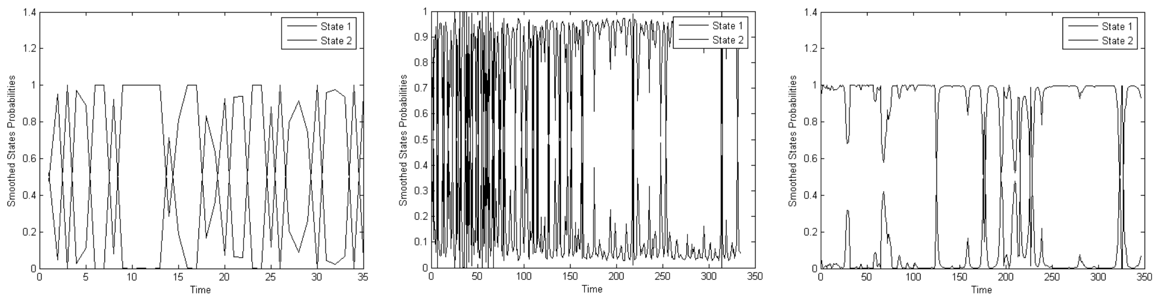
From table 4 we can see that both Chinese and South Korean pass-throughs are obvious while only that in Japan is unobvious as the p-value is 0.61. What's more, the effect is negative in China but positve in South Korea, which result is consistent with that of VEC model and supports our theory in section 2. As the regime variable is accounted, we find the transition probability as follows, from which we can track the change rule of the states.

Figure 3 Transition Probabilities

China		*Japan*		*South Korea*	
0.53 (0.27)	0.58 (0.21)	0.38 (0.10)	0.15 (0.03)	0.95 (Inf)	0.03 (0.02)
0.47 (0.17)	0.42 (0.26)	0.62 (0.13)	0.85 (0.05)	0.05 (0.02)	0.97 (Inf)

* std. errors in parentheses

China Japan South Korea



From the above figures, we can see that there does exist some regime variables which cannot be observed directly. When the state change in a positve probability, the coefficients of variables adjust correspondingly.

V Conclusion

New open economy macroeconomic model provide an appealing framework for examing the exchange rate pass-through to different prices. There is considerable controversy about what features are needed to explain the exchange rate pass-through. Although a number of explanations have been proposed, there has been little work mentioning that the appreciation of exchange rate might lead to inflation instead of deflation. This paper investigates exchange rate pass-through on domestic price level. We think that the interferences in the exchange rate market from the central banks might offest the imports goods price impact which makes the exchange rate pass-through be negative, positve or unobvious. With evidence from Northeast Asia, we show that our theoretical model are supported by the data. In terms of methodology, we use Vector Error Correction and Markov-switching Vector Autoregression respectively and get the consistent results to support our theory.

There is a growing literature that uses new open economy macroeconomic model to measure the degree of exchange rate pass-through on the domestic price level. Based on such foundation, we build a regulated exchange rate market model. The extensive directions of this paper include that 1) use new methodology such as Bayesian Vector Autoregressive model or SETAR etc. 2) further our exploration to cover more parameters such as interest rate, money supply and output ect. 3) explore the determinants, such as control powers of the central bank in different countries, which make the exchange rate pass-through various.

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Reference

Bailliu, J., Fujii, E., (2004). "Exchange rate pass-through in industrialized countries: an empirical investigation". Bank of Canada Working Paper.

Betts, C., Devereux, M., (2001), "The international effects of monetary and fiscal policy in a two-country model". In: Calvo, G., Dornbusch, R., Obstfeld, M. (Eds.), Essays in Honor of Robert A. Mundell, 9–52. MIT Press, London. Cambridge.

Campa, J., Goldberg, L., (2002), "Exchange Rate Pass-Through into Import Prices; A Macro or Micro Phenomenon?" NBER Working Paper No. 8934.

Clements, M.P. and Hendry, D.F.(1999), "Forecasting Non-stationary Economic Time Series: The Zeuthen Lectures on Economic Forecasting", Cambridge, Mass: MIT Press.

Corsetti, G., Pesenti P., (2005), International Dimensions of Optimal Monetary Policy. Journal of Monetary Economics. 53, 281–305.

Dornbusch, R., (1976), "Expectation and Exchange Rates Dynamics", Journal of Political Economy, P1161-1176.

Dornbusch, R, (1987), "Exchange Rates and Prices", American Economic Review, P93-106.

Eaton, J. and Turnovsky, S. (1983), "Covered Interest Rate Parity, Uncovered Interest Rate Parity and Exchange Rate Dynamics", Economic Journal, P555-575.

Engel, C., (2002), "The responsiveness of consumer prices to exchange rates: a synthesis of some new open economy macro models". The Manchester School 70, 1–15.

Feenstra Robert C.（1989）, "Symmetric Pass-Through of Tariffs and Exchange Rates Under Imperfect Competition: An Empirical Test," Journal of International Economics, Aug. 27(1,2), pp. 25-45.

Froot ,K and P. Klemperer, (1989), "Exchange Rate Pass-Through: When Market Share Matters", American Economic Review, P637-654.

Froot, K A.; KIM, Michael and Rogoff, Kenneth（1995）. "The Law of One Price over 700 Years." NBER Working Paper No. 5132, May.

Gagnon, J., Ihrig, J., (2004), "Monetary policy and exchange rate pass-through. Board of Governors of the Federal Reserve System", International Finance Discussion Paper No. 704.

Isard, Peter（1977）, "How Far Can We Push the 'Law of One Price ?" American Economic

Review, 67(5), pp. 942-48.
Johansen, Søren and Katarina Juselius (1990). “Maximum Likelihood Estimation and Inferences on Cointegration—with applications to the demand for money,” Oxford Bulletin of Economics and Statistics, 52, 169–210.
Kichian, M., (2001), “On the Nature and the Stability of the Canadian Phillips Curve”, Bank of Canada Working Paper, No. 2001-4.
Koujianou Goldberg and Knetter（1997）“Goods Prices and Exchange Rates: What Have We Learned”, Journal of Economic Literature, P1243-1272.
Krozig, H. (1997), “Markov Switching Vector Autoregressions. Modelling, Statistical Inference and Application to Business Cycle Analysis,” Berlin: Springer.
Monacelli, T., (2005), “Monetary Policy in a Low Pass-Through Environment”, Journal of Money, Credit, and Banking. 37, 1047–1066.
Mussa, M (1981), “ Sticky Prices and Disequilibrium Adjustment in a Rational Model of the Inflationary Process”, American Economic Review, P1020-1927
Papell, David, (1993), “Exchange Rates and Prices—An Empirical Analysis”, Journal of International Economics,P397-410.
Papell, D (1997), “Cointegration and Exchange Rates Dynamics”, Journal of International Money and Finance, P445-460.
Parsley, D.C and Popper, H.A. (1998), “Exchange Rates, Domestic Prices, and Central Bank Actions: Recent U.S. Experience”, Southern Economic Journal, April, 957–72.
Richardson J. David.（1978）"Some Empirical Evi- dence on Commodity Arbitrage and the Law of One Price," Journal of International Economics, May 8 (2), pp. 341-51.
Smets, F., Wouters, R., (2002), “Openness, imperfect exchange rate pass-through and monetary policy”, Journal of Monetary Economics 49, 947–981.
Sutherland, A., (2005), “Incomplete Pass-Through and the Welfare Effects of Exchange Rate Variability”, Journal of International Economics. 65, 375–399.