Central Bank Intervention and Exchange Rate Dynamics:

A rationale for the regime-switching process of exchange rates

Hsiu-Yun Lee*

Professor, Department of Economics, National Chung Cheng University, Taiwan

and

Wen-Ya Chang

Professor, Department of Economics, Fu-Jen Catholic University, Taiwan

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*Corresponding author. Tel.: 886-5-2720411 ext. 34106, Fax: 886-5-2720816, E-mail: ecdsyl@ccu.edu.tw. Mailing address: Department of Economics, National Chung Cheng University, 168 University Road, Ming-Hsiung, Chia-Yi, Taiwan.

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Abstract

By proposing a stochastic intervention model of exchange rate determination, this paper provides an alternative rationale for the success of the Markov-switching model in explaining exchange rate dynamics. One extreme case is a pure floating rate model while the other extreme one is a driftless random walk model. The relation between the exchange rate and the future fundamentals under a non-intervention state is looser than the one under a pure floating exchange regime. This article also provides a method for detecting a central bank's interventions when intervention data are not available. Applying the stochastic intervention model to the monthly NT$/US$ exchange rates in 1989M1-2004M6, we find that it outperforms both the pure floating rate model and the random walk model in terms of the likelihood value and the diagnostic test of heteroscedasticity. In addition, with the constructed intervention state index in this article, the estimation of the stochastic intervention model is found to be consistent with the hypothesis that the regime switches of exchange rates are due to a central bank's (non-) interventions.

JEL Classification: E58, F30.

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1. Introduction

The most popular exchange rate arrangement nowadays is a regime between undisputed floats and undisputed pegs. In order to capture reality, this paper first proposes a simple model of short-run exchange rate determination with stochastic interventions from a central bank. Under stochastic intervention, the exchange rate sometimes is endogenously determined by market fundamentals, while it sometimes is manipulated by the central bank. The implied state-dependent exchange rate adjustment of a theoretical model can then be shown to conform to the actual exchange rate process for emerging market countries such as Taiwan.

Due most likely to "fear of floating" and "fear of pegging," many emerging market countries that claim they are floating actually manage their exchange rates (ref. Calvo and Reinhart, 2002, and Levy-Yeyati and Sturzenegger, 2002). The stylized facts are the following. Unlike monetary authorities in the European Monetary System or those in several major industrialized countries, the central banks in emerging market countries do not coordinate with other countries to undertake joint currency interventions. Moreover, the central banks in these emerging market countries neither publicly announce the target of their exchange rate policy nor promise under what situation that they will undertake intervention operations. Even if they announce it ex ante, they can undo it ex post. In addition, intervention data are seldom revealed by the central banks of emerging market countries.[1]

In constructing the analytical framework, we consider that the central bank undertakes intervention operations on a case-by-case basis, but the exact timing and

[1] Sarno and Taylor (2001, pp.851-2) provide a detailed discussion for data availability on intervention.

the exact magnitude of the intervention are not known by market participants *ex ante*. These characters make our model distinct from those in Hsieh (1992), Natividad-Carlos (1994), Lewis (1995), and other target-zone models. In these previous papers the central bank follows an explicit target rate or intervention rule, through which the central bank affects the exchange rate by changing monetary aggregates or reserves. The stochastic intervention model in this paper implies a state-dependent adjustment of the exchange rate. In the case of non-intervention, market fundamentals and the probability of the central bank's continual non-intervention shall determine the exchange rate. On the other hand, in the case of intervention the exchange rate is assumed to follow a driftless random walk process as the monetary authority usually pursues a stable value of its home currency in practice.

Because the process of exchange rates displays non-linearity, many empirical studies employ regime-switching models to fit exchange rate data or to forecast future exchange rates (e.g., Engel and Hamiliton, 1990; Engel, 1994; Engel and King, 1999; Bollen et al., 2000; Dewachter, 2001; and Clarida et al., 2003). The research works in Sill and Wrase (1999) and Sarno et al. (2004) appear to be the first approach that provides a rationale for using Markov-switching models in explaining exchange rate dynamics; that is, market fundamentals themselves are regime switching.[2] This article provides an alternative theoretical base for using regime-switching models. In a stochastic intervention model, the exchange rate is sometimes pegged by the central bank and at other times determined by market fundamentals. Therefore, the exchange rate follows a regime-switching process even though the market

[2] Vigfusson (1997), on the other hand, claims that the process of exchange rates in an economy mixed with chartists and fundamentalists can be approximated with a Markov regime-switching model.

fundamentals do not. Furthermore, the relation between the exchange rate and the fundamentals under a non-intervention period is different from the one under a pure floating exchange regime. Specifically, the relation between the exchange rate and the future fundamentals under a non-intervention state is looser than the one under a pure floating exchange rate regime while that between the exchange rate and the contemporary fundamentals under a non-intervention state is closer than the one under a pure floating exchange rate regime. This feature has never been embedded in empirical studies.

In order to find our model's empirical usefulness, we apply our stochastic intervention model to Taiwan data. Specifically, this article derives the theoretically-implied dynamic process of an exchange rate, which is state-dependent, for the empirical time-series analysis. This makes our article distinct from pure empirical models. The (non-) intervention periods for the exchange rate of the New Taiwan dollar (hereafter, the NT dollar) against the U.S. dollar can be estimated from the smoothing probability after estimating the Markov-switching model. Findings from our empirical investigation are consistent with the practice that Taiwan's central bank conducts stochastic intervention in the foreign exchange market. By comparing our stochastic intervention model with two extreme cases, a driftless random walk model and a pure floating rate model, we find that the stochastic intervention model has the highest likelihood value and removes the autoregressive conditional heteroscedasticity effects in the residuals for the depreciation rate of the NT dollar. In addition, the stochastic intervention model has a statistically equal out-of-sample forecasting power as its random walk counterpart. Finally, the process of a constructed intervention state index in this paper is found to be consistent with the non-intervention periods estimated by the smooth probability with the stochastic intervention model.

The remainder of the paper is organized as follows. Section 2 describes the stochastic intervention model. After solving the model, we show that parameters of the depreciation rate are state-dependent and hence provide a rationale for using the Markov-switching model in estimating the exchange rate process. Section 3 begins with the data's description. The state-dependent depreciation rates, together with exogenous market fundamentals, are then examined empirically. Some comparisons among the stochastic intervention model, a random walk model, and a pure floating model are conducted. Finally, a comparison between the estimated intervention periods of the stochastic intervention model and a constructed intervention state index is conducted, too. Conclusions are summarized in the last section.

2. A short-run model of exchange rate determination with central bank stochastic intervention

In order to describe the short-run fluctuation of exchange rate behavior, we assume that: (i) the open economy is specified to be small in the sense that it cannot influence foreign interest rates and foreign prices of its imports; (ii) domestic output remains at its fixed level; (iii) the domestic price is taken as given as it adjusts with a lag; (iv) market participants form their expectations rationally; (v) the monetary authority controls domestic interest rates as a policy instrument.

In accordance with these assumptions, the analytical framework for the short-run exchange rate behavior can be only described by the following constraint of the foreign exchange market:[3]

[3] The theoretical model we shall develop can be regarded as a variant of Frenkel and Rodriguez's (1982) model. It can help us to get a clear and simple relationship between the exchange rate and its

$$\alpha + \beta(p_t^* + e_t - p_t) + \gamma(r_t - r_t^* - E_t e_{t+1} + e_t) = \mathrm{BOP}_t, \tag{1}$$

where α is an autonomous term of trade balance which reflects the net demand for domestic goods due to changes in domestic and foreign expenditures, β and γ are positive structural parameters, p_t^* is the logarithm of the foreign price level at time *t*, e_t is the logarithm of the exchange rate (defined as the domestic currency price of foreign currency), p_t is the logarithm of the domestic price level, r_t and r_t^* denote the domestic and foreign nominal interest rates, respectively, and BOP_t is defined to be the balance of payments. In addition, $E_t x_{t+i} \equiv E[x_{t+i} \mid \Omega_t]$ is the mathematical expectation of any variable x_{t+i} for time $t+i$, based on the information set available at time *t* (Ω_t).

If the central bank lets the exchange rate freely adjust in the foreign exchange market, then the balance of payments is in equilibrium and $\mathrm{BOP}_t = 0$ holds. As in Frenkel and Rodriguez (1982), equation (1) can describe the movement of exchange rate behavior. In the short run, domestic prices are rigid and hence the key factor affecting the current period's exchange rate is the difference between the returns on domestic bonds and those on foreign bonds. In the long run, domestic prices adjust over time and thus the relative price of foreign to domestic goods becomes another key factor in determining the equilibrium exchange rate.

The goal of this paper is to propose a simple framework of a stochastic rational expectations exchange rate with central bank stochastic interventions, which can be further consistent with the time-series features of short-run movements in the exchange rate. Assume that, at any point of time, the central bank stochastically

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fundamentals.

switches between intervening and non-intervening in the foreign exchange market. In the case of intervention we assume that the central bank in general pegs the exchange rate at its last-period level so as to avoid a big jump in the value of its currency.

The behavior of the managed exchange rate thus obeys the following process:

$$e_t^i = e_{t-1} + \varepsilon_t , \tag{2}$$

where e_t^i is the managed exchange rate, and ε_t is a serially-uncorrelated, normal-distributed disturbance with zero mean and variance σ_ε^2. Equation (2) reflects the fact that the central bank may determine its currency to depreciate or appreciate at an acceptable magnitude if the central bank intervenes in the foreign exchange market. However, on average the objective of the central bank's intervention is to keep the value of its home currency intact.

Whether or not the central bank undertakes intervention operations in practice depends on a persistently changing economic environment. Thus, it seems reasonable not to assume that the intervention of the central bank is an independent event. Let q_0 be the probability that the central bank still intervenes in the foreign exchange market at time $t+1$ if it intervenes at time *t*. In the case of intervention the expected exchange rate is the following:

$$E_t^i e_{t+1} = (1-q_0) E_t^i e_{t+1}^n + q_0 E_t^i e_{t+1}^i ,$$

where e_t^n is the market-determined exchange rate, and E_t^i is the expectations operator based on the information set when the central bank intervenes at time *t*.

On the other hand, if the central bank does not intervene in the foreign exchange market, then the market-determined exchange rate (e_t^n) is affected by the interest differential between domestic and foreign interest rates ($r_t - r_t^*$), the price differential of domestic and foreign prices ($p_t - p_t^*$), and the expected exchange rate for the next period. Specifically, equation (1) with $\mathrm{BOP}_t = 0$ gives:

$$e_t^n = (1+\beta')^{-1}[E_t^n e_{t+1} - (r_t - r_t^*) - \alpha' + \beta'(p_t - p_t^*)], \qquad (3)$$

where E_t^n is the expectations operator based on the information set when the central bank does not intervene at time *t*, $\beta' \equiv \beta / \gamma$, and $\alpha' \equiv \alpha / \gamma$. Let q_1 be the probability that the central bank still lets the exchange rate freely adjust at time $t+1$ if it does not intervene in the foreign exchange market at time *t*. In this case the expected exchange rate is as follows:

$$E_t^n e_{t+1} = (1-q_1) E_t^n e_{t+1}^i + q_1 E_t^n e_{t+1}^n .$$

Since $E_t^n e_{t+1}^i = e_t^n$ in the case of non-intervention, the above equation can be rewritten as:

$$E_t^n e_{t+1} = (1-q_1) e_t^n + q_1 E_t^n e_{t+1}^n . \qquad (4)$$

Combining equations (3) and (4) gives:

$$e_t^n = \frac{q_1}{q_1 + \beta'}[E_t^n e_{t+1}^n - \frac{1}{q_1}(r_t - r_t^*) - \frac{\alpha'}{q_1} + \frac{\beta'}{q_1}(p_t - p_t^*)].$$

Solving the equilibrium exchange rate forward, we have

$$e_t^n = c - \frac{1}{q_1 + \beta'} E_t^n \sum_{j=0}^{\infty} d^{-j} \left[(r_{t+j} - r_{t+j}^*) - \beta'(p_{t+j} - p_{t+j}^*) \right], \qquad (5)$$

where $c = -\alpha' / \beta'$ and $d = 1 + \beta' / q_1$. Thus, the key factors affecting the equilibrium exchange rate are the current-period interest differential, the expected interest differential for future periods, this period's price differential, the expected price differential for future periods, and the probability of the central bank continuing its non-intervention.

Based on equation (5), an increase in the interest differential originating from a contraction monetary policy of the home country leads to an appreciation of the home currency. A depreciation of the home currency is associated with an increase in the price differential as a result of a relative slowdown in domestic productivity growth. The discount factor d^{-1} in equation (5) is an increasing function of the probability of staying in a non-intervention period. Therefore, a higher possibility that the exchange rate is kept floating by the central bank increases the effects of the future interest differentials and the future price differentials on the equilibrium exchange rate. In the extreme case where q_1 approaches unity, i.e., the central bank will not intervene in the foreign exchange market for future periods if it does not intervene during the current period, equation (5) turns out to be the standard solution of the exchange rate process under a pure flexible exchange regime. In the other extreme case where q_1 equals zero, i.e., the central bank does not intervene during the current period, but it will peg the exchange rate for future periods, equation (5) reduces to a solution of the exchange rate under the situation where the expected depreciation rate is zero.

We are now ready to construct an econometric model on the basis of theoretical equations (2) and (5), which can fit short-run movements of the exchange rate. We rewrite equations (2) and (5) in terms of their difference forms for empirical purposes since most macroeconomic variables are integrated of order one:

$$e_t^i - e_{t-1} \equiv \Delta e_t^i = \varepsilon_t \,, \tag{6}$$

$$e_t^n - e_{t-1} \equiv \Delta e_t^n = c - \beta'^{-1} X_{t-1} - e_{t-1} - \beta'^{-1} E_t^n \sum_{j=0}^{\infty} d^{-j} \Delta X_{t+j} \,, \tag{7}$$

where $X_t \equiv (r_t - r_t^*) - \beta'(p_t - p_t^*)$. Equation (7) clearly implies a co-integration relationship between the equilibrium exchange rate, the interest differential, and the price differential under the non-intervention situation.

To solve the expected present value of the interest differential and the price differential explicitly, this paper assumes a vector autoregressive (VAR) process for these differenced fundamentals. For illustrative simplicity, the dynamic behavior of differenced data is assumed to follow the VAR(1) process:

$$\Delta Z_t = A_0 + A_1 \Delta Z_{t-1} + U_t , \qquad (8)$$

where $\Delta Z_t \equiv [\Delta(r_t - r_t^*) \ \ \Delta(p_t - p_t^*)]'$, A_0 is a 2×1 scalar vector, A_1 is a 2×2 coefficient matrix, and U_t is a 2×1 serially-uncorrelated error term vector. From equation (8) it is easy to show that $E_t^n \Delta Z_{t+j} = \Sigma_{i=0}^{j-1} A_1^i A_0 + A_1^j \Delta Z_t$, for all $j > 0$. Substituting this resulting relationship into the expectation terms of equation (7), we have:

$$E_t^n \sum_{j=0}^{\infty} \left(\frac{q_1}{q_1 + \beta'} \right)^j \Delta X_{t+j} = c' + [1 \ -\beta'] \left[I - \left(\frac{q_1}{q_1 + \beta'} \right) A_1 \right]^{-1} \Delta Z_t ,$$

where I is a 2×2 identity matrix, and c' is a function of elements in matrices A_0 and A_1.

In order to put the model into empirical investigation, there are two equations that should be estimated simultaneously. The first equation is the specification for fundamental variables as the following VAR(n):

$$\Delta Z_t = A_0 + \sum_{i=1}^{n} A_i \Delta Z_{t-i} + U_t , \qquad (9)$$

where A_0 is a 2×1 constant term vector, and A_i , $i = 1, 2, ..., n,$ are 2×2 coefficient matrices. The second one is the depreciation rate of the home currency, which follows:

$$\Delta e_t = S_t c - S_t \beta'^{-1} (r_{t-1} - r_{t-1}^*) + S_t (p_{t-1} - p_{t-1}^*) - S_t e_{t-1} - S_t \beta'^{-1} E_t^n \sum_{j=0}^{\infty} d^{-j} \Delta (r_{t+j} - r_{t+j}^*)$$

$$+ S_t E_t^n \sum_{j=0}^{\infty} d^{-j} \Delta(p_{t+j} - p_{t+j}^*) + (1 - S_t)\varepsilon_t + S_t \nu_t \,, \tag{10}$$

where ν_t may be interpreted as measurement errors due to using proxy variables for foreign ones with $\nu_t \sim N(0, \sigma_\nu^2)$, S_t denotes an unobservable state variable, $S_t = 0$ is the state of intervention, and $S_t = 1$ is the state of non-intervention. The state S_t follows a first-order Markov chain with the following transition matrix:

$$\begin{bmatrix} \Pr(S_t = 0 \mid S_{t-1} = 0) & \Pr(S_t = 1 \mid S_{t-1} = 0) \\ \Pr(S_t = 0 \mid S_{t-1} = 1) & \Pr(S_t = 1 \mid S_{t-1} = 1) \end{bmatrix} = \begin{bmatrix} q_0 & 1 - q_0 \\ 1 - q_1 & q_1 \end{bmatrix}.$$

The expectations terms in (10) should be consistent with the conditional expectations implied by equation (9).

It is worth mentioning that stochastic intervention from a central bank changes the public’s exchange rate expectations and thus makes the effects of fundamentals on the exchange rate depend on the probability of future intervention. The relation between the exchange rate and the future fundamentals under a non-intervention state is looser than the one under a pure floating exchange regime. Conversely, the relation between the exchange rate and the contemporary fundamentals under a non-intervention state is closer than the one under a pure floating exchange regime. This is a feature that the paper takes up, but which has never been embedded in empirical studies before.

3. Data and empirical results

The theoretical model of exchange rate movements with stochastic intervention in Section 2 suggests that the exchange rate behavior exhibits two different patterns over time, depending on whether or not the central bank intervenes in the foreign exchange market. In the case of intervention the exchange rate follows a driftless random walk process, while in the case of non-intervention it is determined by market fundamentals. A Markov-switching model, provided by Hamilton (1989), is suitable for this empirical task and is constructed by combining two dynamics models via a Markovian switching mechanism. As noted in the previous section, in the state of non-intervention the exchange rate is determined by the expected present value of the fundamentals. The process of exogenous fundamentals therefore must be estimated.

This paper applies Taiwan's data to our empirical study. Taiwan is a developing, small, open economy with a foreign exchange market subject to stochastic intervention by its central bank. The principle of the central bank's exchange rate management is based on the resolution of its Board of Directors such that the NT dollar exchange rate is determined by market forces. However, the central bank claims that when irregular factors result in a market imbalance, it will step in to regulate the market so as to maintain the dynamic stability of the exchange rate. Nonetheless, it neither publicly states the equilibrium value of its currency, nor does it consistently behave in accordance with what it claims.[4] Taiwan's stable economic environment enables us to focus on the exchange rate dynamics without considering much political opposition or social conflict as in some developing countries.

3.1 Data sources and a preliminary analysis of data

Data for exchange rates of the NT dollar against the U.S. dollar (e_t) and data for

[4] For example, after the Plaza Accord in 1985, Taiwan's central bank admitted to an under-valuation of the NT dollar, but it stepped into the foreign exchange market and bought U.S. dollars to stop the appreciation of its currency. During the 1997 Asian financial crisis, Taiwan's central bank announced a solid attempt to defend the NT dollar against massive speculation pressure in September 1997, but surprisingly let the NT dollar float on October 17. The NT dollar depreciated from 28.518

the discount rate of Taiwan (r_t) are obtained from *Financial Statistics, Taiwan District, Republic of China*, while those for Taiwan's general export price index (p_t) come from *Commodity-Price Statistics Monthly in Taiwan Area of the Republic of China*. In this paper U.S. data are used as the foreign data. The discount rate (r_t^*) and the U.S. export price index (p_t^*) come from the *International Financial Statistics, International Monetary Fund*. Due to data availability of the export price index, our data frequency is monthly. The data begin in 1989M1 and end in 2004M6.[5] Based on results of the conventional ADF test, a unit-root hypothesis cannot be rejected at a 5% level for the exchange rate, the interest differential, and the price differential.

3.2 Estimation of the Markov-switching model and its mis-specification test

The lag order of fundamentals in equation (9), which is selected by the Schwarz criterion ex ante, is one. The joint estimation results for equations (9) and (10) are reported in Table 1, indicating that all parameters in the depreciation rate function of the home currency are significant and correct in sign. Therefore, the hypothesis that the exchange rate is affected by fundamentals when the central bank does not intervene, while it follows a driftless random walk process when the central bank does intervene, can be confirmed. In addition, estimates of the transition

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to 30.45 in just three days.

[5] Although Taiwan established its first currency market on February 1, 1979 and the central bank claimed that it adopted a managed floating regime, it kept intensive controls on the foreign exchange market until the mid-1980s. The currency of Taiwan, along with other main Asian currencies, was then forced to sharply appreciate against the U.S. dollar during the period from 1985 to 1988 after the G-5 Plaza Accord of September 1985. To avoid the "structural change" problem, this paper drops data before 1989.

probabilities in the depreciation rate equation are 0.986 and 0.944 for intervention and non-intervention states, respectively, indicating very high persistence in state.

This paper also investigates the appropriateness of our state-space, Markov-switching specification. Under the assumption of no state dependence for the change of the currency value, $q_0 + q_1 = 1$ should indeed hold. We find that the Wald statistic for such a hypothesis is 574.3, which indicates a rejection for the hypothesis at the 1% level of significance. This finding supports the appropriateness of our two-state, Markov-switching specification.

The Markov-switching estimation reveals information about unobserved states. As a matter of fact, economists can use the estimated smoothing probability to get an optimal inference about whether a period is under an intervention state or under a non-intervention one. Figure 1 plots the currency depreciation rate and its smoothing probability estimated from the model. At first glance, we find that even after financial deregulation and releasing its foreign exchange control, Taiwan's central bank undertook intervention operations for most of the sample periods. However, the intensive interventions are not inconsistent with international behavior. Many major industrial countries have taken an even more active role in the management of currencies since the mid-1980s (Dominguez, 1998).

There do exist two periods in the sample when Taiwan's central bank let the exchange rate be determined by market forces. The first is roughly from 1991 to 1993. In the beginning of 1991 Taiwan was recovering from a recession and began rapidly growing, along with its trade balance surplus that began to rise. In the capital account, market participants expected an appreciation of the NT dollar against the U.S. dollar and therefore a huge amount of hot money flowed into Taiwan. Taiwan's central bank appeared to satisfy the internal and external situations and let the exchange rate float.

The other period of non-intervention is estimated in 1997. In the beginning of 1997 Taiwan's economic environment was very much like that in 1991. With high economic growth and low inflation in association with a growing trade balance surplus, Taiwan's central bank kept its exchange rate afloat before mid-1997, but then things changed suddenly in the fall of 1997. The Asian financial crisis began to spread out, with speculative attacks creeping upon the NT dollar. Many efforts to combat them by the Taiwan central bank proved to be ineffective, and the bank had to allow the NT dollar to depreciate in October 1997. Taiwan's central bank basically gave up and let its currency float until the end of the first and the most important speculation of the Asian financial crisis.

Even though the period of the 1997 Asian financial crisis is estimated as being a non-intervention regime, we should be more conservative about this finding. The estimating result of the second non-intervention period may only reflect the low possibility of being in a zero depreciation rate regime. However, a preliminary test with dummies in both the intercept and the variance of the depreciation rate has been conducted and the intercept dummy is found to be insignificant. Nonetheless, the variance dummy indeed improves some explanation power of the model, but it worsens the results of the diagnostic tests.

3.3 Comparisons of a variety of models

To further investigate the explanatory power of our model, this paper considers a pegged exchange rate regime and a pure floating rate regime as well. In the extreme case whereby the central bank always intervenes in the foreign exchange market to keep a stable value of its home currency, we refer to this as a pegged exchange rate regime. In the other extreme case when the central bank never intervenes in the foreign exchange market, we call this a pure floating rate regime.

Under the pegged exchange rate regime, which can be given as the special case of $q_0 = 1$ and $q_1 = 0$, the exchange rate follows a driftless random walk model. This model is particularly interesting since Meese and Rogoff's (1983) pioneering work finds that a naive random walk forecast beats several well-known models of exchange rate determination. The other extreme case of the stochastic intervention model is for $q_0 = 0$ and $q_1 = 1$, where the exchange rate is determined solely by market fundamentals and the model reduces to a pure floating rate model.

The parameters corresponding to the depreciation rate of the NT dollar in the above-mentioned three models are listed in Table 2.[6] Every estimate is significant and correct in sign. In the table we also find that our stochastic intervention model (Markov-switching model) has the highest log-likelihood value for the depreciation rate of the NT dollar, while the pure floating rate model has the lowest one. This evidence suggests that our stochastic intervention model outperforms the pure pegged model (random walk model) and the pure floating rate model in explaining the dynamics of exchange rates.[7]

Some diagnostic tests are shown in the last four rows of Table 2. There are serial correlations in the residuals for the pure pegged model, the pure floating model, and the stochastic intervention model. The failure reminds us of conducting further endeavors to narrow the gap between theoretical implications and the real data. On

[6] In the case of pure floating, the depreciation rate equation and the VAR model of the fundamentals are jointly estimated as in the stochastic intervention model.

[7] Even though a random walk model and a pure floating model are theoretically nested in the stochastic intervention model, they cannot be tested by standard hypothesis tests due to the "nuisance parameter" problem in testing Markov-switching models.

the other hand, the ARCH test reveals significant heteroscedasticity at the 5% level in the estimated residuals of the random walk model and the pure floating rate model in both the short run and the long run. However, there is no ARCH effect in the residuals of the Markov-switching model. Findings from Table 2 imply that the stochastic intervention model captures more features in the exchange rate dynamics than a random walk model and a pure floating rate model do.

Another method of judging the appropriateness of these exchange rate models is to test their out-of-sample predictability. Based on the criteria of the mean absolute errors (MAE) and the mean square errors (MSE), we compare the out-of-sample forecast accuracy of the stochastic intervention model with that of a pure floating rate model and a random walk model. Our results (shown in Table 3) point out that the random walk model and the stochastic intervention model have lower MAE and MSE for all of the forecasting months than the pure floating model does.

This article also uses Diebold and Mariano's (1995) test of the hypothesis that the forecasts from two competing models are equally accurate. The Diebold-Mariano statistic has an asymptotic standard normal distribution. In constructing Diebold-Mariano statistics, consistent estimates of the spectral density of the absolute error differential and the square error differential at frequency zero can be obtained by using the method of Newey and West (1987). It is seen from Table 3 that the Diebold and Mariano (1995) test indicates statistically equal forecasting power between a random walk model and the stochastic intervention model and the lowest power in the floating rate model. Note that the exchange rate follows a random walk process under a pegged rate regime. The evidence that the stochastic intervention model has a statistically equal forecasting power is therefore not surprising since the exchange rate of NT$/US$ in all of the forecasting periods belongs to a pegged rate regime.

3.4 Comparisons of estimated non-intervention period and data

In order to further investigate the driving force of the regime switches of the exchange rates, this article constructs an index for intervention state. This index comes from the change of foreign reserves in a central bank, which is a theoretically proper variable, but empirically an inaccurate proxy for intervention. Sarno and Taylor (2001) point out that monetary authorities' international reserves may change merely due to interest receipts on foreign reserves and/or valuation changes while they are kept intact when the authorities use hidden reserves to intervene.

Interest receipts on foreign reserves are the most important income of the government in Taiwan. We thus use the foreign reserves of Taiwan and the London inter-bank offer rate on 3-month U.S. dollar deposits to estimate the interest receipts and deduct them from the change of foreign reserves to arrive at the adjusted reserve change data. In addition, the foreign exchange reserves (and their change) data reported in the *Financial Statistics Monthly, Taiwan District, Republic of China* have been adjusted according to exchange rate changes. The valuation change problem in using reserves change data as the intervention proxy is less serious for the data reported in the *Balance of Payments Quarterly* than those reported in the *Financial Statistics Monthly* (the latter includes both the change in reserves and the change in valuation). This article thus chooses to use the foreign reserves data revealed in the *Balance of Payments Quarterly* to alleviate the valuation change problem. However, as we have no clue as to the authority's hidden reserves, the constructed intervention data with a change in foreign reserves should be interpreted with caution.

Following the idea of Levy-Yeyati and Sturzenegger (2002) which uses an average of the absolute monthly change in net dollar international reserves relative to the monetary base in the previous month as a classification variable to identify the exchange rate regimes for a country in each calendar year, this paper divides our quarterly adjusted reserves change data by the monetary base in the previous quarter and takes a 5-quarter moving average of it to calculate an intervention state index. Comparing the estimated intervention states of smooth probability with the 5-quarter moving average of the adjusted reserves change data appears to be more suitable than comparing it with the original form of the adjusted reserves change data. Figure 2 plots the intervention state index data. Since the periods of the intervention index being relatively small are close to the estimated non-intervention periods, the process of the intervention state index is consistent with the (non-) intervention period estimated by the smooth probability with the stochastic intervention model. As a consequence, these empirical results confirm the hypothesis that the regime switches of exchange rates are due to a central bank's (non-) interventions.

4. Concluding remarks

In order to model the exchange rate process in a small-open economy with frequent central bank interventions, this paper provides a short-run, central bank stochastic intervention model to explain the state-dependent dynamics of the depreciation rate. One extreme case of the model is a pegged exchange rate regime in which the exchange rate follows a zero-drift random walk process. Another extreme case of the model is a pure floating rate regime in which the exchange rate is determined by market fundamentals. The theoretical implication is that stochastic intervention behaviors of a central bank change the public's exchange rate expectations and thus make the effects of fundamentals on the exchange rate depend on the probability of future intervention. Specifically, under a non-intervention state, the relation between the exchange rate and the future fundamentals is looser than under a floating exchange rate regime while that between the exchange rate and the contemporary fundamentals is closer than under a floating exchange rate regime. This is a feature that the article embeds in the empirical estimation.

With the guide of the theoretical model this article provides a method in detecting a central bank's intervention without any intervention data. That is, we use a regime-switching model for estimating the unobservable intervention and non-intervention states. This is especially important since most of the emerging market countries do not publish their intervention data. Applying Taiwan's data to our stochastic intervention model, we find that our model outperforms both the random walk and the pure floating rate models in terms of the likelihood value of the NT$/US$ depreciation rate and diagnostic tests. In the estimated pegged period, our stochastic intervention model has a forecasting power statistically equal to its random walk counterpart in terms of Diebold and Mariano's (1995) statistics. In addition, with the constructed intervention state index in this article, the estimation of the stochastic intervention model is found to be consistent with the hypothesis that regime switches of exchange rates are due to a central bank's (non-) interventions.

What this article has done is the first step to closely relate time-series analysis to an economic model. Empirical results including the failure of the diagnostic tests for the serial correlation and the possibility of a lack of robustness in the estimated non-intervention period during the 1997 Asian financial crisis have to be improved.

Sarno and Taylor's (2001) marvelous work focuses on the effectiveness of exchange rate intervention - one important topic not addressed by us. In fact, in this article we implicitly assume that all central bank interventions are successful. The exchange rate process under ineffective intervention together with non-intervention is attributed to market fundamentals. However, the process of exchange rates governed by market forces and the process of exchange rates under ineffective intervention (e.g., during the 1997 Asian financial crisis) may not be the same. An extension for distinguishing the two cases constitutes the subject of future research.

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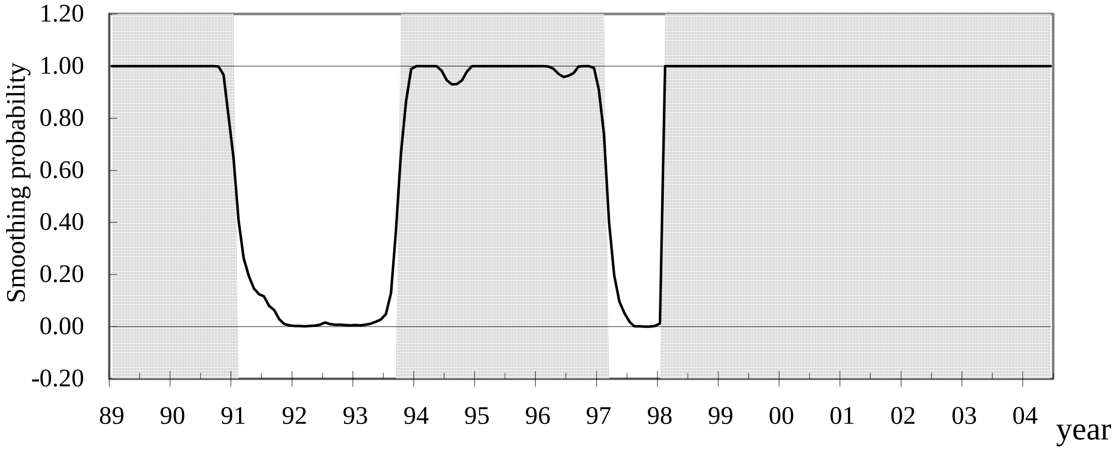
Figure 1: The smoothing probability of intervention and the depreciation rate



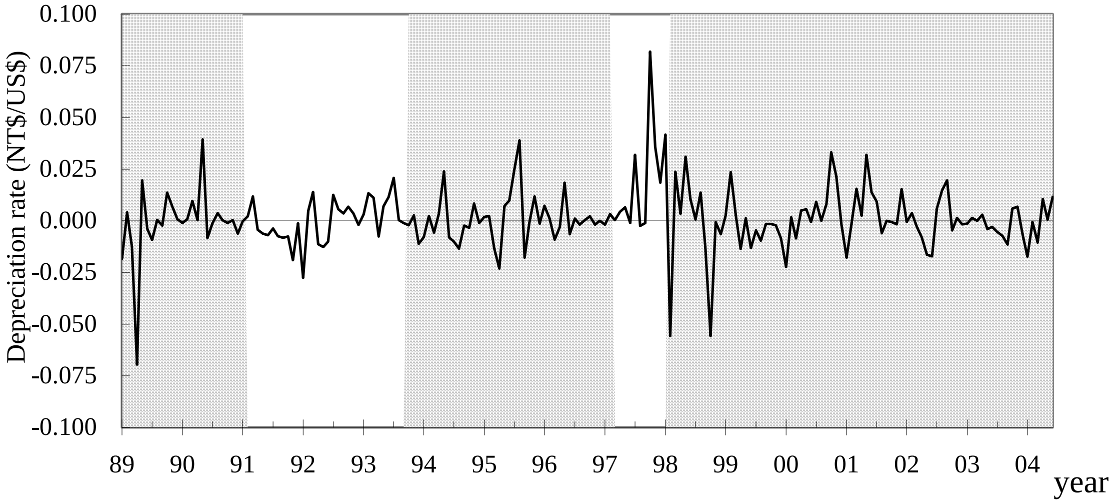


Figure 2: The intervention state index

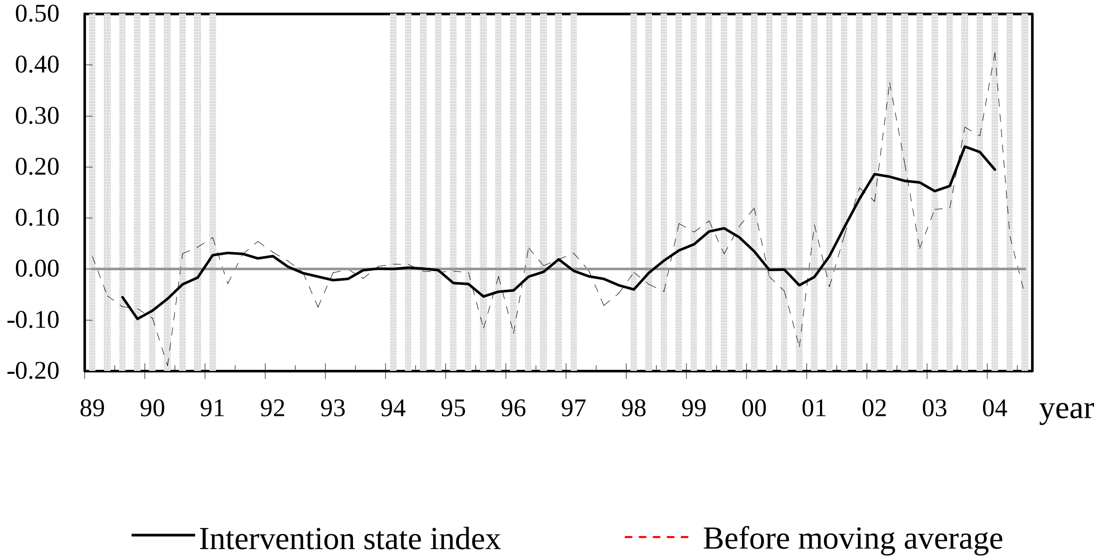


Table 1: Joint estimation of equations (9) and (10) (1989M1-2004M6)

Equation (9): VAR(1) for interest and price differentials					
a_{01}	a_{11}	a_{12}	a_{02}	a_{21}	a_{22}
8e-5	-0.068	8e-4	-3e-4	-0.164	0.240*
(2e-4)	(0.072)	(0.015)	(9e-4)	(0.309)	(0.056)
Equation (10): Depreciation rate of the NT dollar					
c	β'	σ_{ε}	σ_{ν}	q_0	q_1
3.347*	0.169*	0.012*	0.006*	0.986*	0.944*
(0.002)	(0.049)	(7e-4)	(7e-4)	(0.010)	(0.036)

Note: 1. The numbers listed in the table are the estimates of coefficients, with standard errors in parentheses. Term * denotes significance at the 5% level.

2. a_{ij} 's are the coefficients in A_0 and A_1, where $A_0 = \begin{bmatrix} a_{01} \\ a_{02} \end{bmatrix}$ and $A_1 = \begin{bmatrix} a_{11} & a_{12} \\ a_{21} & a_{22} \end{bmatrix}$.

Table 2: Estimation of models under different regimes (1989M1-2004M6)

	Pegged (Random walk)	Pure floating	Stochastic intervention (Markov-switching)
c		3.414*	3.347*
β'		0.364*	0.169*
σ_{ε}	0.013*		0.012*
σ_{ν}		0.083*	0.006*
q_0			0.986*
q_1			0.944*
Log likelihood	536.8	199.0	580.2
Q(1)	38.8*	178.6*	16.6*
Q(24)	114.2*	1987.7*	58.9*
ARCH(1)	11.4*	168.3*	3.6
ARCH(24)	43.9*	156.1*	29.1

Note: 1. In the pure floating and the stochastic intervention regimes, the depreciation rate equation is estimated together with the VAR model of the exogenous variables, while it is estimated alone in the pegged exchange rate regime.

2. "Log likelihood" denotes the logarithm of the likelihood value of the depreciation rate of the NT dollar.

3. Q(n) is the Ljung-Box autocorrelation test statistics for up to nth-order autocorrelation. ARCH(n) denotes the statistics of Engle's Lagrange multiplier test for up to nth-order autoregressive conditional heteroscedasticity. Both have a chi-square distribution with n degrees of freedom. The residuals of the stochastic intervention model are the weighted residuals with a weight equal to the smoothing probability. Term * denotes significance at the 5% level.

Table 3: Out-of-sample forecasting

MAE	Pegged (Random walk)	Pure floating	Stochastic intervention (Markov-switching)
1-month	0.008349	0.140429* [0.0000]	0.008275 [0.3935]
6-month	0.025258	0.148715* [0.0000]	0.024520 [0.7255]
12-month	0.039252	0.157085* [0.0000]	0.039327 [0.9880]
24-month	0.048342	0.157246* [0.0000]	0.048942 [0.9408]

MSE	Pegged (Random walk)	Pure floating	Stochastic intervention (Markov-switching)
1-month	0.000112	0.021093* [0.0000]	0.000112 [0.9418]
6-month	0.000988	0.023102* [0.0000]	0.000970 [0.8505]
12-month	0.002624	0.025559* [0.0000]	0.002563 [0.8548]
24-month	0.003216	0.025080* [0.0000]	0.03227 [0.9880]

Note: MAE denotes mean absolute errors of forecast. MSE denotes mean square errors of forecast. The forecast errors are obtained by a recursive estimation of out-of-sample dynamic forecasts up to k = 1, 6, 12, and 24 months ahead over the period 2000M7-2004M6. Figures in brackets are the P values of Diebold-Mariano statistics comparing the forecast errors with the one obtained by the random walk model. Term * denotes significance at the 5% level.