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GLOBAL INTEGRATION OF INDIA'S MONEY MARKET: INTEREST RATE PARITY IN INDIA

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Foreword

Financial openness exists when residents of one country are able to trade assets with residents of another country, i.e. when financial assets are traded goods. A weak definition of complete financial openness, which one might refer to as financial integration, can be given as a situation in which the law of one price holds for financial assets- i.e. domestic and foreign residents trade identical assets at the same price. A strong definition would add to this the restriction that identically defined assets e.g. a six-month Treasury bill, issued in different political jurisdictions and denominated in different currencies are perfect substitutes in all private portfolios. The degree of financial integration has important macroeconomic implications in terms of the effectiveness of fiscal and monetary policy in influencing aggregate demand as well as the scope for promoting investment in an economy.

The paper shows that the short-term (up to 3 month) money markets in India are getting progressively integrated with those in the USA even though the degree of integration is far from perfect. Covered interest parity is found to hold for while uncovered interest parity fails to hold. The difference between the two can be attributed to the existence of an exchange risk premium over and above the expected depreciation of the currency. Analysis of RBI interventions in response to foreign exchange shocks suggests that these may play a role in the deviations from interest parity. Further work needs to be done however on this as well as on instruments of other maturity such as 1 month and 6 month (for which consistent data was not available).

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1 INTRODUCTION

Financial openness exists when residents of one country are able to trade assets with residents of another country, i.e. when financial assets are traded goods. A weak definition of complete financial openness, which one might refer to as financial integration, can be given as a situation in which the law of one price holds for financial assets- i.e. domestic and foreign residents trade identical assets at the same price. A strong definition would add to this the restriction that identically defined assets e.g. a six-month Treasury bill, issued in different political jurisdictions and denominated in different currencies are perfect substitutes in all private portfolios. The degree of financial integration has important macroeconomic implications in terms of the effectiveness of fiscal and monetary policy in influencing aggregate demand as well as the scope for promoting investment in an economy.

The free and unrestricted flow of capital in and out of countries and the ever-increasing integration of world capital markets can be attributed to the process of Globalization. The benefits of such integration are liquidity enhancement on one hand and risk diversification on the other, both of which are instrumental in making markets more efficient and also facilitate smooth transfers of funds between lenders and borrowers. India began a very gradual and selective opening of the domestic capital markets to foreign residents, including non-resident Indians (NRIs), in the eighties. The capital market opening picked up pace during the nineties. In this paper we try and estimate the degree of financial integration between India and the rest of the World, by focussing on the degree of integration of the Indian money market with global markets.

Frenkel (1992) in his review of Capital Mobility measurement outlined four different definitions of perfect capital mobility that are in widespread use, of which three are of relevance to the current paper. These are real interest parity, uncovered interest parity and covered interest parity. (i) Real interest parity hypothesis states that international capital flows equalise real interest rates across countries. (ii) Uncovered interest parity states that capital flows equalise expected rates of return on countries' bonds regardless of exposure to exchange risk. (iii) Covered interest parity states that capital flows equalise interest rates across countries when contracted in the same currency. Frenkel (1992) shows that these three definitions are in ascending order of specificity in the following sense. Only definition (iii) that the covered interest differential is zero is an unalloyed criterion for "capital mobility" in the sense of the degree of financial market integration across national boundaries. Condition (ii) that the uncovered interest differential is zero requires that (iii) hold and that there be zero exchange risk premium. Condition (i) that the real interest differential be zero requires condition (ii) and in addition that expected real depreciation is zero.

2 LITERATURE REVIEW

The uncovered interest parity (UIP) theory states that differences between interest rates across countries can be explained by expected changes in currencies. Empirically, the UIP theory is usually rejected assuming rational expectations, and explanations for this rejection include that expectations are irrational, see Frankel and Froot (1990) and Mark and Wu (1998), or that time-varying risk premia are present, see Domowitz and Hakkio (1985) and Nieuwland et al. (1998), respectively. In a survey of 75 published estimates, Froot and Thaler (1990) report few cases where the sign of the coefficient on interest rate differentials in exchange rate prediction equations is consistent with the un-biased-ness hypothesis and not a single case where it exceeds the theoretical value of unity. This resounding unanimity on the failure of the predictive power of interest differentials is virtually unique in the empirical literature in economics.

A third explanation was provided by McCallum (1994a), who observes that regressing the change in spot exchange rates on the forward premium, one typically finds a negative regression parameter of -4 to -3 contrary to the expected parameter of +1. McCallum argues, however, that this finding may be consistent with the UIP theory, if one introduces policy behavior. Assuming policymakers adjust interest rates in order to keep exchange rates stable, and that they are interested in smoothing interest rate movements, McCallum derives a reduced form equation for the spot exchange rate under rational expectations. In fact, this results in a negative theoretical relationship between the change in the spot exchange rate and the forward premium consistent with his empirical findings. Christensen, M. (2000) extend the data set used by McCallum to include the recent 8 years and find that \$/DM, \$/£ and \$/Yen for the period 1978.01m to 1999.03m behave amazingly well according to the modified UIP theory developed by McCallum. However, when he estimates the policy reaction function, its structural parameters are inconsistent with the UIP relationships estimated.

Nevertheless, there appears to be overwhelming empirical evidence against UIRP, at least at frequencies less than one year (see Hodrick (1987), Engel (1996) and Froot and Thaler (1990)). Fama (1984) focuses on statistical properties of this relation. He finds that from the end of August 1973 to the end of 1982, the variance of the exchange risk premium has been large, exceeding the variance of expected future spot rates changes of the dollar against each of ten other major currencies (over monthly intervals). On the other hand Frankel and Froot (1987), among others, propose an explanation of UIP deviations based on the existence of asymmetries between currencies. Using survey data to approximate the exchange rates' behaviour, they show that agents were expecting a 10% depreciation of the Dollar against the Mark over 1981-85 whereas the differential in corresponding interest rates was only around 4%. Given that this empirical evidence has not stopped theorists from relying on UIRP, it is fortunate that recent evidence is more favourable. Bekaert and Hodrick (2001) and Baillie and Bollerslev (2000) argue that doubtful statistical inference may have contributed to the strong rejections of UIRP at higher frequencies. Chinn and Meredith (2001) marshal evidence that UIRP holds much better at long horizons. They test this hypothesis using interest rates on longer-maturity bonds for the U.S., Germany, Japan and Canada. The results of these long horizon regressions are much more positive — the coefficients on interest differentials are of the correct sign, and most are closer to the predicted value of unity than to zero. Ravi Bansal and Magnus Dahlquist (2000) conclude that **t**he often found negative correlation between the expected currency depreciation and interest rate differential is, contrary to popular belief, not a pervasive phenomenon. It is confined to developed economies, and here only to states where the U.S. interest rate exceeds foreign interest rates

The covered interest parity (CIP) postulates that interest rates denominated in different currencies are equal once you cover yourself against foreign exchange risk. Unlike the UIP, there is empirical evidence supporting CIP hypothesis. Empirical studies such as Frenkel and Levich (1975, 1977, 1981), Frankel (1989), among others, find that the CIP holds in most cases on the Eurocurrency market (where remunerated assets have similar default and political risk characteristics) since the collapse of the Bretton Woods regime in early 1970's. Lewis(1995) shows that risk premia do not vary significantly and often switch sign, contrary to what the observed stability of the countries' global creditor or debtor status would predict. However she explains that not only the conditional variance of exchange rate is not significant enough to account for risk premia movements, but also that risk premia examined in the short run should concern capital flows and investors with similar temporal horizons, such as currency traders, hedge funds and mutual funds managers. Frankel (1991) reports mean covered interest differentials (CIDs) for the period 1982 to 1987 for a selection of developed and developing economies using monthly observations of the 3-month local money market rate against the equivalent Eurodollar rate. Focusing on the East Asian economies in the sample – Japan, Hong Kong, Malaysia and Singapore – the null of a zero differential is rejected for the first three economies, though only marginally in that the CIDs are very low. Chinn and Frankel (1992) found that the CIDs were small for Japan, Hong Kong and Singapore, but large for Malaysia.

In the Indian context, Varma (1997) has undertaken an analysis of the covered interest parity. His posits a structural break in the money market in India in September 1995, with CIP become effective from that point on for the first time in the Indian money market. The structural break itself is attributed to interplay between the money market and the foreign exchange market. The period after 1995 is however witness to several deviations from the CIP. Varma has used rates on Treasury bills, certificates of deposit and commercial paper and call money rate to analyse the Indian money market. For the foreign rate he has calculated an implicit euro-rupee rate for six, three and overnight maturity. Thus he uses a mix of actual and constructed rates of different maturity. A rigorous test requires use of interest rates on identical instruments (e.g. maturity, risk) and a consistent forward rate (period of forwards should be identical to that of instruments). This is perhaps the first time that such a test is being carried out for India.

3 MODEL AND ESTIMATION

3.1 Estimating Equations

One of the key implications of international financial integration is on the degree of movement/co-movement of interest rates in countries over time and their comparison in terms of convergence or having a common trend. The relationship between two countries' interest rates is termed as interest rate parity.

The interest rate theory proposes that given perfect capital mobility, perfect capital market and fixed exchange rates the interest on identical assets (identical in terms of maturity etc) would be equal across countries. However, in the real world with capital controls, flexible exchange rates and imperfect capital markets divergence between interest rate is frequently observed and persist over long periods. Given the reality of non-frictionless capital markets and flexible exchange rates the recent versions of the interest parity theorem attribute this divergence to the expectation about exchange rate movements. Based on the preference individuals have for risk there are two versions of this basic relation:

a) Uncovered Interest Rate Parity- Assume that individuals are risk neutral. With no capital controls and perfect capital markets the interest differential between two countries is equal to change in exchange rate:

$i_t – i_t^* = S_{t+1}-S_t$.

where

i_t is domestic interest rate

i_t^*/ is foreign interest rate on similar asset (identical in all respects except for yield and currency denomination)

S_t is the spot exchange rate.

A risk neutral person would replace S_{t+1} by his expectation about future exchange rate. So we get

$i_t – i_t^* = E(S_{t+1}) – S_t$

Any deviation from UIP can be attributed to currency associated risks in the absence of hedging agreements- namely currency premium and expectation bias.

b) Covered Interest Parity- Assume that individuals are risk averse. Such an individual would like to cover himself for any unexpected currency fluctuation during the tenure of the deal. Given the forward contract market, he would purchase a forward contract and use the exchange rate mentioned in the contract. Then any difference in interest rate should be equated to forward premium. This is called CIP:

$i_t – i_t^* = F_t- S_t$

or

$i_t – i_t^* = f_t$

where F_t is forward rate and f_t is forward premium.

Any deviation from CIP would suggest that the markets are inefficient, regulations like capital controls exist and costs like sovereign risk, individual borrowing constraints are not accounted for.

3.2 Econometrics

To test the basic relation of interest rate parity we can think of a linear regression of the following type:

Equation 1: $\Delta S_t = \alpha + \beta (i_t - i_t^*) + \varepsilon_t$

For Uncovered Interest Parity we would expect α to be 0 and β to be equal to 1.

For covered interest parity we would use the following regression:

Equation 2: $f_t = \alpha + \beta (i_t - i_t^*) + \varepsilon_t$

and then test for $\beta = 1$.

The problem with using Ordinary Least Square as an estimation technique relates to the issue of non-stationarity of the time series involved in the above equation. In case of non-stationary times series the estimate of β would be spurious and biased. However if we can show that the two variables in question are cointegrated than the OLS estimates are super consistent and would converge to their true value faster (see). Thus before drawing inferences based on the results of ordinary least squares it is imperative to check the variables namely F (3-month forward premium) and IDIFF (3-month TB auction rate differential between India and U.S). In case the two series are integrated of the same order we can then test for cointegartion between the two non-stationary variables.

For covered interest parity we need to test for $\beta = 1$ where β is the coefficient of IDIFF. Formally,

Ho: $\beta = 1$ - Covered Interest Rate Parity holds

H1: $\beta \neq 1$ – There is no interest rate convergence.

The above test uses a standard t- statistic given by:

$t = (\beta - 1)/\sigma_\beta \sim t_{n-2}$

where σ_β is estimated standard error of β. Under the null hypothesis the above statistic follows a t-distribution with n-2 degrees of freedom.

3.3 Data Sources

The paper use the monthly data on following variables:

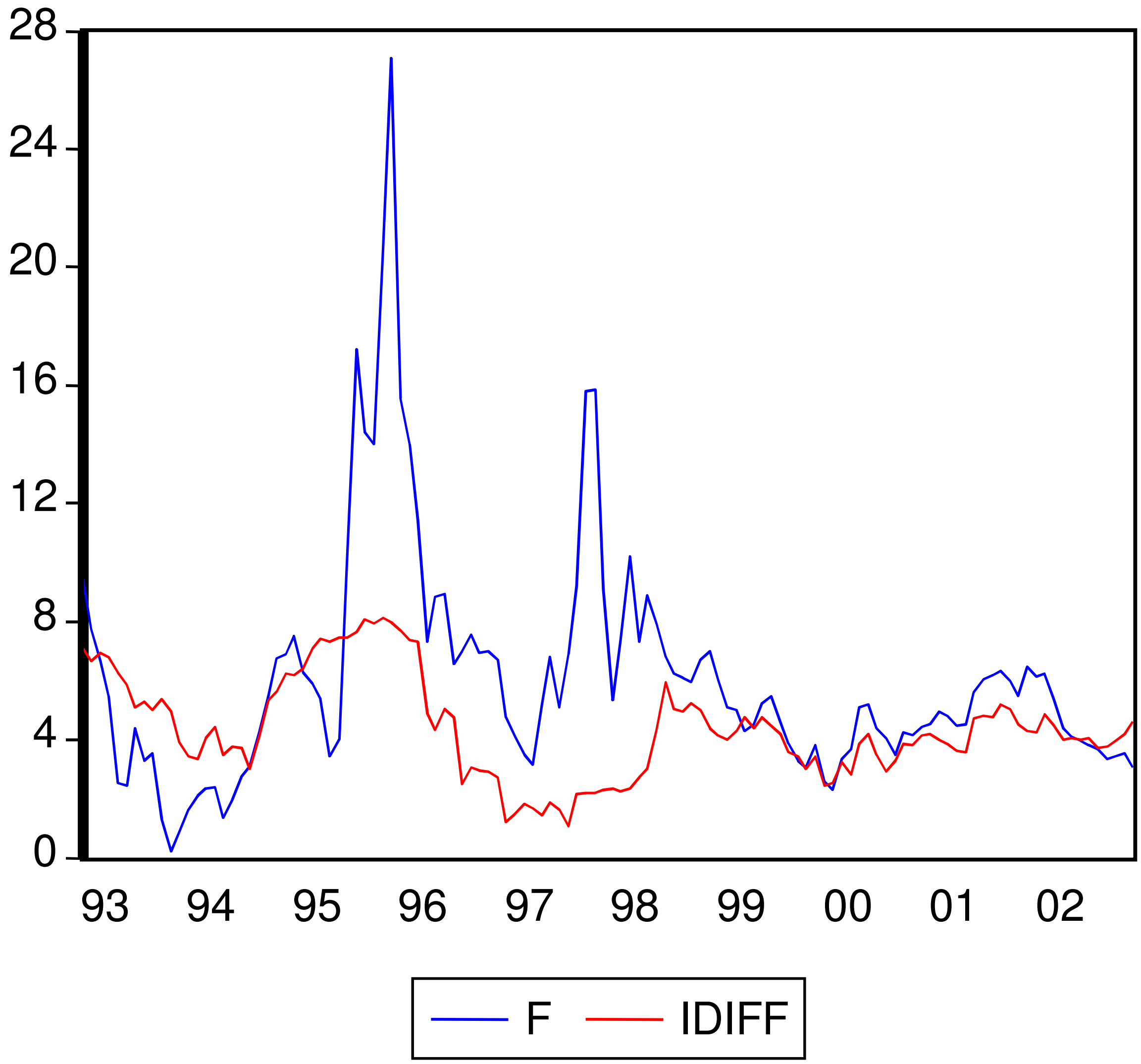
1) 3-month TB Auction rate for India- Apr 1993 to Mar 2003- Source: RBI Bulletin.
2) 3-month Forward Premia-Apr1993 to Mar 2003-Source: Handbook of Statistics, RBI
3) 6-month Forward Premia-Apr1993 to Mar 2003-Source: Handbook of Statistics, RBI
4) Call Money Rate-Apr1993 to Mar 2002-Source: Handbook of Statistics, RBI
5) 3-month TB Auction rate for U.S.- Apr 1993 to Mar 2003-Source: http://www.publicdebt.treas.gov/of/ofrespr.htm

List of Variables used in the Analysis:

Variable Name	Description
IDIFF	3-month TB interest differential between India and U.S
EDIFF	Change in Rs/Dollar Exchange Rate
F	3-Month Forward Premium
DCALL	Change in Call Money Rate
Sign1	Dummy Variable, which assumes value 1 if DCALL is positive.
Sign2	Dummy Variable, which assumes value 1 if DCALL is negative.

The key variables involved in CIP are plotted in Fig 1.This suggests that the degree of integration was low till mid-1998 and has increased dramatically since then.

Figure 1: 3-Month Forward Premium (F) and India-US Interest Differential (Idiff)



Note: *The sharp upward spikes represent RBI Interventions (short term tightening) in the financial market.

4 RESULTS

4.1 Stationarity and Co-integration

Since we are using high frequency time series data it is necessary to test for stationarity of the variables involved in above regressions. In case of non-stationarity, we need to show that the variables of same order of integration are cointegrated. The results for unit root tests are summarised in Table 1 which shows that variables F and IDIFF are I (1) processes while DCALL is I(0).

Table 1 *: Unit Root Test Results*

Variable	Test Type	Model Type[1]	t - value	5% Critical value	Interpretation
IDIFF	ADF	No Constant, No Trend	-1.09	-1.95	Unit Root with no constant and no trend in the data
	PP	No Constant, No Trend	-1.17	-1.95	Unit Root with no constant and no trend in the data
F	ADF	No Constant, No Trend	-1.49	-1.95	Unit Root with no constant and no trend in the data
	PP	No Constant, No Trend	-1.97	-2.58*	Unit Root with no constant and no trend in the data
*1% critical value					
DCALL	ADF	No Constant, No Trend	-2.09	-1.95	No Unit Root
	PP	Constant, No Trend	-7.74	-2.86	No Unit Root
Unit Root Tests in first differences of IDIFF and F					
IDIFF	ADF	Constant, No Trend	-3.46	-2.86	No Unit Root
	PP	Constant, Trend	-10.1	-3.41	No Unit Root
F	ADF	Constant, No Trend	-3.79	-2.86	No Unit Root
	PP	Constant, Trend	-9.62	-3.41	No Unit Root

After ascertaining that the variables are integrated of the same order, we select the order of the VAR. The AIC, SBC, and the likelihood ratio test collectively suggest an optimal lag length of 2. The next step is to test for Co-integration between IDIFF and F using Johansen's procedures. Both the maximum and trace eigen value statistics strongly reject the null hypothesis that there is no cointegration between the variables (i.e. r = 0), but do not reject the hypothesis that there is one cointegrating relation between the variables (i.e. r = 1) {Table 2}. Hence using least squares would yield super-consistent estimators. Note that DCALL is stationary and thus can be included as an exogenous policy variable in the interest parity equation.

Table 2: *Results of Johansen's Co-integration Test*

A) Co-integration LR Test Based on Maximal Eigen value of the Stochastic Matrix

118 observations from 1993 M6 to 2003 M3. Order of VAR = 2.
List of variables included in the co-integrating vector: F and IDIFF

Null	Alternative	Statistic	90% Critical Value	Interpretation
r = 0	r = 1	13.5006	12.9800	Co-integration at
r<= 1	r = 2	5.1903	6.5000	10%

B) Co-integration LR Test Based on Trace of the Stochastic Matrix

118 observations from 1993M6 to 2003M3 . Order of VAR = 2.
List of variables included in the co-integrating vector: F and IDIFF

Null	Alternative	Statistic	90% Critical Value	Interpretation
r = 0	r>= 1	18.6910	15.7500	Co-integration
r<= 1	r = 2	5.1903	6.5000	at 10%

4.2 Covered Interest Parity

Empirically, one finds more evidence for the CIP, which equates forward premium to the interest rate differential. For the period April 1993 to March 2003 we

[1] The choice of the model for conducting the unit root test is based on the sequential unit root testing procedure. See appendix for details.

have regressed F (3-montb forward premium) on IDIFF (Differential between 3-month TB rate for India and U.S). The estimated equation is:

Equation 3: F = 2.0 + 0.93* IDIFF

(2.1)** (4.7)*** [2]

R^2= 0.157, R^2 (adjusted) = 0.14, DW= 0.35.

After adjusting for AR(1) the equation becomes:

Equation 4: F = 3.0 + 0.65* IDIFF + 0.83*AR(1)

(1.7)* (2.0)** (15.8)***

R^2= 0.729, R^2 (adjusted) = 0.724, DW= 1.7.

The coefficient of IDIFF in above equation is 0.65. The calculated absolute value of t for the hypothesis test is 1.09, which is less than the critical value 2. So we can accept the Ho at 5% level of significance and conclude that CIP holds for the period under consideration. This shows that short-term money markets (3-month) in India are getting integrated with global (US) money markets even though the integration is far from perfect.

We would have liked to test the hypothesis for 1-month, 6-month and 1 year treasury bills, but a completely consistent data set is not available. In our view hybrid data sets do not provide a rigorous test (e.g. using 6 month forwards to test integration between one year securities).

4.3 Un-covered Interest Parity

The interest rate parity hypothesis postulates that with flexible exchange rates and non-frictionless capital markets the difference between the yield on identical assets in two countries could be explained by expected change in the exchange rate. Assuming perfect foresight we can test for uncovered interest rate parity by regressing change in spot exchange rate on interest rate differential and testing for the coefficient of interest rate differential being equal to 1. The estimated equation is as follows.

Equation 5: EDIFF = 0.229 - 0.021* IDIFF

[2] Terms in the brackets are t-ratios for respective parameter estimates. Significance at 10%, 5% and 1% level is represented by one, two and three stars respectively.

(1.35) (-0.52)[3]

R^2= 0.006, R^2 (adjusted) = -0.021, DW = 1.56.

The coefficient on interest rate differential is negative and close to 0.[4] Thus the UIP hypothesis fails in India. Given that CIP has been shown to hold during the same time period, this implies that the exchange risk premium for the Indian rupee is not zero (i.e. it is positive).

There have been a number of recognised external shocks during the nineties, such as the Mexican crisis and the Asian crises, that lead to heightened external uncertainty and increased foreign exchange risk perception. These were also situations in which the Central bank (RBI) intervened in the financial markets. The next section analyses the outcome.

4.4 Exchange Risk and RBI Intervention

As per the declared policy of the Reserve Bank of India (RBI), RBI intervenes to smooth out short term fluctuations in demand-supply balances arising from lumpy demand for foreign exchange (e.g. large repayment of debt) that it thinks will lead to excessive volatility given the thinness of the market. This intervention is commonly done through sale/purchase of foreign exchange. If the behaviour of the RBI is completely symmetric with zero sterilisation, we would expect symmetric effects on call markets (increased/reduced liquidity) and on forward rates (higher/lower reserves). The higher the degree of sterilisation the less the effect of foreign inflow on liquidity and more asymmetric the relationship between call rates and forward rates (i.e. rising call rates have larger co-efficient than falling ones).

The RBI also intervenes to counter sharp adverse changes in expectations, like those arising from domestic and global political developments (e.g. post Pokharan sanctions, Kargil war) and external crisis such as the Mexican and Asian crisis. This intervention is commonly done through short-term instruments (overnight and 7-day repos, bank rate/moral suasion of banks), and translates into sharp upward movement in the inter-bank call money market rates. These in turn are reflected in a rise in foreign exchange forward rates. It is only at the time of the next auction, however, that these developments get reflected in the T-bill auction rates.[5] Such tightening is

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3 Terms in the brackets are t-ratios for respective parameter estimates

4 Similar results are obtained after adjusting for auto regression (DW rises to 1.9 with AR).

5 The secondary market rates on T-bills are available for too short a period to do statistically credible tests.

generally followed in due course by a loosening to the starting position, but forwards may not revert to the original level given the residual uncertainty.

In order to gauge the impact of policy changes on the interaction between forward rate and interest differential we have re-estimated the above equation in the following form:

$$F_t = \alpha + \beta 1 \text{ IDIFF} + \beta 2 * \text{DCALL} + \varepsilon_t$$

where DCALL- is the change in the call money rate. The estimated equation corresponding to Equation 4 is given by,

Equation 6 : F = 3.7 + 0.58*IDIFF + 0.10*DCALL + 0.84*AR(1)

(2.0)* (1.8)* (3.4)*** (15.8)***

R^2= 0.75, R^2 (adjusted) = 0.74, DW= 1.6.

The estimated coefficient of the interest differential has now fallen from 0.65 to 0.58. However, to see whether it is statistically different from 1 we would perform the t- test for the restriction $\beta = 1$ again. Under Ho of $\beta = 1$ the t-statistic mentioned in (iii) follows t-distribution with n-3 degrees of freedom. The calculated value for the test $\beta = 1$ is 1.28, which is less than the critical value so that the CIP hypothesis still holds.

External shocks and RBI exchange market stabilisation efforts through the short-term money market seem to loosen the link between the domestic and foreign money markets.

5 CONCLUSIONS

The paper shows that the short-term (up to 3 month) money markets in India are getting progressively integrated with those in the USA even though the degree of integration is far from perfect. Covered interest parity is found to hold for while uncovered interest parity fails to hold. The difference between the two can be attributed to the existence of an exchange risk premium over and above the expected depreciation of the currency. Analysis of RBI interventions in response to foreign exchange shocks suggests that these may play a role in the deviations from interest parity. Further work needs to be done however on this as well as on instruments of other maturity such as 1 month and 6 month (for which consistent data was not available).

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7 APPENDIX: TESTS

7.1 *Order of integration*

7.1.1 Sequential ADF Test for unit root

Step 1: Estimate

$\Delta y_t = a_0 + a_2 t + \gamma y_{t-1} + \sum \beta i \Delta y_{t-i+1} + \in_t$

Test-Ho : $\gamma = 0$ using ι_t statistics

If Ho is rejected, no need to proceed. Conclude that {Yt} is stationary. If Ho is not rejected it is necessary to determine whether too many deterministic regressors were included. First test for the significance of trend:

Ho:a_2=0 given $\gamma = 0$ use $\iota_{\beta\iota}$

We may also gain additional information by testing

Ho: $a_2 = \gamma = 0$ use ϕ_3

If trend is significant the retest for the presence of a unit root (i.e. γ=0) using the standardized normal distribution. If the Ho is rejected proceed no further-conclude that y_t

is stationary otherwise it is non-stationary.

If a_2 is not significant move to step 2.

Step 2:Estimate

$\Delta y_t = a_0 + \gamma y_{t-1} + \sum \beta i \Delta y_{t-i+1} + \in_t$

Test Ho: $\gamma = 0$ use ι_μ

If Ho is rejected, conclude that {y_t} is stationary. If Ho is not rejected then test for the significance of drift a_0:

Ho: a_0=0 given $\gamma = 0$ use $\iota_{\alpha\mu}$

We may also gain additional information by testing

Ho: $a_0 = \gamma = 0$ use ϕ_1

If drift is significant the retest for the presence of a unit root (i.e. γ=0) using the standardized normal distribution. If the Ho is rejected proceed no further-conclude that y_t

is stationary otherwise it is non-stationary.

If a_0 is not significant move to step 3.

Step 3:Estimate

$\Delta y_t = \gamma y_{t-1} + \sum \beta i \Delta y_{t-i+1} + \in_t$

Test Ho: $\gamma = 0$ use ι

If Ho is rejected, conclude that $\{y_t\}$ is stationary otherwise it is nonstationary.

7.1.2 Phillips-Perron Test

This test rests on very mild assumptions regarding the distribution of the errors and can be used even if there is serial correlation and hetroscedasticity. In the ADF tests we assumed that the errors are white noise i.e. they are statistically independent and have a constant variance.

Consider $y_t=a_0+a_1y_{t-1}+\mu_t$

$y_t=b_0 +b_1yt-1+b_2(t-T/2) +\mu_t$

T-number of observations

E $(\mu_t)=0$ but there is no requirement that disturbance term is serially uncorrelated and homogeneous.

This test develops the test statistics for testing the presence of unit root by assuming that y_t can be generated under Ho by a random walk process.

Ho: $y_t = y_{t-1} + \mu_t$

The Phillips-Perron test statistics are the modifications of the DF statistics that take into account the less restrictive nature of the error process. Some useful test statistics are:

$Z(ta_1)$: used to test Ho:$a_1=0$ - use ι_μ

$Z(tb_1)$: used to test Ho: $b_1=0$ -use ι_ι

$Z(\phi_3)$: used to test $b_1=0$ and $b_2=0$ -use ϕ_3

7.2 Test for Co-integration: Johansen's Methodology

Given a group of non-stationary series we may be interested in determining whether the series are co-integrated.

Step 1: Test for the order of integration using the DF, ADF, PP tests

Step 2:Selection of the appropriate lag length. The result of the test can be quite sensitive to the lag length

The most common procedure is to estimate a VAR using undifferenced data. Then use lag length tests as in VAR:

a) $AIC = T \ln|\Sigma| + 2N$

Where N total number of parameters estimated in all equations

$|\Sigma|$ - natural log of the determinant of the var-cov matrix of the residuals.

b) $SBC = T \ln|\Sigma| + N \ln T$

c) $LR = (T-c)[\ln|\Sigma_r| - \ln|\Sigma_u|]$

where T-number of usable observations

c= number of parameters in the unrestricted system

$\ln|\Sigma_i|$ - natural log of determinant of Σ_i $\qquad$ i= u, r

The above statistic follows χ^2 with degrees of freedom equal to the number of restrictions imposed.

Step 3: Estimate the model and determine the number of cointegrating relationships. If you have k endogenous variables, there can be from zero to k-1 linearly independent cointegrating relations. There are two formal test for determining the number of cointegarting relationships:

a) Trace test :

$Qr = -T \sum \log(1-\lambda i)$ for r =0,1…k-1

where λi is the i-th largest eigenvalue. This statistic tests the Ho® against Ha (k).

To determine the number of cointegrating relations r, subject to the assumptions made about the trends in the series, we can proceed sequentially from r = 0 to r = k-1 until we fail to reject.

b) Maximum Eigen Value test:

$Q_{max} = -T \log (1-\lambda_{r+1}) = Qr - Q_{r+1}$

This statistic is used to test the Ho of r co-integrating relationships against r+1 relations.

7.3 Call Money Assymmetry

If the behaviour of the RBI is completely symmetric with zero sterilisation, we would expect symmetric effects on call markets (increased/reduced liquidity) and on forward rates (higher/lower reserves). The higher the degree of sterilisation the less the effect of foreign inflow on liquidity and more asymmetric the relationship between call rates and forward rates (i.e. rising call rates have larger co-efficient than falling ones). Historically there has been incomplete sterilisation and therefore we expect some asymmetry in the relationship between call money rates and forward rates, though these interventions are very short term and may not even appear in data of monthly frequency.

RBI tightening is generally followed in due course by a loosening to the starting position, but forwards may not revert to the original level given the residual uncertainty. Thus the absolute impact co-efficient relating call rates to forwards is likely to be larger on the upside than on the downside. As the time period of such intervention is of the order of a week and movements are quite sharp, such asymmetry can be observed in monthly data.

$F_t = \alpha + \beta 1 \text{ IDIFF} + \beta 2 \text{ (Sign1*DCALL)} + \beta 3 \text{ (Sign2*DCALL)} + \varepsilon_t$

where DCALL- is the change in the call money rate and sign1 and sign2 are dummies that separate positive changes in call money rates from negative changes

The estimated equation corresponding to Equation 3 is given by,

F = 1.88 + 0.688*IDIFF + 0.654*(SIGN1*DCALL) – 0.454*(SIGN2*DCALL)

(2.3)*** (3.9)*** (6.3)*** (-4.7)***

R^2= 0.44, R^2 (adjusted) = 0.42, DW= 0.61.

The estimated coefficient of interest differential has now fallen from 0.93 to 0.69. However, to see whether it is statistically different from 1 we would perform the t- test for the restriction $\beta = 1$ again. Under Ho of $\beta = 1$ the t-statistic mentioned in (iii) follows t-distribution with n-4 degrees of freedom. The calculated value for the test $\beta = 1$ is 1.78 which is less than the critical value so that we can accept the CIP hypothesis. The same conclusion follows after adjusting for the auto regressive nature of the error term.

F = 3.1 + 0.58*IDIFF + 0.35*(SIGN1*DCALL) – 0.11*(SIGN2*DCALL) + 0.83*AR(1)

(1.9)* (2.0)* (5.9)*** (-2.1)** (15.2)***

R^2= 0.80, R^2 (adjusted) = 0.79, DW= 1.9.

The calculated value for the test $\beta = 1$ is 1.40, which is less than the critical value so that we can accept the CIP hypothesis.