INTERNATIONAL AND INTRA REGIONAL INTEREST RATE INTERDEPENDENCE IN ASIA: METHODOLOGICAL ISSUES AND EMPIRICAL RESULTS

by

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Abstract

We highlight a number of disconnects in recent policy discussions and empirical studies on financial integration and capital mobility. We focus on methodological issues in measures of interest rate interdependence. To assist in the critical analysis, we estimate international interest rate interdependence in Korea using money market rates. In our estimation, we pay specific attention to the instability/non-robustness of the coefficient estimates arising from the choice of control variables and different types of dynamic specification, the role of intra Asian interest rate interdependence, and implications for policy discussions.

Key words: *Asia, interest rates, exchange rate regimes, exchange rate expectations, Korea*

JEL Classification: *F31, F32, F33*

1. Introduction

There have been a number of major disconnects in recent discussions of Asian monetary issues. We wish to emphasize four. A number of observers have offered grand visions of monetary union and a common currency among a large group of Asian countries, in some cases within a decade. Such views often seem to reflect reactions to the creation of the Euro and the dislocations generated by the crisis of 1997-98, rather than careful analyses of the conditions enumerated in the literature on optimal currency areas (OCAs) for a common currency to be economically desirable. Recent interpretations of OCA analysis suggest that these criteria can be used more generally to investigate the weights that should be given to developments in the foreign exchange market in setting domestic monetary policy (Willett 2003a, b). The weight within a currency union would be one hundred percent. By extension, the weights that should be given to different bilateral exchange markets in a multi country world would depend in part on the degree of economic and financial integration with the corresponding countries. Thus, a broad intermediate range between fully independent policy making and a currency union should include various degrees of monetary and fiscal policies coordination. However, despite the arguments being made for monetary union within Asia, there seems to be relatively little discussion of the need for macroeconomic policy coordination. Unless one adopts the most extreme forms of new classical macroeconomic or endogenous optimal currency area models, substantial monetary and at least some fiscal policy coordination needs to precede the formation of a currency area if it is to operate smoothly. The failure to focus on such issues pertaining to macro policy coordination now is consistent with a skeptical view that advocates of currency union are acting as visionaries stressing future benefits and overlooking or minimizing costs.

In general, the case for monetary coordination is stronger, the greater is the degree of capital mobility (Rajan, 2004a). Here, however, we find another disconnect between policy advocacy and the empirical evidence: while advocates of monetary union in Asia over the medium term tend to assume high levels of capital mobility and financial integration, there is vast disagreement over the degree of capital mobility and financial integration. A number of arbitrage based empirical studies have produced estimates that are consistent with perfect capital mobility for some Asian countries (Table 1), but it is far from clear how much weight should be given to them. Willett *et al*. (2002) have argued that one popular class of such research was subject to strong upward biases, and many other estimates suggest substantially lower capital mobility. Still other papers that are not subject to these particular biases (such as Frankel *et al*., 2000 and Cheung *et al*., 2002) have also found results consistent with perfect capital mobility. Such findings are, however, inconsistent with another class of studies that has found considerable amounts of sterilization in Asia (see for example, Fry, 1988; Sarjito, 1996; Rooskaareni, 1998; Hutchison, 2003; and Willett *et al*., 2002). Thus we have another major disconnect between different types of studies.

A final disconnect that we wish to highlight is that most empirical studies on Asian monetary integration have emphasized the degree of capital mobility between various Asian countries and the global money market, usually proxied by the United States, and sometimes Europe and/or Japan. Even if highly precise, such studies tell us little about the direct monetary interdependence among Asian countries. For example, Indonesia and Korea could both be heavily influenced by developments in U.S. interest rates, while monetary developments in Indonesia might have only a modest impact on Korea and vice versa. As long as Asian countries retain some degree of national monetary independence, it is these intra Asian interdependencies that are most relevant for monetary policy coordination within Asia. Such intra Asian interest rate interdependence is an important and understudied area.

In view of the important policy lessons often drawn from studies on financial integration and capital mobility, this paper critically reviews the empirical literature on the subject, focusing on methodological issues and concerns. Section 2 briefly summarizes the various forms and definitions of financial integration and capital mobility. While numerous methods have been employed to measure financial integration, the focus of this paper will be on price measures, and specifically on arbitrage measures of interest rate interdependence. We focus on short term interest rates since these are the most relevant for monetary policy. Section 3 briefly highlights some examples of empirical irregularities in existing literature on interest rate interdependence. Section 4 turns its attention to key methodological issues in measuring interest rate interdependence. To assist in the analysis, we estimate the extent of international interest rate interdependence of Korea. In our estimations we pay specific attention to the instability/non-robustness of the coefficient estimates arising from the choice of control variables; the use of *levels* versus *changes* in interest rates; the use of dynamics in general, including the use of lagged dependent variables; and measures of exchange rate expectations. Section 5 focuses on intra Asian interest rate interdependence. The final section offers a few concluding remarks.

2. Measures of Financial Integration[1]

2.1 Brief Review of Non-Arbitrage Based Measures of Financial Integration

As noted, the focus of this paper is on arbitrage based price measures (referred to as interest rate interdependence). There are, however, a number of other measures of financial integration. Non-arbitrage price based measures include stock market correlations (both direct correlations as well as the extent to which international CAPM holds), or news based measures (i.e. extent to which interest rates and other financial market variables are impacted by common shocks versus country specific ones).[2]

There are at least two other broad non-price categories of financial integration (Figure 1). One are regulatory or institutional factors such as capital controls, prudential regulations and extent of internationalization of the banking system (or barriers to foreign bank entry, thereof). An obvious limitation of these measures is the difficulty of obtaining good proxies to measure such barriers or regulatory impediments that prevent financial integration. Recent research has found that the frequently used zero-one measure of capital controls can be highly misleading. More graduated measures of capital controls have recently been developed, although there are also not problem-free (see Nitithanprapas *et al.*, 2003; Willett *et al.*, 2004).

Next there are quantity based measures, including savings-investment (*S*-*I*) correlations, current account dynamics, consumption correlations and dynamics of private

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[1] This section draws partly on Cavoli *et al*. (2003) and Rajan (2004a).

[2] See Flood and Rose (2003) for a recent extension of the equity market integration literature using a general intertemporal asset pricing model.

capital flows. The argument regarding *S-I* correlations is simply that in a closed economy, by definition, savings must equal investment (i.e. correlations between savings and investment should be very high). At the other extreme, in highly integrated capital markets with a single world interest rate, domestic investment should be largely independent of domestic savings since the former can be financed via foreign savings (Feldstein and Horioka, 1980). There are significant empirical and theoretical shortcomings with the Feldstein-Horioka criterion and it remains a controversial measure of financial integration (for instance, see Coakley *et al*., 1998 and references cited within).[3]

A related strand of the literature has focused on current account dynamics and, in particular, whether the current account is stationary. Simply put, the argument here is that if savings and investment are cointegrated, their difference, which is the current account, ought to be stationary (Ghosh, 1995). The problem with this line of reasoning is that a finding of stationarity could either imply that an economy is not financially integrated (thus suggesting the existence of a long run relationship between savings and investment), or that the open capital market is imposing a solvency constraint on the country.[4]

The most recent literature on quantity based measures of financial integration has focused on the extent to which consumption co-moves between countries (pioneered by Obstfeld, 1989). The intuition behind tests of consumption correlations ("international risk sharing") is that financial openness ought to afford individuals the opportunity to smoothen consumption over time as they can borrow and lend on international financial markets. Thus,

[3] The RIP is a necessary but not sufficient condition for *S-I* to hold.

[4] Insofar as one would expect domestic savings and investment to converge over time due to binding intertemporal budget constraints, a better test would be to examine short term dynamics, i.e. error (equilibrium) correction models (ECM).

consumption in any one country should co-move less with income over time, and, if preferences are similar, consumption should be correlated across countries.[5] Conceptually, while consumption based Euler equation tests of capital mobility are attractive when attempting to discern whether a region is ripe for monetary union (as the degree of business cycle synchronization may be less relevant if agents can share consumption risks across borders), they are based on a number of restrictive assumptions that limit their practical usefulness.[6]

Another quantity based measure looks at the actual magnitude and dynamics of private capital flows (Rajan and Siregar, 2002). While there is always some merit in examining actual movements of cross-border capital flows, it is probably of limited use as a measure of financial integration. For instance, a country that is highly integrated with international capital markets - in the sense of there being no significant difference in domestic and international rates of return - may experience little if any international portfolio capital flows (at least debt related flows).[7]

2.2 Measures of Interest Rate Interdependence

The basic intuition behind parity conditions is that in a perfectly integrated financial market, arbitrage should equalize the prices of identical assets traded across different markets, i.e.

[5] A weaker form of risk sharing states that the degree of cross-country consumption co-movements exceeds that of output co-movements.

[6] These limitations notwithstanding, Kim *et al*. (2004) find that consumption risk sharing in East Asia is quite low (about 20 percent).

[7] Related to this, some studies have examined the quantitative effects of changes in monetary conditions on capital flows. Direct estimation of interest rate effects on capital movements has gone out of fashion, but modern approaches to estimates of domestic monetary sterilization of capital and reserve flows produce, as a by-product, estimates of offset coefficients, the extent to which an exogenous change in the domestic monetary base is offset by capital flows. These offer a too often neglected estimate of capital mobility.

the law of one price holds. There is, in general, no reason for nominal interest rates between two countries to be equal (money market rates differ, for example, between the U.S. and Korea). Similarly interest rates in one country do not have to change by the same amount following the change in another country, even the base country (Korea's central bank lowered interest rates recently while the Federal Reserve increased interest rates). A convenient starting point to understand the conditions under which a wedge may be driven between the interest rates of two countries, is the interest rate parity condition.

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

where R_t is the domestic interest rate, R_t^f is the foreign (base) interest rate (U.S. rate unless otherwise stated) and s_t is the current spot rate (both exchange rates in logs),[8] rp_t is the risk premium. Both the expected exchange rate and the risk premium are unobservable, in general, and we therefore have to find proxies for them.

In this section, we remind ourselves of the three common interest parity conditions resulting from (1), viz. CIP, UIP and RIP.[9]

a) The Covered Interest Parity (CIP) Condition

The CIP sets $E_t s_{t+n} = f_{t+n}$, where f_{t+n} is the (log of the) forward exchange rate for n periods into the future. In the absence of a risk premium, it is stated as follows:

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

[8] Throughout this section the exchange rate is quoted as the domestic price of foreign currency.

[9] Another arbitrage condition is the closed interest parity condition which essentially states that the returns on identical instruments of the same currency but traded in different markets (such as onshore and offshore markets) should be equalized. The measurement of the closed interest differential is difficult for developing economies as it requires that a particular asset is traded sufficiently for there to be a liquid offshore market for it (see Obstfeld, 1998; Frankel and Okwongu, 1996).

The CIP indicates that the difference between the current spot rate and the forward rate will equal the interest differential between similar assets measured in local currencies.

Two methods are often used to measure CIP. One is the mean covered interest differentials (CIDs), with the null hypothesis being that of a zero differential. The second method tests the CIP hypothesis by running the following regression:

$$R_t = \alpha + \beta(R_t^f + f_{t+n} - s_t) + \varepsilon_t \tag{3}$$

Under the null, $\alpha = 0$, $\beta = 1$.

While there have been a number of studies on the CIP involving industrial economies, there have been fewer ones pertaining to developing economies (exceptions include de Brouwer, 1999, and Bansal and Dahlquist, 2000). This is primarily attributable to the fact that many developing economies do not have sufficiently liquid forward foreign exchange markets, or if they do exist, the data on forward rates are not easily available. In any case, as Willett, *et al*. (2002) observe:

> "(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." (pp.424-5)

b) The Uncovered Interest Parity (UIP) Condition

The UIP is often represented as follows:

$$R_t = R_t^f + E_t(s_{t+n} - s_t) \tag{4}$$

The nexus between the UIP and the CIP is apparent by decomposing eq. (3) as follows (Frankel, 1991):

$$R_t - R_t^f - E_t(s_{t+n} - s_t) = [R_t - R_t^f - (f_{t+n} - s_t)] + (f_{t+n} - E_t s_{t+n}) \tag{5}$$

where the first bracketed term on the right hand side is the CIP (sometimes referred to as country or political risk premium) and the second term is the currency risk premium. If the CIP holds but the UIP is rejected, this would imply that forward rates are biased predictors of future exchange rate.[10]

Before formally testing eq. (3), researchers need to find a measure for the expected future exchange rate. The following four specifications for the expected change in the exchange rate have typically been used for estimation:

(i) perfect foresight	$E_t(s_{t+n} - s_t) = s_{t+n} - s_t$
(ii) extrapolative expectations	$E_t(s_{t+n} - s_t) = s_t - s_{t-n}$
(iii) static expectations	$E_t(s_{t+n} - s_t) = 0$
(iv) survey data	$E_t(s_{t+n} - s_t) = s_{t+n}^{survey} - s_t$

where s^{survey} is the (log of the) expected spot rate obtained from a survey of market participants, e.g. the Economist Intelligence Unit (EIU) Currency Consensus Forecast.[11]

Each one of these specifications has its advantages and drawbacks. One way to make the leap from theory to empirical operationalization is by using *ex-post* differentials. This may be justified by assuming Rational Expectations. This assumption - that the actual or *ex-post* spot exchange rate equals the expected spot exchange plus an uncorrelated error term - is a practical way of overcoming the problem of non-observable expected exchange rate changes. Alternatively, the static expectations hypothesis is based on the idea that exchange

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

[11] See Frankel and Froot (1987, 1989), Taylor (1989), MacDonald and Torance (1990), Cavaglia et al. (1993), among others. For an application to four Asian exchange rate markets, see MAS (1999).

rates follow a random walk. Or, perhaps, agents have a backward looking behavior. Although this is not very likely from a theoretical perspective, the difference in the results between using the perfect foresight model and extrapolative expectations is quite small when using monthly data and periods during which we do not observe vast exchange rate fluctuations and trends.

Another approach is to use surveys of exchange rate expectations of market agents (e.g. Chinn and Frankel, 1995). Generally speaking, survey data is notoriously difficult to obtain. Interestingly for the study of Asian economies, a paper by the Monetary Authority of Singapore (MAS, 1999) attempts to model the formation of exchange rate expectations when these can be observed from survey data for Malaysia, Thailand, Indonesia, and the Philippines. Various expectations behaviors were considered (extrapolative, static, and adaptive). The study concludes that "market participants focused mainly on recent changes in the spot rates as the basis for making projections into the future" (p.20). In fact, the change in the expectations actually moves in the opposite direction to recent changes in the exchange rate. This suggests the possibility that some sort of technical analysis is being used in the forecasting.

We return to the issue of measuring exchange rate expectations in Section 3. However, apart from measurement issues, interpretation of the UIP test is not without its problems. First, the test for UIP is in fact a joint test for the CIP and the currency risk premium. Second, tests for UIP generally assume that all agents form expectations rationally. Thus, the failure of the UIP to hold (in the sense that there exist large and persistent UIDs) could be because (a) the Covered Interest Parity (CIP) does not hold; (b) there may be large and time varying country or political risk premia (imperfect asset substitutability); or (c) there are problems with proxies/measures of exchange rate expectations (or that the market consists of heterogeneous agents).[12]

c) The Real Interest Parity (RIP) Condition

For completeness, and although we will not use real interest rates in our estimation below, we will mention RIP arbitrage. This condition may be derived by first taking the following UIP equation:

$$E_t(s_{t+n} - s_t) = R_t - R_t^f \tag{6}$$

and substituting it into an expression for relative purchasing power parity (PPP):

$$s_t = p_t - p_t^f \quad \text{or} \qquad E(s_{t+n} - s_t) = E_t(\Delta p_{t+n} - \Delta p_{t+n}^f) \tag{7}$$

(where p is the (log of the) price level). Combining the equations (6) and (7) with the Fisher equation, $r_t = R_t - E_t \Delta p_{t+n}$ (where r is the real interest rate) yields the expression for the RIP:

$$r_t = r_t^f \tag{8}$$

Clearly for the RIP to hold, UIP, PPP and the Fisher hypothesis need to be in place simultaneously. This is unlikely given the lack of empirical success of either the UIP and PPP or the short to medium terms. Thus, the RIP is generally considered a very long-run interest parity condition encompassing both real and financial linkages.[13]

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[12] 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).

[13] In fact, the UIP itself may also be more valid over longer time horizons -- over one year (see Madarassy and Chinn, 2002 and Meredith and Chinn, 1999). For an interesting application which links RIP deviations to UIP deviations and PPP deviations, see Cheung *et al.* (2003).

3. Examples of Empirical Irregularities in Existing Literature

Close examination of some of the studies reported in Table 1 reveals the existence of numerous contradictions depending on methodology, time periods, and variables used in each study. Consider some examples.

Edwards and Khan (1984) and Ahn (1994) use the same methodology and a similar data set to measure the degree of financial openness of Singapore. Edwards and Khan find that the U.S. interest rate is the only determinant (coefficient 0.922) of the Singapore interest rate. On the other hand, Ahn finds that the foreign interest rate can explain only 30 percent of the movements of local interest rate. Ahn further finds that a lag of the local interest rate itself plays an important part in determining Singapore interest rate (approximately 43 percent), as well as real money supply and real income.

Different types of market determined interest rates can generate different outcomes. For example, Chinn and Frankel (1994) report that the foreign interest rate is significant with a coefficient -0.989 if the rate on monetary stabilization bonds is used. However, there is no influence of foreign interest rates if the corporate bond rate is included instead.

Realizing that finding the right money market determined interest rate might be difficult for developing countries, Haque and Montiel (1991) and Dooley and Mathieson (1994) overcome this problem by determining the effect of foreign interest rate on the local economy from estimation of the money demand function. Dooley and Mathieson report a large coefficient of 0.95 in both the OLS and IV estimates (indeed they cannot reject perfect capital mobility for Korea using the OLS estimate). These results of high-to-perfect capital mobility reported by Dooley and Mathieson do not agree with Ahn's (1994) finding, where only domestic variables (real money supply, real income, and past value of its own interest rate) influence the Korean interest rate and the foreign interest rate plays no role. Furthermore, Willett *et al*. (2002) point out that the Hague-Montiel and Dooley-Mathieson approaches assume no sterilization of capital flows. Estimates of actual levels of sterilization result in a substantial lowering of their estimates by as much as one half in some cases.

4. Methodological Issues and New Empirical Results for Korea

We believe that there is no better way of highlighting the various methodological concerns and ambiguities in the empirical literature on interest rate interdependence than by presenting our own set of estimation results and highlighting the problems along the way. In this section we try to determine the degree of international interest rate interdependence for Korea using money market rates.

When it comes to estimating the relationship between the Korean money market rate R_t^{KOR} and a base ("foreign") rate R_t^f, there are several methodological issues that need to be addressed. Although other variables may enter the relationship between the two variables, our primary parameter of interest will be the coefficient on the base rate. The size of the coefficient has been used to classify degrees of floating/pegged exchange rate regimes and capital mobility, both across countries and across time within a country.

4.1 Exchange Rate Expectations

Equation (1) can be modified under various conditions (Shambaugh, 2004):

a) Credible peg exchange rate plus small, constant, or uncorrelated with shocks to base rate risk premium:

$$\Delta R_t^{KOR} = \Delta R_t^f \,. \tag{9}$$

This equation can also be expressed in levels for the case of a constant risk premium.

b) Non-credible peg plus variable risk premium:

$$\Delta R_t^{KOR} = \Delta R_t^f + \Delta E_t(s_{t+1} - s_t) + \Delta rp_t; \tag{10}$$

which, under fixed exchange rates, reduces to:

$$\Delta R_t^{KOR} = \Delta R_t^f + \Delta E_t s_{t+1} + \Delta rp_t. \tag{11}$$

c) Floating exchange rate:

$$E_t(s_{t+1} - s_t) = R_t^{KOR} - R_t^f - rp_t. \tag{12}$$

For this case, interest rate differentials can persist and result in expected exchange rate changes.

We use monthly data for Korea from January 1990 (1990:1) to June 2003 (2003:6). It makes sense to treat the pre and post crisis periods as separate regimes. As we shall see, these sub-periods yield substantially different estimates.[14] The period has been classified by Reinhart and Rogoff (2004) as containing an intermediate exchange rate regime until November 1997 (pre announced crawling band until November 1994, and *de facto* crawling peg to U.S. dollar until November 1997), and after that a floating exchange rate regime ("freely falling" from December 1997 - June 1998, and freely floating from July 1998 - December 2001).[15] Alternatively, and based on officially reported exchange arrangements, Frankel *et al*. (2002) and the IMF report a managed floating (intermediate) classification until November 1997. However, there is considerable movement in the exchange rate during October and November. As a result we use September 1997 as the end point of the pre crisis period.

[14] The other approach is to estimate time varying parameters. Recent studies on capital openness allowing time-varying risk premium are by Kim and Lee (2004) and Lixing (2000).
[15] This category is used for countries with a 12-month rate of inflation above 40%.

The Reinhart and Rogoff classification of the won as freely floating from July 1998 onwards is inappropriate because large changes in reserves clearly indicate that there has been a good deal of management of the exchange rate. Reinhart and Rogoff only consider the behavior of the exchange rate in forming their classification. However, McKinnon and Schnabl (2003), Hernandez and Montiel (2003), Willett and Kim (2004) and others show that despite some fear of floating, the won has been a good deal more flexible after the crises than before. To be on the safe side, we start the post crisis estimation period in 1999:1.[16]

In addition, changes in the stringency of capital controls should be expected to influence the degree of interest rate interdependence, just as could changes in the exchange rate regime. As with exchange rate regimes, we cannot expect monetary policy to remain the same across varying degrees of capital control. Ideally we would allow an index of capital controls to interact with the base rate. However, there is no index that we are aware of which is available at the monthly frequency.

Table 2 displays regression results for equation (1) in combination with (i), (ii), and (iii) (numbered (12) to (17)). Equation (12) in Table 2 shows the slope coefficient in the case of static expectations being greater for the peg period than for the float. For the two periods, the slope coefficients are statistically significant. Note that the slope coefficient for the peg is quite far from unity. The intercept could be interpreted as a risk premium. One surprising result is the low regression adjusted R^2 for the peg, particularly when compared to the float period. Figures 2a and 2b suggest an explanation: the scatter is much tighter for the post crisis period due to the smaller variation in R_t^{KOR} for the second period (note that the scale is identical for both figures).

[16] See Flood and Rose (2002) for an analysis of how well UIP works during a crisis period compared to a period or normalcy.

Equations (14) and (15) add the perfect foresight and extrapolative expectations measures to the base rate. Clearly, in comparing these two specifications to equation (12), the estimated relationship becomes substantially weaker. For the pre crisis period, the slope coefficient is insignificant in both (14) and (15). For the post crisis period, it is reduced in size to between one and eight hundredth of its size in (12). The same is true for the post crisis period, although the slope coefficient remains statistically significant. From Figure 3, 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, as shown in the regression results. The results in (14) and (15) are quite discouraging when faced with finding a proxy for exchange rate expectations.

4.2 Issues in Estimation: Levels versus Changes

The issue of estimating interest rate interdependence for a panel of countries using levels (equation (1)) or differences (equation (9)) directly pits Frankel *et al.* (2000) and (2002)[17] against Shambaugh (2004). We would be less concerned about the method employed if economic policy conclusions were mainly unaffected irrespective of choice. Unfortunately, as we shall see below, this is not the case for Korea, and we suspect it not to be the case for many other countries.

Briefly, the econometric considerations are as follows. It is well known that regressing two I(1) series on each other often results in spurious regressions (Granger and Newbold, 1974; Phillips, 1986). Furthermore, statistical inference is tricky in this situation, as *t*-statistics are not normally distributed, not even asymptotically. As a result, Shambaugh

[17] There are, of course, others who have estimated the relationship in levels, see, e.g., Hausmann *et al*. (1999).

(2004) tests hypotheses based on (9) rather than (1) in his panel estimation using annual data, since most of his interest rate series are integrated of order one (contain a unit root).[18] Frankel *et al.* are, of course, aware of this problem and address it. They primarily note that they view (1) as a long run relationship, even though they do not formally test for cointegration in their monthly data panel estimation. As an aside, Frankel *et al.* (2002; 12) note that “a priori we would expect interest rates to be I(0) variables,” and refer to the argument made by Cochrane (1991). We also find it somewhat hard to think of nominal interest rates being I(1) variables in general and especially over longer samples. If this were so then either real interest rates and/or inflation rates (inflationary expectations) would have to be I(1). It is not clear why real interest rates should be non-stationary in general, and inflation rates are often found to be I(0).[19] Of course if interest rates were I(0) variables, then Shambaugh’s argument for using (9) instead of (1) is no longer valid. But does it matter, for practical purposes, if we estimate the relationship in levels or difference?

4.2.1. Level Regressions

We already discussed equation (12) in Table 2 above: the size of the slope coefficients correspond to prior expectations given the postulated regimes at hand for Korea, but the relative size of the two regression R^2 does not. The crucial question now is whether we view equation (1) as a long-run relationship or whether we look at it as a regression between two stationary variables.

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[18] Interestingly enough, the *p*- value for R_t^{KOR} containing a unit root for the pre crisis period using the ADF test is only 5.14%, meaning that we can almost reject the null hypothesis of a unit root at the 5% level. For R_t^f the *p*-value is close to 20%.The respective *p*-values are much higher for the post crisis period.

[19] 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.

If we viewed the level equation (1) as representing a long term monetary equilibrium and full policy reactions rather than short-term financial market integration, then static regressions of this type have made a comeback when testing series for cointegration (Engle and Granger, 1987). To determine whether the two interest rates are cointegrated, we estimate the static regression as shown in equation (12) 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. This result does not change once other domestic variables such as inflation, money growth, and income growth are added to the cointegrating equation. However, for the post crisis period, the EG-ADF statistic is -3.73, allowing us to reject the null hypothesis of no cointegration at the 5% (but not at the 1% level).[20] 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. $R_t^{KOR} = \beta_0 + \theta R_t^f + \sum_{j=-p}^{p} \delta_j \Delta R_{t-j}^f + \varepsilon_t$. 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 quite 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,

[20] 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.

Shambaugh reports an average value of $\hat{\theta}$ of 0.84, with the majority of them between 0.8 and 1.2, i.e. close to unity. Note that for Korea, we can reject the null hypothesis of θ = 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 adding a lag dependent variable, e.g. through a partial adjustment assumption. Equation (1) would be replaced by

$$R_t^* = R_t^f + E_t(s_{t+n} - s_t) + rp_t \tag{1a}$$

where

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

Equation (1a) 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.

Equation (13) in Table 2 shows the partial adjustment results for the pre and post crisis estimation period. 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 close to three months, while for the peg it is roughly 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 the estimation methodology.

4.2.2. Difference Regressions

Where the likelihood of spurious correlations is high, estimation in levels is typically not recommended. If there is evidence of cointegration, then an ECM specification should be used. Instead Shambaugh (2004) suggests first differencing the data instead since (a) he does not find cointegration for all country pairs involved, and (b) "hopes the dynamics have largely settled" given that the panel data is annual. Although we have higher frequency data, we will also use the differencing methodology as a first attempt to test for short-run financial integration.

Results for first differences are reported in equation (18). We added a constant and allowed for the slope to be different from one in equation (9).[21] The results are as expected, with the slope coefficient substantially higher for the peg compared to the float.

$$\Delta R_t^{KOR} = \underset{(0.00009)}{0.00006} + \underset{(0.598)}{1.736^{***}}\ \Delta R_t^{US} \tag{18}$$

$R^2 = 0.054$; t = 1990:1 – 1997:9.

$$\Delta R_t^{KOR} = \underset{(0.00003)}{-0.00004} + \underset{(0.100)}{0.179^{**}}\ \Delta R_t^{US}$$

$R^2 = 0.030$; t = 1999:1 – 2003:6.

[21] Shambaugh (2004, pp.306-8) presents arguments for the slope to be different from unity for a pegged exchange rate regime, and for the slope to be different from zero for a floating exchange rate regime.

Using a Wald test, we find that the *p*-value for the null hypothesis of a zero intercept and unit slope is approximately 0.40 for the pre crisis period. Since equation (9) implies that the difference between the two interest rates $\Delta R_t^{KOR} - \Delta R_{t-1}^{f}$ should be zero, we also tested for the presence of a unit root using the ADF statistic. The null hypothesis of a unit root is rejected at the 1% level.[22] The (change of the) interest rate difference therefore appears to be stationary.

Kim and Lee (2004) report a slope coefficient of 2.946 for the pre crisis period using a slightly longer sample.[23] One possible theoretical explanation for a coefficient of this magnitude is that "increases in the base rate make investors doubt the peg's stability" (Shambaugh, 2004, p.306). However, it is also instructive to glance back at Figure 2a. It now becomes immediately clear why the slope coefficient is this large: for a given change in the U.S. Treasury Bill, there are corresponding large changes in the Korean interest rate. In addition, the regression adjusted R^2 is only about a half of the average size reported by Shambaugh (2004) for the pegs (0.13) in his panel of 103 countries. On the other hand, the post-crisis results indicate a slope half as large as those reported by Shambaugh and a higher regression R^2 (0.35 and 0.006 respectively). Still, the relative order of the two regression R^2s is as expected, with the lower value for the float perhaps indicating that the monetary authority has more autonomy during a float.

[22] Compared to Shambaugh's (2004, p.322) results for pegs and nonpegs, the slope coefficient in (18) is roughly three times as large as the average coefficient in his annual results (0.56).

[23] Kim and Lee (2004) estimate the break date endogenously and time it at August 1997. Their sample period is 1987:1 - 2002:4.

4.2.3. Further Dynamics and ECM

Shambaugh (2004) finds that the results based on testing equation (9) are not adequate when using monthly data. He allows for dynamics to enter by adding lags of ΔR_t^f. We follow this procedure here and allow for the maximum lag length to be determined by the BIC criteria. Unfortunately the BIC was at a minimum for no additional lag so that no further insights can be gained from this exercise. At face value, the absence of lags suggests that the domestic interest rate adjusts *immediately* to changes in the base rate. This result is quite different from the conclusion we drew from the level regressions. Of course, and as mentioned by Shambaugh (2004), there is the possibility that the correlations among the levels of interest rates across countries may be spurious with both responding in part to third factors. Since interest rates are endogenous variables, the pattern of correlations among these may vary substantially depending upon the pattern of shocks.

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 following relationship

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

where the maximum lag length of past changes in the domestic and base interest rate are determined by either the AIC or BIC. Instead we follow Frankel, Schmukler, and Serven (2002), who use the LSE/Hendry specification search of general to specific since the result will be more parsimonious.[24]

[24] 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).

Having settled for a set of explanatory variables, a general Autoregressive Distributed lag model (ADRL) is estimated first. Given that we work with monthly interest rate data, we felt that an ADRL(5,4) was sufficient. Table 3 shows the results for the General Unrestricted Model (GUM).[25] As can be expected from the number of regressors, the 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). However, 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:

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

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

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

Since the parsimonious equation will be 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

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

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

It is comforting to see how close this result is to the long run solution calculated from the partial adjustment equation and the DOLS estimation of the cointegrating vector.

5. Intra-Regional Financial Integration

In this section we extend our analysis to allow for influences of regional interest rates. For simplicity, we only study the robustness of the level and difference specifications. That is, we will estimate equations of the following form:

$$R_t^i = \alpha + \beta_1 R_t^{US} + \beta_2 R_t^{\mathrm{Reg}} + u_t \qquad (18)$$

where we consider Indonesia, Malaysia, the Philippines, and Thailand as countries that have a potentially significant influence on Korean interest rates (i = KOR here). We do not include Singapore and Hong Kong as regional centers because their high degree of financial integration with the U.S. could cause severe multicollinearity problems if both were included in the same regression. We also allow any of these regional interest rates to appear on the l.h.s. of the equation, with Korean and other regional interest rates entering as explanatory variables. This brings the total of estimated equations to twenty. Alternatively, we will use the differenced form of (18).

5.1. Estimation in Levels

While it is tempting, at first, to estimate equation (18) without the U.S. interest rate as a control, the omitted variable bias is bound to be substantial, since many of the regional interest rates are somewhat correlated with the U.S. interest rate. Therefore we do not report these results here.[26] We do want to emphasize, however, how misleading such regressions can be. For example, Malaysian and Thai interest rates would appear to be significantly affected by Indonesian interest rates and to some extent even by interest rates in the Philippines (for the post crisis-period).

Table 5 presents the estimated coefficients. The rows indicate which of the regional interest rates appears as the dependent variable in equation (18). The columns show which regional variable is included in addition to the U.S. interest rate. For example, without the influence of the regional interest rate variables, the coefficients of the U.S. interest rate were 0.320 for the pre-crisis period (equation (12)) and 0.230 for the post-crisis period (equation (17)) for Korea. For the most part, adding the regional interest rates as an explanatory variable does not significantly alter the estimated coefficients of the U.S. interest rates in the Korean row. However, there are several cases for which other regional interest rates exert a significant influence in addition to the U.S. rate. For example, Korean interest rates significantly affect interest rates in Malaysia and the Philippines.

Note that this statement is sensitive to the starting point of the sample period in some cases. For example, while it seems appropriate to start the post crisis period in January 1999 for most of the countries, Indonesia still appears to have been in the aftershocks of the crisis, most likely from political unrest within the country. The volatility of the Indonesian interest rate with respect to U.S. and Korean interest rates persists until June 1999. Hence we set the starting date of the post-crisis period for Indonesia to July 1999. Without this adjusting of the starting date, the estimated coefficients of the U.S. and regional interest rates become unreasonably large. Furthermore, they are statistically significant. For example, the coefficients are well over 11 in the case of Korean or Malaysian regional interest rates. The

[26] The results are available from the authors on request.

implication is that a small change in the regional rate will cause a substantial change in Indonesian interest rates.

As in the earlier case in the text, it is instructive to examine the scatter plot of interest rates between Indonesia and Korea during the post-crisis period. Even though much of the observations are clustered towards the origin, the high volatility of Indonesian interest rates during the first six months of 1999 causes some of the observations to stretch far to the right on the horizontal axis. Note that in addition the estimated coefficient of the U.S. interest rate becomes negative (and statistically significant).[27] Adjusting the sample period makes most of the abnormalities disappear. This is an excellent example of just how sensitive the results are to somewhat marginal changes, in this case in the sample period.

Looking at the other rows, the following results stand out: during the pre-crisis period, there is evidence of a feedback between Indonesia and Malaysia, Korea and Malaysia, Malaysia and Thailand, and Thailand and Indonesia. This leaves the Philippines, which is considered to have the least developed financial markets among the East Asian countries. According to the estimates, the Philippines plays little to no role either as an influence to or being influenced by other countries. These findings are consistent with the results reported by Anoruo *et al.* (2002), who find feedbacks among East Asian countries with the exception of the Philippines. On the other hand, the degree of regional correlation is relatively small during the pre crisis period, with the exception of Malaysia acting as the regional center.

For the post crisis period, the evidence of regional integration among the East Asian countries is stronger. There are several cases for which estimates of we find relatively large

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[27] When Indonesia is a regional center, the estimated coefficients of the US and regional interest rates are not sensitive to the starting date of the post-crisis period.

effects of regional interest rates on the domestic interest rates. More specifically, first Korea and Malaysia seem to have a strong and significant effect on the other countries with the exception of Indonesia. Second, neither Indonesia nor the Philippines seem to affect the other countries in a significant manner. Third, we experimented further with the specification by including the lag dependent variable and domestic macroeconomic variables (results not shown here). The inclusion diminishes the effect of the regional interest rates. Taken together, the results suggest a dominant role of both Korea and Malaysia within the region. This is most noticeable in the case of the Philippines. These results suggest the counter intuitive result that as the Philippines adopted the floating exchange rate regime after the crisis, its interest rates follow the interest rate in the base country more closely. One possible explanation that deserves further research is that the authorities have been sensitive to exchange rate fluctuations and have adjusted interest rates to reduce them, i.e. that they have been subject to substantial fear of floating.

5.2. **Estimation in Differences**

Table 6 reports the regression results for the difference model. As in Table 5, the rows indicate which of the regional interest rates appeared as the dependent variable in the differenced form of equation (18). Proceeding as we did in the previous section, we first note that in the absence of the regional interest rate variables, the coefficients of the U.S. interest rate were 1.736 for the pre-crisis period (equation (18)) and 0.179 for the post-crisis period for Korea. For the pre-crisis period, the U.S. coefficient remains very much the same irrespective of which regional interest rate variable is added. Not unexpectedly, none of the coefficients for the regional variables are even remotely significant. These results change for the post-crisis period. Several of the regional variables now seem to have a significant, albeit small, effect on the Korean interest rate, even when controlling for the U.S. interest rate. On the other hand, Korean interest rates appear to have, in general, a larger, and often statistically significant, effect on Malaysia, the Philippines, and Thailand for the post crisis period, and when compared to other regional rates.

Looking at the other rows of Table 6, Malaysia remains as an important regional center. It continues to influence the interest rates of Korea, the Philippines, and Thailand. However, when comparing the results to the levels estimation, Malaysia's coefficient is now only half as large as before. This apparent importance of Malaysia does not conform with our judgmental expectations about its relative importance. It is neither large nor a major financial center. Thus we believe these results will need further examination. Finally, interest rates in the Philippines continue to be significantly correlated with both the changes in the U.S., and Korean and Malaysian interest rates.

6. Concluding Remarks

The UIP approach can be adopted to provide a continuous metric for interest sensitive capital mobility in terms of the coefficient of the effects of foreign interest rates on the domestic interest rates. However, as we have shown, there can be a great degree of variability in estimates of patterns of interest rate interdependence due to differences in methodologies and time periods considered (with financial market developments and changes in capital controls, and exchange rate regimes).[28] In our judgment, there needs to be

[28] To further illustrate the importance of time periods, Frankel et al. (2000), Chinn and Frankel (1994), Dooley and Mathieson (1994) mostly agree that Thailand has high capital mobility, particularly in the 1990s. Even though Kim and Lee (2004) report the same finding for Thailand pre-crisis period, their results indicate that the foreign interest rate plays no role in interest rate determination of Thailand during post-crisis period.

much more focus on evaluating the appropriateness of the different methodologies that have been widely used. We see the recent paper by Shambaugh (2004) as a valuable step in this direction and we hope that our paper contributes to furthering this process.

There is no obvious single clearly best dynamic specification or proxy for exchange rate expectations, but if we had to choose at present, we are inclined to give the most weight to estimates in difference form when using *panel* data at *annual* frequency. This recommendation is made despite the unresolved issue of whether nominal interest rates are I(1) or I(0) variables. Except where there are substantial trends under flexible rates or there is substantial disequilibrium under pegged rates, we suspect that static expectations is the best simple proxy for exchange rate expectations. Much of the recent variability in Asian interest rates appears to have been unexpected, making actual *ex post* changes a poor proxy for *ex ante* expectations. In principle when looking at the influences of foreign on domestic interest rates, we should control for other influences on domestic rates. As we move from pegged to more flexible exchange rate regimes, a major channel of monetary interdependence becomes *changes* in exchange rates.

Our recommendation regarding the variables in differenced form when using annual panel data is based on the idea that dynamics of short term capital flows should work themselves out within a year. Of course there may be exceptions such as the presence of a fixed exchange rate regime coinciding with a reaction function tied to a balance of payments constraint. With moderate capital mobility this would yield low interest rate interdependence in the short run if countries sterilized, but the long run balance of payments constraint would force eventual adjustments. Interest rate equations of the type estimated in the text then would not reflect short-term financial adjustments. But these considerations probably more than anything reflect our doubts about the appropriateness of using panel methods for the issue at hand. Perhaps, and at a more general level, we should realize that the simple relationship between the domestic and base interest rate is asked to reveal a lot of information.

When using *monthly* data, we suggest using levels equations to test for cointegration. Cointegration between the domestic and base interest rates would yield supportive evidence for a *long-run* relationship. Evidence of cointegration, however, does not necessarily imply strong evidence for a currency peg. For example, such a long run relationship might be due to the limited range in which interest rates fluctuate in low inflation countries rather than any direct interdependence. Short-run dynamics then should be established using an ECM. For truly financial integration, including, of course, the case of no capital controls, adjustments should be completed in the short-run, and the error (equilibrium) correction term should carry a non-significant coefficient. This would, in essence, reduce the ECM form to an equation using differenced variables for both the domestic and base interest rate.

However, there are issues other than the data frequency, levels vs. differences, appropriate sample periods, adjustment costs, etc. which need serious consideration. For example, once we consider less than perfect financial integration, we have an identification problem in that a coefficient on the foreign interest rate variable can be a combination of the monetary reaction function and the degree of financial integration. This poses a potential problem in addition to other possible regime changes such as the wide spread shift to lower level inflation levels and thereby interest rate regimes. Again, to reveal much information is a lot to ask from such a simple relationship.

Our results present strong evidence that despite the rapid growth of international capital mobility, for most Asian countries it is far less than perfect. Our results are consistent with our priors that most Asian countries are more highly financially integrated with the

United States than with each other, and that the levels of direct interest rate interdependence within Asia vary substantially across pairs of countries. We have little confidence, however, in our current estimates of the precise strengths of these relationships. For the present, any such estimates need to convey a warning label “Dangerously Unstable, Handle With Care.”

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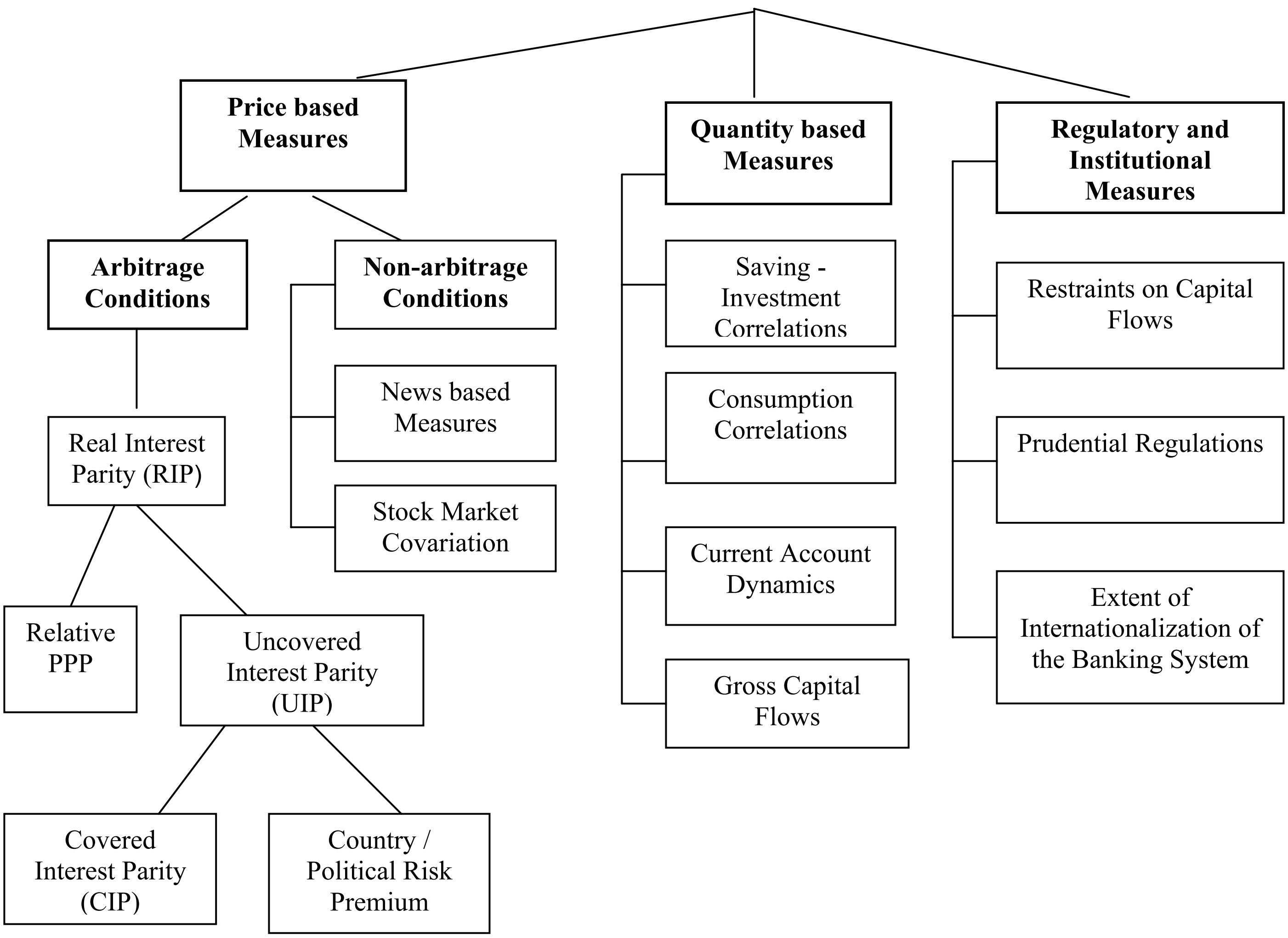
Figure 1

Categorizing Measures of Financial Integration: A Simple Framework



Source: Cavoli, Rajan and Siregar (2003)

Figure 2a

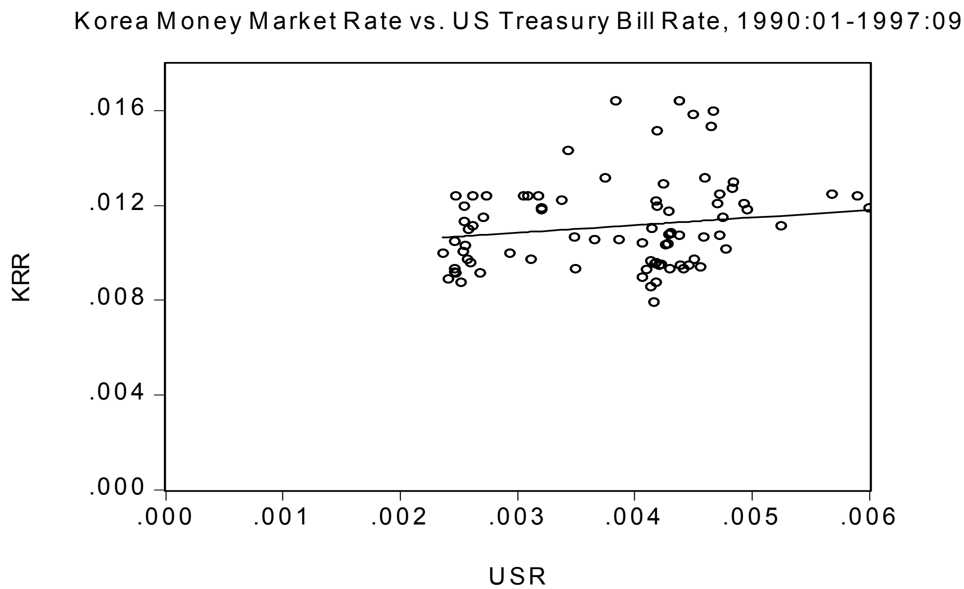


Figure 2b

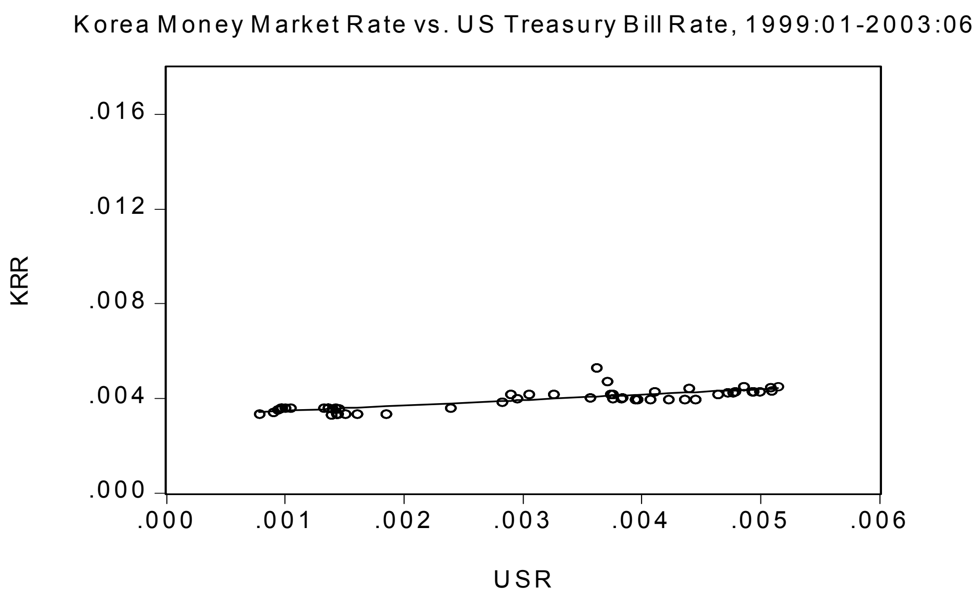


Figure 3
R_t^{KOR}, R_t^{US}, and the Percentage Change in the Exchange Rate

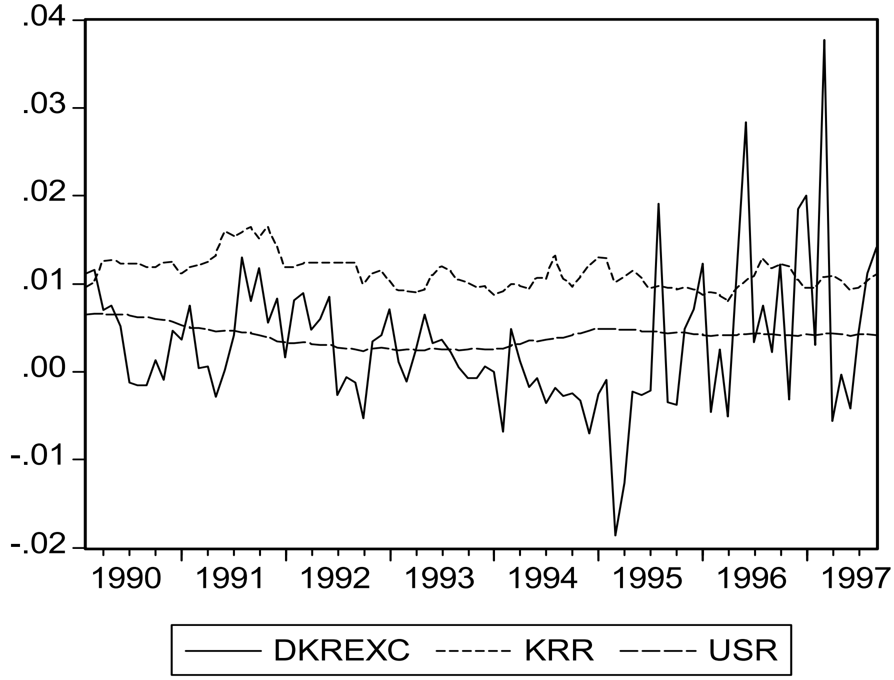


Table 1
Selected Empirical Findings on Capital Mobility in East Asia

	Data range	Method	Hong Kong	Indonesia	Korea	Malaysia	Philippines	Singapore	Thailand
Edwards and Khan (1984)	1976:03 - 1983:04	OLS	-	-	-	-	-	0.922	-
Ahn (1994)	1976:03 - 1983:04	OLS	-	-	-	-	-	0.304	-
	1979:02 -1993:04	OLS	-	-	(0.018)	-	-	0.253	-
Haque and Montiel (1991)	1969 - 87	IV	-	0.865 (not different from 1)	-	0.638 (not different from 1)	0.577 (not different from 1)	-	-
Dooley and Mathieson (1994)	1964-89	OLS	-	0.972 (less than perfectly mobile)	0.95 (not different from 1)	0.91 (not different from 1)	1.11 (not different from 1)	-	0.91 (not different from 1)
		IV	-	0.66	0.95	0.94	1.13	-	0.85
Chinn and Frankel (1994)	1982:09 - 1992:03	OLS with time trend	1.276	(0.037)	-0.989 / (-0.066)	(0.013)	-	(0.809)	0.997
Frankel, Schmukler, and Serven (2000)	1990:01-1999:12	OLS	1.07	1.26	-	-	1.29	0.86	1.42
Frankel, Schmukler, and Serven (2002)	1990:01-1999:12	ADL / EC	0.91	-	-	-	2.16	0.99	1.4
Kim and Lee (2004)	1987:01 - 2002:4	First differencing	1.641	0.043	Before 2.946 After (-0.29)	Before (0.152) After (- 0.42)	1.731	(0.118)	Before 3.514 After (0.392)

Note: The estimated coefficients in parenthesis are not statistically significant from zero; "not different from one" refers to statistical significance.

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

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

Explanatory variable		(12) Static Expect.	(13) Static Expect.	(14) Perfect foresight	(15) Extrapolative Expectations	(16) Static Expect.	(17) Static Expect.
R_t^{USA}	*pre*	0.320** (0.180)	0.170*** (0.065)	0.003 (0.0189)	0.029 (0.027)	-0.263 (0.275)	0.043 (0.103)
	post	0.230** (0.025)	0.078*** (0.0124)	0.006*** (0.002)	0.004** (0.002)	0.322*** (0.049)	0.110*** (0.025)
Constant	*pre*	0.010*** (0.0008)	0.001** (0.0008)	0.011*** (0.0003)	0.011*** (0.0003)	0.052*** (0.013)	0.012** (0.005)
	post	0.003*** (0.00008)	0.001*** (0.0002)	0.004*** (0.00009)	0.004*** (0.0001)	-0.004 (0.003)	-0.001 0.002
R_{t-1}	*pre*	-	0.822*** (0.070)	-	-	-	0.735*** (0.067)
	post	-	0.659*** (0.042)	-	-	-	0.649*** (0.042)
$\ln y_t - \ln m_{t-1}$	*pre*	-	-	-	-	0.007*** (0.002)	0.002** (0.0008)
	post	-	-	-	-	-0.001* (0.0006)	- 0.0004* (0.0003)
INF_t	*pre*	-	-	-	-	-0.087* (0.040)	-0.026 (0.024)
	post	-	-	-	-	0.006 (0.005)	0.003 (0.003)
Summary Statistics							
Adj. R^2	*pre*	0.031	0.709	-0.011	0.006	0.368	0.717
	post	0.667	0.954	0.104	0.035	0.661	0.952

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
Effects of the U.S. and Regional Rates on Local Interest Rates

		US	IND	US	KOR	US	MAL	US	PHI	US	THA
IND	*Pre*			1.147*** *(0.458)*	0.120 (0.153)	1.452*** *(0.409)*	0.941*** *(0.219)*	1.193*** *(0.469)*	0.017 *(0.089)*	0.698* *(0.483)*	0.329** *(0.153)*
	Post[a]			0.043 (0.769)	-3.032 (3.365)	-0.129 (0.393)	5.773* (4.207)	-0.635* (0.446)	0.480 (0.383)	-0.282 (0.356)	1.644 (1.698)
KOR	*Pre*	0.237 (0.220)	0.073 (0.076)			0.512*** (0.174)	0.669** (0.322)	0.371** (0.179)	0.125* (0.074)	0.187 (0.255)	0.092 (0.129)
	Post	0.224*** (0.026)	0.016* (0.010)			0.223*** (0.020)	0.332*** (0.092)	0.132*** (0.044)	0.116** (0.051)	0.223*** (0.028)	0.275 (0.234)
MAL	*Pre*	0.462*** (0.103)	0.151*** (0.062)	0.340*** (0.094)	0.178*** (0.052)			0.270*** (0.109)	0.034 (0.092)	0.449*** (0.112)	0.112** (0.039)
	Post	-0.004 (0.014)	0.067*** (0.027)	- 0.303*** (0.075)	1.413*** (0.381)			-0.199** (0.035)	0.261** (0.151)	0.002 (0.052)	0.798* (0.505)
PHI	*Pre*	-0.432 (0.339)	0.045 (0.238)	-0.552* (0.341)	0.536*** (0.202)	-0.223 (0.440)	0.552 (0.587)			-0.326 (0.465)	-0.035 (0.186)
	Post	0.811*** (0.190)	0.091*** (0.026)	0.230 (0.197)	2.677*** (0.637)	0.816*** (0.173)	1.411*** (0.243)			0.813*** (0.190)	1.317** (0.653)
THA	*Pre*	1.177*** (0.291)	0.259** (0.113)	1.446*** (0.303)	0.119 (0.174)	1.639*** (0.277)	0.547** (0.261)	1.480*** (0.311)	-0.011 (0.058)		
	Post	0.018 (0.030)	0.018** (0.010)	-0.082 (0.066)	0.469* (0.282)	0.018 (0.032)	0.320*** (0.065)	-0.057 (0.066)	0.098* (0.067)		

a) Unlike other countries, the post-crisis period of Indonesia begins in July 1999 instead of January 1999 because of the much longer aftershocks following the crisis.

Table 6
Effects of the Changes in U.S. and Selected Regional Interest Rates on Changes in Local Interest Rate

		US	IND	US	KOR	US	MAL	US	PHI	US	THA
IND	*Pre*			0.573 (1.121)	-0.049 (0.135)	0.496 (0.981)	-0.203 (0.314)	0.274 (1.013)	0.062 (0.061)	0.475 (0.980)	-0.043 (0.069)
	Post[a]			-1.017 (1.429)	-2.890 (3.418)	-1.933* (1.429)	5.847 (10.847)	-1.643 (1.459)	-0.078 (0.585)	-1.781* (1.245)	-1.582 (1.473)
KOR	*Pre*	1.768*** (0.611)	-0.019 (0.050)			1.755*** (0.607)	0.086 (0.162)	1.728*** (0.611)	0.009 (0.022)	1.764*** (0.617)	0.016 (0.060)
	Post	0.152* (0.112)	-0.013 (0.010)			0.175** (0.102)	0.091*** (0.027)	0.096 (0.146)	0.076** (0.042)	0.182** (0.097)	0.111* (0.073)
MAL	*Pre*	0.049 (0.267)	-0.015 (0.014)	0.014 (0.265)	0.016 (0.031)			0.006 (0.266)	0.010 (0.010)	0.046 (0.270)	0.015 (0.027)
	Post	0.031 (0.025)	0.117 (0.145)	0.017 (0.097)	0.186** (0.110)			-0.076 (0.111)	0.115 (0.103)	0.055 (0.106)	0.197 (0.196)
PHI	*Pre*	3.221 (2.747)	0.504 (0.620)	3.132 (2.735)	0.190 (0.517)	3.417 (2.984)	1.174** (0.692)			3.410 (2.933)	-0.185 (0.191)
	Post	1.142** (0.541)	0.021 (0.038)	0.910* (0.553)	1.040** (0.492)	1.058** (0.568)	0.766*** (0.144)			1.106** (0.574)	0.346* (0.216)
THA	*Pre*	-0.274 (1.208)	-0.059 (0.078)	-0.405 (1.173)	0.058 (0.215)	-0.315 (1.199)	0.286 (0.510)	-0.196 (1.206)	-0.031 (0.030)		
	Post	-0.029 (0.160)	-0.001 (0.012)	-0.085 (0.158)	0.322 (0.279)	-0.041 (0.161)	0.279*** (0.078)	-0.108 (0.179)	0.074* (0.050)		

Data Appendix

R	Money market rate on short-term interest rates (line 60b of IFS) for every country except for (i) Hong Kong: Interbank Offered Rate (HIBOR) is from the Monetary Authority of Hong Kong (MAHK) and (ii) Singapore: Interbank Offered Rate (SIBOR) is from the Monetary Authority of Singapore (MAS).
R^f	U.S. Treasury bill rate (IFS)
M1	Narrow definition of money (line 34 of IFS), except Hong Kong where M1 is obtained from MAHK.
CPI	Consumer price index (CPI) (line 64 of IFS).
IP	Industrial production Index (line 66 of IFS); except for Thailand where the manufacturing production indexes are from the Bank of Thailand and Indonesia's indexes are from Total Production Indices for Large and Medium Manufacturing series of the Central Bank of Indonesia.
s	End of period of exchange rate (series *ae* from IFS) which expresses as units of national currency per U.S. dollar.
p	price level; defined as ln(cpi)
m1	Real money supply. Defined as $ln(M1) - ln(p)$
y	Real income, defined as $ln(IP)$