WHAT CAN UIP TELL US ABOUT EXCHANGE RATE REGIMES? SOME EMPIRICAL EVIDENCE FROM EAST ASIA

by

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Abstract

We follow the recent approach in the literature to use parameters from the uncovered interest parity condition to determine exchange rate regimes in selected East Asian countries. We first find that, in general, movements in these parameters mirror *de facto* exchange rate regime changes following the Asian Financial Crisis. However, the same holds for a group of control countries that neither changed exchange rate regimes nor were part of the Asian Financial Crisis. Further analysis shows that the results are primarily driven by inflation convergence between countries towards the end of the '90s.

Key words: *uncovered interest parity, exchange rate regimes, open economy trilemma*

JEL Classification: *F31, F32, F33*

1. Introduction

Students of open economy macroeconomics learn early about the concept of the *macroeconomic policy trilemma* or the *impossible trinity* (see, e.g. Krugman and Obstfeld, 2006; 629-30, Mankiw, 2007; 359-60). According to it, countries cannot pursue domestic monetary autonomy, a fixed exchange rate regime, and a policy of allowing free capital movement at the same time. Instead, only two of these three options are available to policy makers. Put differently, once a country decides to fix its exchange rate, it then must choose between an independent monetary policy and allowing free capital flows.

If one believes in the effectiveness of active macroeconomic policy, countries potentially have another policy tool available to conduct stabilization policy: fiscal policy. However, its application is limited by the presence of long inside lags and a wide-spread unwillingness to raise tax rates. In addition, fiscal policy for stabilization purposes is infrequently employed due to the presence of large budget deficits in many countries. This suggests the use of an independent monetary policy to conduct active policy in the presence of free capital mobility. The European Union, the U.S., the U.K., and Canada come to mind as examples in having chosen this option, although countries within the European Union earlier opted, at times, for a fixed exchange rate regime with capital mobility.

But do *all* countries which allow for their exchange rate to be flexible really gain monetary freedom? Assume that monetary policy is set through a short-term interest rate target and that capital markets are tightly integrated. Ignoring the risk premium for the moment, or assuming at least that it does not vary, then setting the domestic short-term interest rate to a level different from that of a base rate (the U.S. treasury bill, say) should result immediately in an expected change in the exchange rate with respect to some base rate. If the domestic country cannot stomach large exchange rate swings resulting from setting domestic interest rates independently, then even those countries that have a flexible exchange rate cannot really conduct independent monetary policy. This phenomenon is referred to as the “Fear of Floating” (Calvo and Reinhart, 2002). In its presence, the trilemma is reduced to a single choice: to fix or not to fix the exchange rate. Selecting one or the other has no implication for monetary freedom, and as a result, the choice of exchange rate regime does not depend on the nation’s desire to pursue an independent monetary policy.

In general, world financial markets have become increasingly integrated across borders. This has important and testable policy implications, especially for smaller and lesser developed countries. In particular, one would postulate that irrespective of the chosen exchange rate regime (and there are more than simply the fixed versus flexible variety), there is no difference in the behavior of domestic interest rates with regard to the base interest rate. What is the evidence? Using the uncovered interest parity (UIP) relationship and panel data, Frankel *et*. *al*. (2000, 2002) find evidence supporting the fear of floating hypothesis, while Shambaugh (2004) rejects it.

The purpose of this paper is to shed further light on this controversy. We do so by using the Asian Financial Crisis of 1997-1998 as a quasi experiment. Many of the East Asian countries included in our sample (Indonesia, Korea, Malaysia, the Philippines, Thailand) changed their exchange rate regime for the post-crisis period while others (Hong Kong, Singapore) did not. Accordingly we would expect corresponding changes in certain parameters of the UIP that is consistent with this behavior. The evidence for the countries in our sample at first seems to support Shambaugh's conclusion in general: countries with more (less) flexible exchange rates appear to have a more (less) autonomous monetary policy. However, we observe the same pattern of change in parameters for countries that were not part of the Asian Financial Crisis, and that did not change exchange rate regimes as a result of it (Australia, New Zealand, Canada).[1] We interpret this result as implying that UIP is not really well suited to test for the type of exchange rate regime and that it cannot tell us as much about the effect of these regimes on monetary policy as is commonly assumed in the literature. Instead the results we observe over the sample period are more consistent with inflation convergence between countries.

The paper proceeds as follows: section 2 presents the relationship between UIP and exchange rate regimes, and related specification issues. In section 3, using monthly data, we start our empirical analysis by looking at a single entity, Korea, a country which is believed to have moved to a more freely floating exchange rate regime following the Asian Financial Crisis. Section 4 analyzes data from a panel of East Asian countries and a control group of countries using both annual and monthly data. A final section concludes.

2. UIP and Its Relationship to Exchange Rate Regimes

Following Frankel *et*. *al*. (2000, 2002) and Shambaugh (2004), we will focus on the UIP relationship to gain insights into whether countries with more flexible (fixed) exchange rates have a higher likelihood to pursue a "more (less) independent" monetary policy. A more independent monetary policy here is defined as having the ability to set and to

[1] From here on, we will use the following symbols for countries: Indonesia – IDN, Korea – KOR, Malaysia – MYS, the Philippines – PHL, Thailand – THA, Hong Kong – HKG, Singapore – SGP, Australia – AUS, New Zealand – NZL, Canada – CAN.

move the domestic interest rate (R_t) independently and at a level which differs from the base rate (R_t^f), i.e. $R_t \neq R_t^f$ and $\Delta R_t \neq \Delta R_t^f$.[2] Generally, nominal interest rates could be different between countries for reasons other than monetary policy. There is also no reason to expect that changes in these interest rates should always be the same even under a fixed exchange rate system.

The difference between the two interest rates can be described by two conditions, covered interest parity (1) and UIP (2).

$$R_t = R_t^f + (f_{t+n} - s_t) \tag{1}$$

$$R_t = R_t^f + E_t(s_{t+n} - s_t) + rp_t \tag{2}$$

where f_{t+n} is the forward exchange rate for n periods into the future, s_t is the current spot rate (both exchange rates in logs),[3] and rp_t is the risk premium. The expected exchange rate and the risk premium are unobservable variables.

Both parity conditions basically assume for the law of one price to hold. The connection between (1) and (2) can be seen through the following decomposition (Frankel, 1991):

$$R_t - R_t^f - E_t(s_{t+n} - s_t) = (f_{t+n} - E_t s_{t+n}) + [R_t - R_t^f - (f_{t+n} - s_t)] \tag{3}$$

The first term on the r.h.s. is often referred to as the currency risk premium. The second term is the country or political risk premium. The latter is zero if CIP holds exactly. Hence if CIP cannot be rejected, then rejection of UIP means that forward rates do not equal expected future exchange rates.[4]

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[2] These are nominal interest rates on similar assets and measured in local currency.

[3] Throughout this paper, the exchange rate is quoted as the domestic price of foreign currency.

[4] For a survey of these results, see Froot and Thaler (1990) and Engel (1995).

In principle, we could use equation (1) to determine the exchange rate regime prevalent in a given country. This would involve running a regression of the following type:

$$R_t = \beta_0 + \beta_1(R_t^f + f_{t+n} - s_t) + u_t \tag{4}$$

where in the case of a fixed exchange rate regime we would expect β_0 and β_1 to be zero and one respectively coinciding with a high regression R^2. In the case of a flexible exchange rate regime, we would expect to find no relationship between the domestic and base rate plus forward premium variable. There are complications for inference if countries follow the same "independent" monetary policy following a common shock, but we leave this for later discussion.

Most studies focusing on equation (4) have used data from industrial countries and have been more interested in testing for CIP rather than choosing it as a tool to classify exchange rate regimes. Few studies on CIP have looked at developing countries (de Brouwer, 1999, and Bansal and Dahlquist, 2000). We see the reason for this in the absence of sufficiently liquid forward foreign exchange markets in developing countries. Moreover, even when they exist, data is not easily available.[5] Regardless, Willett, *et al*. (2002) point out:

> "(S)ubstantial deviations from covered interest parity are a good indication that capital mobility is less than perfect..[However] ... [f]inding that covered interest parity holds … is consistent with either high or low capital mobility, and there is no good reason to presume that the magnitudes of deviations from interest parity will provide a reasonable proxy for the degree of international capital mobility. In terms of modern theory, the appropriate measure of capital mobility is the extent to which uncovered rather than covered interest parity holds." (424-5)

[5] We have been unable to replicate Edwards and Khan's (1985) study and would love to hear from anyone in the profession who has been able to do so and is willing to provide us with the data. See also Ahn (1994).

As a result of these arguments, UIP is used to extract information about exchange rate regimes. Equation (4) is therefore replaced by:

$$R_t = \beta_0 + \beta_1[R_t^f + \mathrm{E}_t(s_{t+n} - s_t)] + u_t \tag{5}$$

where either the intercept or the error term may contain the risk premium, depending on one's view of it.

In a fixed exchange rate regime ($\mathrm{E}_t(s_{t+n}) = s_t$), and testing, at first glance, should be straightforward. With flexible exchange rates, equation (2) is solved for the expected change in the exchange rate. Hence differences between the domestic and base interest rate drive expected changes in the exchange rate. In this case, β_1 should be close to zero and the regression R^2 should be low, unless, of course, there is fear of floating or the presence of capital controls.[6]

An alternative route for using UIP to determine exchange rate regimes is to find proxies for ($\mathrm{E}_t(s_{t+n} - s_t)$) and to include these in UIP regressions. The following four measures have typically been suggested in the literature:

(i) perfect foresight $\quad E_t(s_{t+n} - s_t) = s_{t+n} - s_t$

(ii) extrapolative expectations $\quad E_t(s_{t+n} - s_t) = s_t - s_{t-n}$

(iii) static expectations $\quad E_t(s_{t+n} - s_t) = 0$

(iv) survey data $\quad E_t(s_{t+n} - s_t) = s_{t+n}^{survey} - s_t$

[6] Note, however, that "statistical studies of the relationship between interest rate differences and later depreciation rates show that the interest rate difference has been a very bad predictor, in the sense that it has failed to catch any of the large swings in exchange rates. Even worse, the interest difference has, on average, failed to predict correctly the *direction* in which the spot rate would change." Krugman and Obstfeld (2007; 596).

In (iv), survey data is obtained by interviewing market participants. This is regularly done by entities such as the Economist Intelligence Unit (EIU) Currency Consensus Forecast.[7]

Using any of the assumptions about expected exchange rate changes, we could now estimate equation (5) were it not for some further complications. The first involves the classification of exchange rate regimes. Then there are potential econometric estimation problems frequently encountered in time series analysis.

2.1 Exchange Rate Classification

One way to use equation (5) is to test whether or not countries with a more flexible exchange rate system conduct monetary policy more independently than those with a fixed exchange rate regime. To do this you first have to identify country specific exchange rate regimes as “more flexible” (nonpegged country) or “more fixed” (pegged country). Equation (5) can then be estimated separately for both types of regimes, with the expectation that the pegged countries have a slope closer to unity and a higher regression R^2 when compared to the nonpegged countries. For example, the term ($\mathrm{E}_{\mathrm{t}}(s_{t+n} - s_t)$) does not only vanish under a fixed exchange rate regime, but also for a credible peg. Both Frankel *et.al*. (2000, 2002) and Shambaugh (2004) use pooled regression results (among other techniques) to see if there is empirical relevance to the fear of floating argument.

Alternatively, and more interestingly for our set of countries, you can choose sample periods when some countries are commonly assumed to have switched exchange rate

[7] Examples are Frankel and Froot (1987, 1989), Taylor (1989), MacDonald and Torance (1990), Cavaglia et al. (1993), Chinn and Frankel, (1995). MAS (1999) uses four Asian exchange rate markets.

regimes, while others have not. The slope and goodness of fit parameters in equation (5) should then change correspondingly. As in the case discussed in the previous paragraph, the analysis here requires a specification of exchange rate regime. Most commonly, the literature distinguishes between the International Monetary Fund *de jure* classification and a *de facto* classification based on techniques developed by Calvo and Reinhart (2002) or Reinhart and Rogoff (2004). The IMF identifies three categories and 15 further subcategories between 1975 to 1998. Since 1999, it introduced a new classification which also takes into account actual rather than declared behavior.[8] *De facto* classifications are derived from measures typically related to exchange rate and reserve movements, and exchange rate movements outside certain bands. We will use both *de facto* and *de jure* classifications below.

2.2 Econometric Issues: Levels vs. Differences

Having settled the issue of exchange rate regime classification in the previous section, we could simply estimate equation (5) by OLS. Assuming that the base rate is exogenous, that there are no large outliers in either the dependent or explanatory variables, and that there are no omitted variables, then the OLS estimator is consistent if the interest rate variables have a stationary distribution. This result holds irrespective of whether or not the errors are autocorrelated. Inference is valid, as long as heteroskedasticity- and autocorrelation-consistent (HAC) standard errors are used.

[8] The basic three categories of the earlier specification were pegs, limited flexibility, and more flexibility. The newer eight categories were: exchange rate arrangement with no separate legal tender, currency board arrangement, conventional pegged arrangement (peg against a single currency or a basked of currencies), pegged exchange rate within horizontal bands, crawling peg, crawling band, managed floating with no pre-announced path for exchange rate, and independently floating.

Assuming for the moment that there are no structural breaks during our sample period, then for the domestic and base interest rates to have a stationary distribution requires that neither one of the two series contains a stochastic trend. In other words, we need for both variables to be integrated of order zero (I(0)). If, on the other hand, both were I(1), then inference is problematic, since the *t*-statistics are not normally distributed (even in large samples), and there may be a spurious regression problem (Granger and Newbold, 1974; Phillips, 1986; for an excellent summary see, Stock and Watson (2007), chapter 14).

There are different paths out of this situation in the case of I(1) interest rates. One is to estimate an equation specified in differences, i.e.

$$\Delta R_t = \beta_1[\Delta R_t^f + \Delta \mathrm{E}_t(s_{t+n} - s_t)] + u_t^* \tag{6}$$

In addition, if the two interest rates are cointegrated, then an Error (Equilibrium) Correction Mechanism (ECM) form should be employed. The null hypothesis here is that interest rates are more likely to be cointegrated for pegged countries than for others.

We would be somewhat less concerned about this issue if inference regarding the fear of floating and monetary independence was robust irrespective of using equation (5) or (6). Unfortunately, this is not the case. The issue also separates Frankel *et.al*. (2000, 2002), who estimate levels equations, from Shambaugh (2004), who prefers some form of differenced variable estimation. Both come to different conclusions regarding the fear of floating argument and Shambaugh (2004; 314) in particular feels that this is primarily due to estimating (6) rather than (5). Frankel *et.al*. (2002) do not ignore the potential problem resulting from dynamics and address it by stating that they look at equation (5) as an equilibrium expression, although they do not perform a cointegration test for their monthly panel. Furthermore, they argue that “a priori we would expect interest rates to be I(0) variables,” (Frankel *et.al*., 2002; 12).

While we lean more towards the idea that nominal interest rates, real interest rates, and inflationary expectations are stationary, especially over longer periods and for countries that did not experience episodes of hyperinflation,[9] we will estimate both equations (5) and (6) below to contrast potential differences and sensitivities in the conclusions.

3. Estimation Results: Korea

We begin with our empirical investigation by looking at a single country, Korea. The sample period is January 1990 (1990:01) to June 2003 (2003:06). We focus on a single country first because many of the phenomena which we observe for our group of countries can be analyzed in more detail using a single entity. We picked South Korea for this section because it is one of the most developed countries in our sample of seven East Asian countries and it experienced a change in exchange rate regime.

3.1 Choice of Exchange Rate Regime Classification: Korea

Korea’s exchange rate regimes is classified as follows: during the pre-crisis period of our sample Reinhart and Rogoff (2004) list Korea as having a pre announced crawling band until November 1994, and a crawling peg to the U.S. dollar until November 1997. For the period during which Korea was heavily affected by the Asian Financial crisis (December

[9] For the U.S., we can reject the presence of a unit root in the CPI inflation rate for the sample period 1974:I-2002:IV at the 1% level, and for the sample period 1947:I-2002:IV at the 0.1% level using an Augmented Dickey-Fuller (ADF) test. Furthermore, the ADF test has low power, and by using the more powerful DF-GDL test (Elliott, Rothenberg, and Stock, 1996) strengthens this result.

1997 - June 1998), they list it as “freely falling,”[10] followed by a “freely floating” exchange rate regime (July 1998 to December 2001). The *de jure* classification used by the IMF is “managed floating” until November 1997. Since we found that results for East Asia are very sensitive to the sample beginning and end points due to outliers surrounding the Asian Financial Crisis, we looked at the exchange rate series more carefully and found considerable movement during October and November 1997. To avoid these outliers and their consequences for OLS properties, we chose September 1997 as the last observation of our pre-crisis period. As for the starting point of the post-crisis period, we settled on January 1999. Although the *de facto* classification has the won as “freely floating” starting half a year earlier, this classification is based on exchange rate behavior only. However, there was much variation in reserves beginning in July 1998, which suggests that the central bank intervened heavily in the exchange rate market. Still, there is little doubt that the won has been much more flexible during the post-crisis period than during the pre-crisis period (McKinnon and Schnabl (2003), Hernandez and Montiel (2003), Willett and Kim (2004), Kim, Kim, and Wang (2005)).[11]

There is, of course, the additional issue of capital controls, which has a significant effect on the UIP relationship and its ability to signal monetary independence. Shambaugh (2004) allows for the interaction of capital controls with the base rate variable, and we would like to do the same. However, there is no monthly capital control index available to our knowledge. Even if we had such an index at our disposal, most observers date the opening of capital markets for Korea to 1995. We doubt that

[10] This category is used for countries with a 12-month rate of inflation above 40%.

[11] For a comparison of crisis/normalcy performance of UIP, see Flood and Rose (2002).

interacting the two variables would have much of an effect from then on until the end of our pre-crisis period in September 1997. As a result, we ignore the issue for now.

In summary, we settle on classifying the pre-crisis period as “intermediate” and the post-crisis period as “floating.” We believe that this is consistent with the *de jure* and the *de facto* classification.

3.2 Estimation Results Korea

What really matters for the purpose of our analysis is that Korea moved from a peg (pre-crisis) to a nonpeg (post-crisis). Hence we expect the slope to be closer to unity in the earlier period with a coinciding higher regression R^2. Formally

$$R_t^{KOR} = \beta_0 + \beta_1 R_t^{US} + u_t \text{ or } \Delta R_t^{KOR} = \beta_0 + \beta_1 \Delta R_t^{US} + u_t$$

and

$$H_0 : \beta_1^{peg} > \beta_1^{nonpeg} ; R_{peg}^2 > R_{nonpeg}^2$$

Table (1) presents the regression results for both levels equations (column (1)) corresponding to equation (5) above, and the difference specification (column (2)) corresponding to equation (6). In both cases we assume static expectations as presented above. Columns (3) and (4) are Shambaugh’s panel results using differences, both for his entire sample and for the 1990s only. We entered his slope coefficient for the pegs in the “pre-crisis” row and for the nonpegs in the “post-crisis” row.

The levels equation (column (1)) produces the expected results for the slope coefficients. For Korea, the size of the slope coefficient decreased for the nonpeg period when compared to the peg period. For both periods, the slopes are statistically significant, and the slope for the pre-crisis period is also significantly different from unity. The slope coefficients are pleasantly close to Shambaugh's results for the entire estimation period (column (3)), although Shambaugh does not find a statistically significant slope for the '90s (column (4)). The Korean surprise is the regression R^2, which increases dramatically for the post-crisis period despite a decrease in the slope coefficient. Hence we cannot reject the null hypothesis for the slope, but we reject it for the regression R^2.

Although the results in column (2) are closer in spirit to column (4), there is an extraordinary high slope coefficient during the peg period. It is roughly three times as large as the one found in column (4) for the peg. However, it is not statistically significantly different from unity. Even so, the regression R^2 is surprisingly low. This coefficient drops to half the size in column (4) for the nonpeg period, although it remains statistically significant.

When faced with this type of result, the first inclination is to check for data entry errors. Having ruled out this possibility, it was comforting to find another study which listed a slope coefficient of similar magnitude: Kim and Lee (2004) show an even larger slope coefficient (2.946) for a similar sample period using a different estimator. While one can set out to search for theoretical explanations of coefficients of this magnitude (see, e.g. Shambaugh, 2004; 306) it is often useful to plot the data. Figures 1a and 1b present the scatter plot and the regression line for both the pre- and post-crisis period. Note that both are drawn to the same scale.

The reason for the high regression R^2 is that there was much less variation in Korean interest rates during the post-crisis period when compared to the pre-crisis period. The base rate variation was roughly the same. However, the reason for the higher Korean variation seems to be the result of higher average interest rates during the pre-crisis period. It is well known that the variance of the inflation rate increases with its level. To us it seems that the Korean result is driven more by inflation convergence than by a switch from a peg to a more flexible exchange rate regime. This may also explain the size of the slope coefficient in column (2): a given change in the base rate produces much more of a variation in the domestic rate.

To be fair, we need to point out that, following Frankel *et.al*. (2000, 2002), we have used monthly data, while Shambaugh (2004) works with annual (panel) data. Our sample period is much shorter than his, and Shambaugh "hope[s] the dynamics have largely settled" so that he "can pool the data across countries." (312) We could pursue this issue further here, but find it distracting from our main argument. Appendix 1, however, produces tests for cointegration, a dynamic OLS (DOLS) estimator to identify the cointegrating vector and the speed of adjustment coefficient, a general to specific specification search to find the ECM representation, the use of different expectations hypothesis, and allows for dynamics to settle to an equilibrium over a longer period. The upshot of this analysis is that, according to the tests based on UIP, the post-crisis period rather than the pre-crisis period appears to have been a peg.

4. Multiple Country Analysis

The conclusion from section 3 was that, quite surprisingly, UIP estimation does not suggest that Korea followed a more flexible exchange rate regime in the post-crisis period when compared to the pre-crisis period. However, this does not make sense since there is strong evidence from other sources that it did. We therefore must question the ability of UIP to identify exchange rate regimes. Perhaps there are other reasons for observing the results we did for Korea.

Of course, we could simply state that there is an interesting story behind every observation, and that theory reveals itself in data when we looking at averages for a panel of countries. Indeed this is one of the main assets of comparative economic performance. We will do so in this section by using data for seven East Asian countries (HKG, IDN, KOR, MYS, PHL, SGP, THA), and three control countries that are believed to have had a freely floating exchange rate regime throughout the sample period (AUS, CAN, NZL). Note that the three non-Asian countries were not as much affected by the Asian Financial Crisis as the seven East Asian countries.

Table 5 presents *de jure* and *de facto* exchange rate classifications for the seven East Asian countries for the pre- and post-crisis period. There is much variation in the behavior between countries. Four countries moved from a more pegged situation towards a more floating regime (IDN, KOR, PHL, THA), one went in the opposite direction (MYS), while two remained the same (HKG, SGP).

Table 6 presents the estimation results using annual data for both levels and differences. Focusing on the levels results first, we find the following. For the two countries which were expected to maintain the same slope coefficient, it actually declined. The slope for MYS was expected to increase, and it did. Three of the four countries for which the slope coefficient was supposed to decrease, show a decrease. However, all three free floating developed countries, which were not directly affected by the Asian Financial Crisis, have a substantial decrease in the slope coefficient. Similar results hold for the differences.

Having estimated these equations using annual data, we have to confess that we do not have much faith in regression coefficients obtained from only four observations[12] or so. We simply use these as a descriptive device and would rather not talk about goodness of fit (or inference for that matter). To overcome the small sample size problem, we re-estimated the coefficients using monthly data. Results are reported in Table 7.

Looking at the slope coefficients for the levels regressions first, we find that HKG, IDN, KOR, THA, and MYS follow the expected pattern between the pre- and post-crisis period. However SKP, CAN, NZL, and AUS all show significant decreases in the slope coefficients, with PHL entering with an incorrect sign. The regression R^2 display a mixed pattern, with most of them going against the null hypothesis. Moving to the difference equations next, we observe a coefficient decrease as expected only for PHL and KOR, and only the latter is statistically significant. However, the slope coefficient also decreases for HKG, CAN, NZL, and AUS (again, significantly so only for AUS). Only Korea's goodness of fit decreases as expected. For most countries it increases for the post-crisis period.

We now have a dilemma. We would like to base our conclusions on the monthly data estimates of Table 7, but are worried that, in Shambaugh's words, "dynamics have not settled," and as a result we should report the results from the annual sample period, or perhaps averages across countries. However, this does not make sense when you only have very few annual observations available. But what if the annual coefficients, for which dynamic adjustments have been completed, are similar to the slope coefficients

[12] The phrase "Are you kidding me?" comes to mind if we talked about inference now.

based on the monthly data sample? To investigate this possibility, we ran the following regression for the level and difference regressions combined.

$$\widehat{\beta}_1^{monthly} = \text{-}0.33 + 0.88\ \widehat{\beta}_1^{annual} + e_t$$
$$(0.08)\quad (0.07)$$
t = 1, …, 20, R^2 = 0.88

It does appear that, on average, the slope coefficients from the monthly sample period estimation are very similar to those from the annual sample period. Figure 2 confirms this impression. Hence, and for the purpose of the problem at hand, it seems all right to base the analysis on the monthly data analysis.

Taking the evidence from both annual and monthly data together, as well as levels and differences, it seems fair to say that slope coefficients on average decreased. More importantly, this is the case whether they were supposed to do so when countries moved from a peg to a nonpeg, or remained on the same type of regime. Most damaging to the idea that UIP can tell us something about exchange rate regimes is the fact that slope coefficients of the floaters AUS, NZL, and CAN also decreased. This seems to suggest that we should search for an explanation elsewhere. The question then becomes, what is it that drives these results?

One possible explanation is that there has been inflation convergence throughout the '90s, which would explain the pattern of the slope coefficients concerned. Note that

$$R_t = \beta_0 + \beta_1 R_t^f$$
$$(\rho + \pi)_t = \beta_0 + \beta_1 (\rho + \pi)_t^f$$
$$\Delta p_t = (\beta_0 + \beta_1 (\rho_t - \rho_t^f)) + \beta_1 \Delta p_t^f$$
$$\Delta p_t = \beta_0^* + \beta_1 \Delta p_t^f$$

where ρ is the real interest rate, and π represents inflationary expectations. To get to the final expression, we have assumed that real interest rates are equal between countries (or at least that their difference remains the same over time), and that agents have static expectations.

Now consider a situation where inflation during the pre-crisis period is substantially higher in the domestic country when compared to the base country. Furthermore, think of the inflation rate falling for both countries, but more so for the domestic inflation rate than for the inflation rate in the base country. For illustration purposes, let the inflation rate for the domestic economy fall by five percentage points, while the base country inflation rate fell by 2.5 percentage points during the pre-crisis. This would result in a slope coefficient of -2.[13] Although inflation rates continued to fall during the post-crisis period, they did not fall by as much anymore, since they were already at a lower level at the beginning of the post-crisis period. Again, for illustration purposes, consider an addition fall of 1.5 percentage point for the domestic inflation rate level, and a 1 percentage point change in the base country. In that case, the slope coefficient would be 1.5, i.e. it would have fallen. Table 8 presents the pre-crisis and post-crisis inflation rates for the countries in our sample. Clearly inflation rates between the U.S. and the other countries have converged. A similar result would hold for the formulation in differences.

5. Conclusion

We set out to use UIP to identify potential exchange rate regime changes for seven East Asian countries for the pre and post Asian Financial Crisis period. This seemed to be a

[13] The calculations are based on the idea that the inflation rate falls by this amount in a single period, or every period over a longer sample. You can make reasonable adjustments to find smaller slope coefficients.

worthwhile exercise for two reasons. First, those countries which followed a more flexible exchange rate regime are supposed to have gained more monetary independence unless there is fear of floating. Changes in parameters from the UIP equation should be able to confirm this one way or another. Second, we can think of the episode as a quasi experiment: some of the countries changed exchange rate regimes, while others did not. In addition, we have a control group of countries outside of Asia which is believed by most to have continued with its exchange rate regime regardless of the crisis.

What we found is that UIP is not well suited to identify exchange rate regime changes during the period we analyzed and for the countries we looked at.[14] In our view, authors have been asking UIP to reveal more than can be reasonably expected. There are various reasons for this, such as variation in results depending on assumptions about the stationarity of interest rate variables, changes in capital mobility, time frequency of data used, assumptions about expectations formation, etc. Most importantly, we believe that economic phenomena other than the choice of exchange rate regimes may explain why some authors have found evidence against the fear of floating, which implies that countries with more flexible exchange rates can pursue more of an independent monetary policy. In particular, it is inflation convergence during the later parts of the '90s which seems to explain much of the fall in the slope coefficient of the UIP relationship. Our results strongly suggest that statements regarding monetary independence and choice of exchange rate regime should take this into account more carefully.

[14] MacCallum (1994) also believes that deviations from the UIP may be due to monetary policy decisions of central banks and proposes to include a monetary policy reaction in an expression for the UID. Bird and Rajan (2001) and Rajan, Siregar and Sugema (2002) offer bank-based explanations for persistent interest rate differentials in East Asia. Also see Edwards and Khan (1985) and Willett *et al*. (2002).

Appendix 1: Further Dynamics involving the Korean UIP Specification

The analysis in this appendix continues to analyze various dynamic aspects of the UIP relationship for Korea. The results in section 3 suggest that following the analysis using levels and difference specifications, there was some doubt on whether Korea had pursued a more flexible exchange rate regime, as is commonly assumed in the literature. For example, the regression R^2 in the levels equation rises dramatically from 0.031 to 0.667, i.e. twenty times. Furthermore, the slope coefficient for the difference specification is statistically significant for the post-crisis period with a regression R^2 that is five times as high as Shambaugh's (2004) value for nonpegs. The crucial question now is whether we view regression (5) as a long-run relationship or as a regression between two stationary variables in the short run.

If we viewed the level regression (5) as representing a long term monetary equilibrium and with completed policy reaction rather than short-term financial market integration, then static regressions of this type have made a comeback when testing for cointegration (Engle and Granger, 1987). To determine whether the two interest rates are cointegrated, we estimate the static regression as shown in column (1) of Table 1 and then perform an ADF test on the residuals. For the pre crisis period the EG-ADF test statistic is -3.02. Since the critical value at the 10% level is -3.12, we cannot reject the null hypothesis of no cointegration even at this level. The result holds when other domestic variables such as inflation, money and income growth are added to the cointegrating equation. However, for the post crisis period, the EG-ADF statistic is -3.73, resulting in rejection of the null hypothesis of no cointegration at the 5% (but not at the 1% level).[15] Having found supportive evidence for the presence of cointegration in the post-crisis period, we use the dynamic OLS (DOLS) estimator (Stock and Watson, 1993) to find the cointegrating vector. This involves adding future, present, and lagged changes in the U.S. interest rate to the static regression, i.e. $R_t^{KOR} = \beta_0 + \theta R_t^f + \sum_{j=-p}^{p} \delta_j \Delta R_{t-j}^{US} + u_t$. Our estimate of θ is 0.233 with a standard error of 0.024, indicating a long run relationship between the two interest rates, but also one that is some distance from unity.

These results are opposite to what might be expected for a peg and a float: our prior is that interest rates should be more likely to be cointegrated for the peg than for the float. Shambaugh (2004, p.342) finds 29 cases where the null hypothesis of no cointegration is rejected. Of these, 23 are pegs and only 3 are nonpegs. Using the DOLS estimator, Shambaugh reports an average value of $\hat{\theta}$ of 0.84, with the majority lying between 0.8 and 1.2, i.e. close to unity. Note that for Korea, we can reject the null hypothesis of $\theta = 1$ even at the 1% level.

We find it instructive to investigate how robust this result is if we assumed interest rates to be stationary, i.e. I(0) variables. In that case, dynamics could be introduced by

[15] The Johansen procedure produces the same result, viz. no evidence of a cointegrating vector for the first period, but rejection of no cointegration for the second period.

adding a lag dependent variable, e.g. through a partial adjustment assumption. Equation (5) would be replaced by

$$R_t^* = R_t^f + E_t(s_{t+n} - s_t) + rp_t \tag{5a}$$

where

$$R_t = R_{t-1} + \lambda(R_t^* - R_{t-1})$$

Equation (5a) could be viewed as an equilibrium relationship and the partial adjustment could be the result of the presence of capital controls in the post-crisis period.

Column (2) in Table 2 shows the partial adjustment results for the pre and post-crisis estimation period. (Column (1) is added for comparison.) First note the increase in the regression R^2. As in the static regression results, and according to prior expectations, the slope coefficient for the peg is more than twice as large as for the float. Furthermore, solving both equations for the stationary state equilibrium results in a slope coefficient of 0.96 for the peg and 0.23 for the float. The former is obviously close to unity. Somewhat surprisingly, the speed of adjustment is higher for the float than for the peg. The mean lag in the former is roughly three months, while for the peg it is five and a half months. Also opposite to prior expectations is the relative magnitude of the regression R^2, with the float allowing for less room to maneuver than the peg. Durbin's *h* suggests the presence of further dynamic problems in the float specification, but not for the peg period. Finally, and perhaps most importantly, note the robustness in the results when comparing the estimate obtained from the DOLS estimation above and the post-crisis long-run coefficient for equation (13) (partial adjustment): both are almost identical. This is comforting given the difference in methodology.

Since we found evidence of cointegration for the float period, there should also be an ECM representation for the post crisis period. We could follow Shambaugh (2004) here and simply estimate the relationship

$$\Delta R_t = \theta(\alpha_0 + R_{t-1} - \alpha_1 R_{t-1}^f) + \beta_1 \Delta R_{t-1}^f + \sum_{j=1}^{l} \gamma_j \Delta R_{t-j} + \sum_{k=1}^{m} \delta_k \Delta R_{t-k}^f + u_t$$

Instead we follow Frankel *et.al*. (2002), who use the LSE/Hendry specification search of general to specific since the result will be more parsimonious.[16]

Having settled for a set of explanatory variables, a general Autoregressive Distributed lag model (ADL) is estimated first. Given that we work with monthly interest rate data, we felt that an ADL(5,4) was sufficient. Table 3 shows the results for the General Unrestricted Model (GUM).[17] As can be expected from the number of regressors, the

[16] see Hendry (1995; 269-70) or Gujarati (1995), 485-6 for a convenient summary. For detailed examples of earlier applied studies in consumption, see Davidson, Hendry, Srba and Yeo (1978); in the demand for money, Hendry and Mizon (1978).

[17]"[The GUM] is the most general, estimable, statistical model that can reasonably be postulated initially, given the present sample of data, previous empirical and theoretical research, and any institutional and measurement information available. ... The GUM is also formulated to contain the parsimonious,

standard errors of the parameters are quite large, giving low *t*-statistics. Some would argue that this is due to high multicollinearity among the explanatory variables, although on this see Hendry (1995 pp. 274-8 and p. 365). We are less concerned about uncertainty surrounding the GUM coefficients at this point. Instead we note that all of the previously published interest rate equations are nested in the GUM. Furthermore, we are interested in the pattern of coefficients to derive a more parsimonious model from the GUM, which will be functions of these parameters. This results in the following ECM representation:

$$\widehat{\Delta R_t^{KOR}} = \underset{(0.00009)}{0.00076} + \underset{(0.067)}{0.466^{***}}\ \Delta R_{t-1}^{KOR} + \underset{(0.082)}{0.126^{*}}\ \Delta R_t^{US} - \underset{(0.049)}{0.104^{**}}\ \Delta R_{t-1}^{US}$$

$$-\ \underset{(0.026)}{0.241^{***}}\ (R_{t-1}^{KOR} - R_{t-1}^{US}) - \underset{(0.020)}{0.184^{***}}\ R_{t-1}^{US}$$

t = 1999:1-2003:6, $\overline{R}^2 = 0.780$, SER = 0.00007

Since the parsimonious equation is nested in the GUM, we can use an *F*-test for the validity of the restrictions. The $F(5,43)$ statistic is 0.271. We therefore cannot reject the null hypothesis that the restrictions are valid. The solved form in Table 4 shows that this is not surprising since the major features of the GUM are accounted for.

The long run (stationary state) solution of the ECM equation is

$$R^{KOR} = 0.003 + 0.236 R^{US}$$

Again, it is comforting to see how close this result is to the long run solution calculated from the P.A. equation and the DOLS estimation of the cointegrating vector.

Finally, we experimented with alternative expectations hypothesis regarding the expected change in the exchange rate. The results are reported in columns (3) and (4) of Table 2. Clearly, in comparing these two specifications to column (1), the estimated relationship becomes substantially weaker. For the pre crisis period, the slope coefficient is insignificant in both columns (3) and (4). For the post crisis period, it is reduced in size to between one and eight hundredth of its size in column (1). The same is true for the post crisis period, although the slope coefficient remains statistically significant. Looking at a graph of the relationship, it becomes clear that large changes in the proxied expected exchange rate change are associated with relatively small movements in the Korean interest rate. Furthermore these movements in the exchange rate seem to dominate any change in the U.S. interest rate. The results in columns (3) and (4) are quite discouraging when faced with having to find a proxy for exchange rate expectations.

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interpretable, and invariant econometric model at which it is hoped the modeling exercise will end." Hendry (1995 p. 361)

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Figure 1a

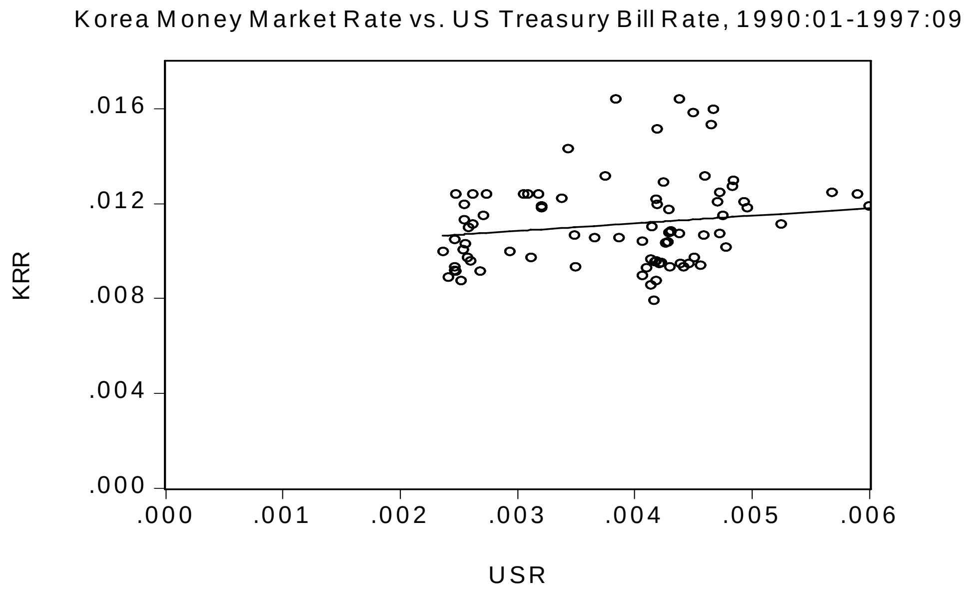

Figure 1b

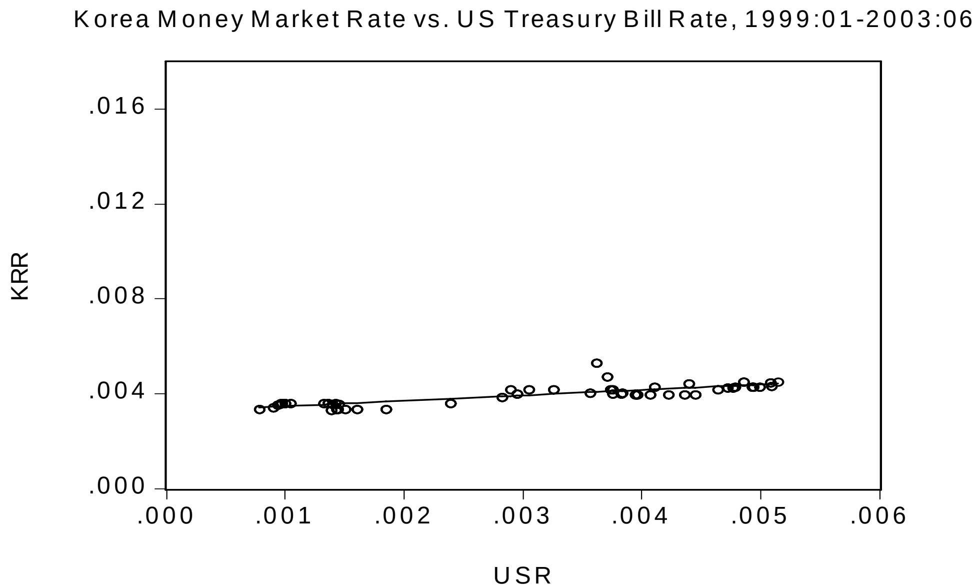

Figure 2: Regression of Monthly Slope Coefficients on Annual Slope Coefficientss

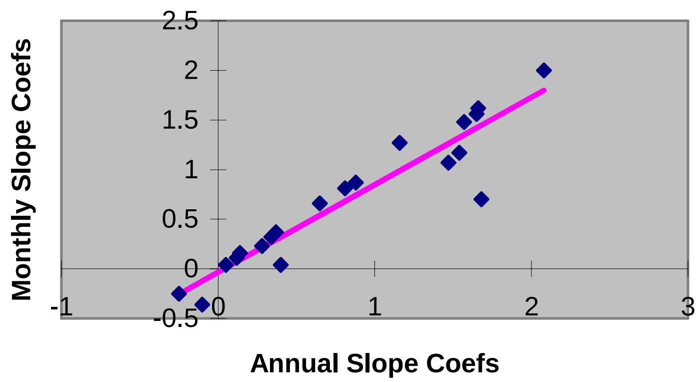

Table 1
Effect of U.S. Interest Rates on Korean Interest Rates, Level and Differences

Dependent Variable columns (1) and (2): Korea Money Market Rate, 1990:01-2003:06.

Explanatory Variable		(1) Levels	(2) Differences	(3) Shambaugh	(4) Shambaugh '90
U.S. Int. Rate	*pre*	0.320** (0.180)	1.736*** (0.598)	0.46** (0.04)	0.56*** (0.06)
	post	0.230** (0.025)	0.179** (0.100)	0.27*** (0.08)	0.35 (0.25)
Constant	*pre*	0.010*** (0.001)	0.00006*** (0.00001)	?	?
	post	0.003*** (0.0001)	-0.00004 (0.00003)	?	?
Summary Statistics					
Adj. R^2	*pre*	0.031	0.054	0.19	0.13
	post	0.667	0.030	0.009	0.006

Note: Columns (1) and (2): Pre-crisis and post crisis periods for columns are 1990:01-1997:09 and 1999:01-2003:06. Numbers in parenthesis are Newey-West HAC standard errors. *** indicates significance at 1% level, ** at 5% level, and * at 10% level of a one-sided *t*-test. Columns (3) and (4) are from Shambaugh (2004) and for comparison only. The sample is a panel of 103 countries for the sample period 1973-2000, where the dependent variable is the local interest rate. The explanatory variable is not always the U.S. interest rate. The result for the "pre-crisis" period is Shambaugh's peg slope coefficient, while the "post-crisis" period is his result for nonpegs. Column (4) presents Shambaugh's results for the 1990s.

Table 2
Effect of U.S. Interest Rates on Korean Interest Rates (Level)

Dependent Variable: Korea Money Market Rate, 1990:01-2003:06

Explanatory Variable		(1) Static Expect.	(2) Static Expect.	(3) Perfect foresight	(4) Extrapolative Expectations
R_t^{USA}	*pre*	0.320** (0.180)	0.170*** (0.065)	0.003 (0.019)	0.029 (0.027)
	post	0.230** (0.025)	0.078*** (0.012)	0.006*** (0.002)	0.004** (0.002)
Constant	*pre*	0.010*** (0.001)	0.001** (0.0008)	0.011*** (0.0003)	0.011*** (0.0003)
	post	0.003*** (0.0001)	0.001*** (0.0002)	0.004*** (0.0001)	0.004*** (0.0001)
R_{t-1}	*pre*	-	0.822*** (0.070)	-	-
	post	-	0.659*** (0.042)	-	-
Summary Statistics					
Adj. R^2	*pre*	0.031	0.709	-0.011	0.006
	post	0.667	0.954	0.104	0.035

Note: Pre-crisis and post crisis periods are 1990:01-1997:09 and 1999:01-2003:06. Numbers in parenthesis are Newey-West HAC standard errors. *** indicates significance at 1% level, ** at 5% level, and * at 10% level of a one-sided t-test.

Table 3: General Unrestricted Model (GUM) of the Interest Rate Equation

lags j	ΔR_{t-j-1}	ΔR_{t-j}^{US}
0	1.184 (0.10)	0.133 (0.101)
1	-0.378 (0.152)	-0.204 (0.129)
2	-0.128 (0.144)	0.168 (0.101)
3	0.105 (0.093)	-0.043 (0.103)
4	-0.034 (0.035)	0.003 (0.072)

Note: Sample Period 1999:1-2003:6; HAC standard errors in parenthesis, constant not reported here, $\overline{R}^2$ = 0.98, SER = 0.00007.

Table 4: Solved Form of the Interest Rate Equation

lags j	ΔR_{t-j-1}^{KOR}	ΔR_{t-j}^{US}
0	1.225	0.126
1	-0.466	-0.173
2	-	0.104
3	-	-
4	-	-

Table 5: Exchange Rate Classification for Seven East Asian Countries

Country	Period	*De Jure* Regime	*De Facto* Regime	Expectation on Slope
HKG	pre-crisis post-crisis	Fixed Fixed	Fixed Fixed	Constant
IDN	pre-crisis post-crisis	Intermediate Floating	Intermediate Floating	Decrease
KOR	pre-crisis post-crisis	Intermediate Floating	Intermediate Floating	Decrease
MYS	pre-crisis post-crisis	Intermediate Fixed	Intermediate Floating	Increase
PHL	pre-crisis post-crisis	Floating Floating	Fixed Intermediate	Decrease
SGP	pre-crisis post-crisis	Intermediate Intermediate	Intermediate Intermediate	Constant
THA	pre-crisis post-crisis	Intermediate Floating	Fixed Intermediate	Decrease

Table 6: UIP for Seven East Asian Countries plus Three Additional Countries, Annual Data

Country	Levels				Difference			
	Slope		R^2		Slope		R^2	
	Pre-crisis	Post-crisis	Pre-crisis	Post-crisis	Pre-crisis	Post-crisis	Pre-crisis	Post-crisis
HKG	1.54** (0.42)	1.47*** (0.087)	0.77	0.97	1.51** (0.50)	1.35*** (0.23)	0.55	0.81
IDN	1.99*** (0.216)	0.67 (0.70)	0.70	0.27	2.68*** (0.27)	1.50 (1.23)	0.87	0.21
KOR	0.34* (0.14)	0.28*** (0.033)	0.08	0.90	-0.0087 (0.53)	-0.30 (0.43)	0.00004	0.011
MYS	-0.25* (0.13)	0.047 (0.043)	0.096	0.12	-0.55 (0.31)	-0.40 (0.22)	0.22	0.068
PHL	-0.65 (0.56)	0.88*** (0.05)	0.15	0.96	-0.47 (0.32)	0.55* (0.23)	0.13	0.19
SGP	0.81** (0.25)	0.37*** (0.024)	0.65	0.96	0.59* (0.27)	0.26 (0.16)	0.46	0.076
THA	1.57*** (0.12)	0.12** (0.037)	0.91	0.56	1.57*** (0.28)	-0.47 (0.51)	0.73	0.019
AUS	2.08*** (0.36)	0.14 (0.080)	0.81	0.34	1.63*** (0.20)	0.49*** (0.035)	0.88	0.88
NZL	1.66*** (0.14)	0.04 (0.087)	0.90	0.01	-0.0006 (0.11)	0.4** (0.18)	0.00	0.15
CAN	1.65** (0.4)	0.65*** (0.040)	0.66	0.94	1.33*** (0.26)	0.59*** (0.11)	0.82	0.64

Note: Sample Period is 1990-2004. Pre-crisis period is 1990-1997 for HKG, KOR, SGP, THA, AUS, NZL, CAN; 1990-1996 for MYS; 1991-1996 for IND and PHL. Post-crisis period is 1999-2004 for all countries except for Indonesia. ***, ** and * stand for being significant at 1%, 5% and 10%, respectively.

Table 7: UIP for Seven East Asian Countries plus Three Additional Countries, Monthly Data

Country	Levels				Difference			
	Slope		R^2		Slope		R^2	
	Pre-crisis	Post-crisis	Pre-crisis	Post-crisis	Pre-crisis	Post-crisis	Pre-crisis	Post-crisis
HKG	1.17*** (0.040)	1.07*** (0.05)	0.87	0.94	1.02*** (0.30)	0.86*** (0.17)	0.08	0.19
IDN	1.65*** (0.41)	0.70* (0.35)	0.46	0.099	0.01 (0.4)	-1.56 (1.16)	0.007	0.016
KOR	0.32* (0.18)	0.286*** (0.028)	0.041	0.65	1.74*** (0.60)	0.15* (0.077)	0.064	0.038
MYS	-0.25* (0.14)	0.037 (0.047)	0.087	0.017	0.08 (0.16)	0.07 (0.092)	0.001	0.005
PHL	-1.04 (0.59)	0.87*** (0.17)	0.07	0.63	1,67 (1.77)	1.00** (0.45)	0.005	0.12
SGP	0.81*** (0.12)	0.37*** (0.018)	0.58	0.86	0.41* (0.23)	0.61*** (0.19)	0.02	0.22
THA	1.48*** (0.30)	0.11*** (0.036)	0.42	0.19	-0.30 (1.17)	0.059 (0.13)	0.00059	0.002
AUS	2.0*** (0.24)	0.16*** (0.057)	0.73	0.32	0.79*** (0.19)	0.16*** (0.057)	0.2	0.32
NZL	1.62*** (0.15)	0.04 (0.087)	0.80	0.01	0.70** (0.34)	0.39*** (0.089)	0.033	0.20
CAN	1.56*** (0.289)	0.66*** (0.047)	0.51	0.87	0.79*** (0.29)	0.58*** (0.15)	0.072	0.35

Note: Pre-crisis period: HKG, KOR, SGP, AUS, NZL, CAN: 1990:1-1997:9; IDN: 1991:6-1997:7; MYS: 1990:1-1997:4, PHL: 1991:2-1997:6; THA: 1990:1-1997:6. Post-crisis period: 1999:1-2005:4 except for IDN, which ranges from 1999:7 – 2005:4. New-West heteroskedasticity consistent covariance standard errors are reported in parenthesis.

Table 8: Inflation Rate Averages for the Pre-crisis and Post-crisis period

Country	Pre-crisis	Post-crisis	Difference
HKG	9.0	-4.0	-13.0
IDN	8.8	8.0	-0.8
KOR	6.6	0.8	-5.8
MYS	3.9	2.7	-1.2
PHL	9.2	5.9	-3.3
SGP	2.4	0.0	-2.4
THA	5.0	0.3	-4.7
AUS	2.5	1.5	-1.0
NZL	2.1	-0.1	-2.2
CAN	2.1	1.7	-0.4
USA	3.1	2.2	-0.9