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Asymmetric Arbitrage and Default Premiums Between the U.S. and Russian Financial Markets

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Deviations from covered interest rate parity (CIP) and from a generalized form of CIP involving forward-forward arbitrage between the Russian Treasury bill (GKO) market and the U.S. Treasury bill market are modeled nonlinearly. We find a no-arbitrage band within which deviations are random, outside of which deviations revert to the edge of the band. The band is asymmetric, implying that small profit margins trigger arbitrage into the dollar, but large profit margins are needed to trigger arbitrage into the ruble. The bandwidth rises and the speed of mean reversion falls as the maturity increases. The findings are consistent with the existence of Russian default premiums. [JEL F31, G15]

This paper examines covered interest rate arbitrage—that is, interest rate arbitrage where the risk of foreign exchange rate movements are covered in the forward exchange market—between the Russian ruble Treasury bill (GKO) market and the U.S. Treasury bill market, using daily data for the period December 1996 to August 1998. In contrast to numerous studies of covered interest rate arbitrage between developed financial markets,[1] significant deviations from covered interest rate parity (CIP) between the ruble and the dollar are discovered. Interestingly, however, these deviations are asymmetric in the sense that much larger deviations from parity are in arbitrage involving borrowing in dollars and lending in rubles. We therefore interpret these deviations as evidence of significant default

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[1]See, for example, Taylor (1992) or Sarno and Taylor (2002) for a survey of the relevant literature.

premiums, which proved rational given the foreign debt moratorium announced by the Russian authorities in mid-August 1998. Further innovations in this study include our examination of arbitrage between different points of the maturity term structure, rather than just between instruments of the same maturity, and our allowance for asymmetric adjustment to covered interest rate arbitrage opportunities using recently developed nonlinear econometric techniques.

I. No-Arbitrage Relationships

The CIP condition is normally expressed as the condition that the interest differential between financial assets of the same term denominated in different currencies is equal to the cost of covering in the forward market the currency risk from arbitrage between the two instruments, which arises from possible movements of the exchange rate before the assets mature. In symbols, if S_t is the spot exchange rate (domestic price of foreign currency) at time t, $i(k)_t$ and $i(k)_t^*$ represent the domestic and foreign k-period interest rates on the assets concerned (expressed on a k-period rather than on an annualized basis), and $F(k)_t$ is the forward exchange rate (the rate at which a future exchange of currencies is agreed to now) of the same term as the assets, then the CIP condition may be expressed as[2]

$$\frac{F(k)_t}{S_t} = \frac{1 + i(k)_t}{1 + i(k)_t^*}. \tag{1}$$

If equation (1) does not hold, and if the risk of default of one or more of the counterparties in the transactions is assumed to be zero, then it must be possible to arbitrage risklessly and profitably from one of the currencies into the other—by borrowing one currency, selling it in the spot foreign exchange market for the other currency, lending the other currency, and then selling the proceeds of the maturing (second currency) asset in the forward foreign exchange market. Assuming that the normal laws of supply and demand hold in the relevant financial markets, this arbitrage will tend to ensure that the equality holds. Since the forward transaction eliminates foreign exchange risk exposure, any deviation from CIP must represent either market inefficiency or a premium arising from perceived risk of default.

While the literature has tended to concentrate on the form of CIP expressed in equation (1), a more general form of covered interest rate arbitrage would involve arbitrage along the term structure. For example, by simultaneously lending—at time t—the domestic currency for j periods and borrowing the same amount of domestic currency for k periods, $j < k$, a forward cash flow of domestic currency is created from the beginning of period $t + j + 1$ (or end of period $t + j$) through to the end of period $t + k$. This can then be arbitraged into the foreign currency for periods $t + j + 1$ to $t + k$ by selling the domestic currency in the forward foreign exchange market j periods ahead, simultaneously borrowing and lending the foreign currency for j and k periods, respectively, and selling the foreign currency in

[2]Note that, since the data we shall analyze are averaged over the bid and ask prices, we shall follow much of the literature in ignoring the bid-ask spread for the purposes of exposition. See Taylor (1987, 1989) and Ghosh (1998) for a discussion of the bid-ask spread in this context.

the forward market k periods ahead. Clearly, it may also be profitable to do the arbitrage in reverse, or it may be the case that neither arbitrage is profitable. If neither arbitrage is profitable, then the following generalized CIP condition must hold:

$$\frac{F(k)_t}{F(j)_t} = \left(\frac{1 + i(k)_t}{1 + i(j)_t}\right)\left(\frac{1 + i(j)_t^*}{1 + i(k)_t^*}\right). \tag{2}$$

The CIP condition expressed as equation (2) may be viewed as a generalization of equation (1) by setting $j = 0$ and noting that a zero-period forward exchange rate is just the spot rate, $F(0)_t = S_t$, and the zero-period interest rate must be zero, $i(0)_t = i(0)_t^* = 0$.

Note that our generalized CIP condition may also be viewed as a generalization of the forward-forward CIP condition first examined by Ghosh (1998), assuming that arbitrage ensures parity between synthetic forward interest rates and actual forward interest rates. Ghosh's forward-forward CIP condition may be written in our notation (recalling that $j < k$) as[3]

$$\frac{F(k)_t}{S_t} = \left(\frac{1 + i(j)_t}{1 + i(j)_t^*}\right)\left(\frac{1 + i(j,k)_t}{1 + i(j,k)_t^*}\right), \tag{3}$$

where $i(j,k)_t$ is the forward interest rate at time t for deposits or loans from the end of period $t + j$ through to the end of period $t + k$ and, as before, an asterisk denotes a foreign variable. The equivalence of equation (3) with our general formulation (2) may be seen as follows: Consider a forward deposit agreement whereby an investor agrees at time t to place a deposit of one unit at the rate $i(j,k)_t$ from the end of period $t + j$ to the end of period $t + k$ (a symmetric argument would apply to a forward loan agreement). The same cash flow at time t and at the end of period $t + j$ can be obtained by simultaneously placing a deposit of $1/(1 + i(j)_t)$ for k periods and taking a loan of $1/(1 + i(j)_t)$ for j periods; this is, in fact, a synthetic forward deposit agreement. Given that there is the same degree of risk and commitment of funds as involved in the direct forward deposit agreement, the total return to the synthetic and the direct forward agreements must be the same. The synthetic agreement will produce a cash flow of $(1 + i(k)_t)/(1 + i(j)_t)$ at the end of period $t + k$, while the forward agreement will produce a cash flow of $(1 + i(j,k)_t)$. Equating these, assuming a similar condition holds for the foreign interest rates, and substituting into equation (3), yields equation (1), which we have already shown to be a special case of equation (2).

To a close approximation, (2) can be expressed in the form

$$f(k)_t - f(j)_t = id(k) - id(j)_t, \tag{4}$$

where $f(j)_t$ and $f(k)_t$ denote the natural logarithms of $F(j)_t$ and $F(k)_t$, respectively, and $id(j)_t \equiv [i(j)_t - i(j)_t^*]$ and $id(k)_t \equiv [i(k)_t - i(k)_t^*]$ denote the domestic-foreign

[3]Equation (3) is equivalent to Ghosh's equation (5) (Ghosh, 1998, p. 119), assuming that the profit level from the arbitrage is zero (which is correct in the present context, since in our exposition we are considering no-arbitrage conditions).

j- and k-period interest rate differentials, respectively. This is the general form of CIP that will be tested empirically in this paper.[4] We will consider deviations from both standard CIP, where $j = 0$ and $k > 0$, and deviations from generalized CIP, where $k > j > 0$.

II. Covered Interest Rate Parity: A Brief Review of the Empirical Literature

Probably the earliest written discussion of CIP was published in the early 1920s by John Maynard Keynes, first in an article in a British newspaper (Keynes, 1922) and subsequently in a monograph on monetary economics, the *Tract on Monetary Reform* (Keynes, 1923):

> Forward quotations for the purpose of the currency of the dearer money market tend to be cheaper than spot quotations by a percentage per month equal to the excess of the interest which can be earned in a month in a dearer market over what can be earned on the cheaper. (Keynes, 1923).

Writing of the 1920s international financial markets, Keynes argued that a deviation from the CIP would only generate sufficient profit to generate substantial arbitrage if the deviation exceeded 50 basis points on an annualized basis, and that even then the deviation might be arbitraged relatively slowly.

There are a number of published empirical studies of CIP and covered interest rate arbitrage for the post–World War II period (see Officer and Willett, 1970; Taylor, 1992; and Sarno and Taylor, 2002, for surveys). To date, essentially two types of empirical tests of CIP have been conducted. The first type has involved the use of econometric regression analysis. Thus, if CIP holds, and in the absence of transaction costs, estimation of the following equation:

$$\frac{F(k)_t - S_t}{S_t} = \alpha + \beta[i(j)_t - i(j)_t^*] + u(j)_t \tag{5}$$

where $u(j)_t$ is the regression error, should result in estimates of α and β differing insignificantly from zero and one, respectively, and a non-autocorrelated error. Equation (5) has been tested by a number of researchers for a variety of industrialized countries' exchange rates and time periods (e.g., Grubel, 1966; Branson, 1969; Marston, 1976; Cosandier and Lange, 1981; and Fratianni and Wakeman, 1982), with the general conclusion that, with important exceptions due to factors such as political risk (see Aliber, 1973; Frankel and MacArthur, 1988; and Branson and Jaffee, 1990) CIP is not rejected.[5]

The second type of test relies on computing actual deviations from interest parity to see if they differ significantly from zero. The significance is often defined

[4]Note that it is important that both sides of equation (4) are expressed on the same basis—that is, annualized or per period (e.g., three-month or monthly). In the empirical work below, we converted the forward premium to an annualized basis and also used the interest rate differential in annualized terms so that the deviations from CIP are also in annualized percentage terms.

[5]Officer and Willett (1970) survey much of the earlier postwar literature. See also Taylor (1992, 1995).

with respect to a neutral or no-arbitrage band, which is determined by transaction costs such as the bid-ask spread and brokerage fees (Demsetz, 1968). For example, the classic study of Frenkel and Levich (1975) demonstrates that for arbitrage between sterling, U.S. dollar, and Canadian dollar securities, 80 to 90 percent of apparent profit opportunities in weekly data over the period January 1962 through November 1967 lie within a neutral band when 90-day Treasury bills are used and almost 100 percent when euro-deposit rates are considered.[6] Furthermore, in Frenkel and Levich (1977) it is demonstrated that in periods of turbulence a much smaller percentage of deviations from CIP may be explained by transaction costs. This is interpreted as reflecting higher financial uncertainty in such periods. Clinton (1988) demonstrates that deviations from CIP should be no greater than the minimum transaction costs in one of three markets: the two underlying deposit markets (e.g., euro-marks and euro-dollars) and the foreign exchange swap market (i.e., the market in which a currency can be simultaneously bought and sold forward against another currency). On the basis of analysis of data for five major currencies against the U.S. dollar taken from mid-morning quotes on the Reuters Money Rates Service from November 1985 to May 1986, Clinton finds that the neutral band should be within ± 0.06 percent per year from parity and that although the hypothesis of zero profitable deviations from parity can be rejected, "empirically, profitable trading opportunities are neither large enough nor long-lived enough to yield a flow of excess returns over time to any factor."

Taylor (1987, 1989) uses high-quality, high-frequency contemporaneously sampled data for spot and forward dollar-sterling and dollar-mark exchange rates and corresponding euro-deposit interest rates for a number of maturities and makes allowance for bid-ask spreads and brokerage costs in his calculations for the 1980s and selected postwar periods. He finds, inter alia, that there are few profitable violations of CIP, even during periods of market uncertainty and turbulence. One interesting feature of this research is the finding that where profitable violations of CIP do occur, their frequency, size, and persistence appear to be an increasing function of the length of the period to maturity of the underlying financial instruments. A rationale is offered for this in terms of banks' prudential credit limits: since banks impose prudential limits on the amount of outstanding liabilities they have with counterparties, arbitraging at the shorter maturities will result in limits being filled for shorter periods, leaving dealers on average freer to take advantage of other profit opportunities as they arise.

An interesting but little-highlighted theme in this literature is that the size of the minimum deviation from CIP necessary in order to induce arbitrage appears to have shrunk over time. As noted above, Keynes (1923) and subsequent writers on the interwar financial markets, such as Einzig (1937), suggested that the minimum arbitrage opportunity should be of the order of 50 basis points on an annualized basis. Although this figure appeared largely as a conjecture based on personal experience in the markets, recent work by Peel and Taylor (2002) on covered interest rate arbitrage between the London and New York markets during the 1920s appears to have

[6]Arbitrage with Canadian dollar securities is not included in the Frenkel-Levich analysis of euro-deposits due to a lack of data.

borne this conjecture out empirically, using threshold estimation techniques. For the postwar period, Branson (1969), using data for U.S. dollar arbitrage against the Canadian dollar and pound sterling, found that the minimum arbitrage opportunity appeared to have shrunk to around 0.18 percent on an annualized basis.[7]

In this paper, in what is to the best of the authors' knowledge the first empirical study of arbitrage between the Russian and Western financial markets, we apply nonlinear regression modeling techniques to estimate the minimum deviation from CIP required for covered arbitrage between the Russian ruble and the U.S. dollar. Our study may thus be viewed as a synthesis of the two standard methods of testing for CIP discussed above—that is, regression-based analysis and "neutral band" or "no-arbitrage band" analysis. It is important to note, however, that the "no-arbitrage bands" that we identify in our empirical analysis may be considered as an indication of the market's implicit assessment of the probability of default, rather than merely reflecting transaction costs. Since, in our opinion, they reflect a view of the risk of default that is far from neutral, we prefer to refer to these bands as "no-arbitrage bands" rather than "neutral bands."

III. Default Risk in Russian-U.S. Arbitrage: Background to the Study

This paper considers generalized covered interest rate arbitrage between U.S. dollar Treasury bills and Russian ruble Treasury bills or GKOs. Our sample period of daily data ends in early August 1998. After repeated and sustained downward pressure on the ruble, the Russian authorities allowed the ruble to float in mid-August 1998, widening its trading band to between 6 and 9.5 rubles to the dollar, effectively allowing a devaluation of some 34 percent. At the same time, a three-month moratorium on much foreign debt repayment was imposed. This followed sharp declines on the Russian stock market and a sharp rise in short-term interest rates to above 150 percent on an annualized basis earlier in the same month.

Although these measures followed the announcement of a support package of over $22 billion from the International Monetary Fund and the World Bank that had been announced in July, fears of devaluation and also of at least a temporary debt moratorium had clearly been prevalent in the market. Indeed, in early August a leading and highly influential financier, George Soros, called for a 15–25 percent devaluation of the ruble and the setting up of a currency board (see, for example, the *Economist,* August 15, 1998). Although the sudden devaluation of the ruble would not have affected interest rate arbitrageurs who had sought cover in the forward markets, one might nevertheless have expected to see important deviations from CIP in the period before the August 1998 crisis, as market traders factored in the risk of a foreign debt moratorium on funds they had arbitraged into rubles, and this is the motivation for the present empirical study. As a first attempt to identify these default premiums, and given that we are examining only data spanning a roughly two-year period (although we use daily data in order to increase the num-

[7]Possible reasons for the reduction over time in the minimum deviation from CIP needed for covered arbitrage activity are discussed in Peel and Taylor (2002). See Balke and Wohar (1998) for an application of nonlinear methods to deviations from covered interest arbitrage among major currencies during the post-Bretton Woods floating rate period.

ber of observations), we feel justified in our implicit assumption that they remained roughly constant over the period in question. As we note in our concluding section, the relaxation of this assumption remains an issue for future research.

It is worth reflecting on the use of the terms "arbitrage" and especially "covered arbitrage" and their derivative terms in the context of the present analysis. In particular, the term "covered arbitrage" or "covered interest rate arbitrage" implies that advantage is being taken of different interest rates being offered on assets denominated in different currencies, with the risk involved being "covered" in the forward foreign exchange market. In more standard analyses of covered interest arbitrage, other risks—and in particular default risks—are usually considered negligible. For the kinds of transactions traders were making in the period immediately preceding the Russian crisis, however, default risks were non-negligible, manifold, and impossible to cover and included the risks that one or more legs of the interest arbitrage transaction would not be honored by the various counterparties. Thus, the notion of being able to separate the probability of default from the "loss given default" would be very difficult, because one would need to know what type of counterparty was expected to default. There are at least two types of default risk: the sovereign could default on the GKO securities and/or the Russian banks involved (or the Moscow Interbank Currency Exchange (MICEX) if it was acting as the counterparty) could default on the forward foreign exchange contracts. In the Russian crisis, both types of default in fact occurred. Further research might, therefore, usefully attempt to disentangle these types of default risk, although this is likely to be a very difficult task. What is clear, however, is that the "no-arbitrage bands" that we seek to identify in our empirical analysis are likely to reflect the market's assessment of the combined effect of the risk of default from these various sources, especially with respect to arbitrage from dollars to rubles.

IV. Modeling Deviations from CIP and Generalized CIP

From equation (4), the generalized CIP condition, define the deviation from CIP for arbitrage between maturities j and k, $\delta(k,j,)_t$, as[8]

$$\delta(k,j,)_t \equiv [f(k)_t - f(j)_t] - [id(k)_t - id(j)_t] \tag{6}$$

Our empirical methodology involves estimating nonlinear time series models for $\delta(k,j,)_t$. We consider both departures from the traditional CIP relationship, where arbitrage occurs between the spot ($j = 0$) and a forward market ($k > 0$), and departures from the generalized form of CIP discussed above, where arbitrage occurs between the spot and a forward market ($k > j > 0$). In particular, we wish to estimate the no-arbitrage band, within which arbitrage does not take place, and the speed of mean reversion of deviations from CIP outside the no-arbitrage band. A parametric model that may capture this nonlinear behavior—and that nests both instantaneous and slower mean reversion toward the band—is the threshold

[8]As discussed in footnote 4, it is important that the forward premium and the interest rate differential be expressed on the same basis. In our empirical work, they were each expressed on an annualized basis.

autoregressive (TAR) model (Tong, 1990; Granger and Teräsvirta, 1993). A simple TAR formulation—known as the asymmetric band-TAR model—that allows deviations from covered interest parity, δ_t (where we have dropped the *j,k* arguments for clarity), to follow a random walk within a band with an upper threshold of κ_1 percent and a lower threshold of κ_2 percent ($\kappa_1 > \kappa_2$) while exhibiting mean-reverting behavior outside of the band may be written

$$\delta_t = \delta_{t-1} + \varepsilon_{1t} \qquad \text{if } \delta_{t-1} < \kappa_1 \text{ and } \delta_{t-1} > \kappa_2, \tag{7a}$$

$$\delta_t - \kappa_1 = \rho(\delta_{t-1} - \kappa_1) + \varepsilon_{2t} \qquad \text{if } \delta_{t-1} \geq \kappa_1, \tag{7b}$$

$$\delta_t - \kappa_2 = \rho(\delta_{t-1} - \kappa_2) + \varepsilon_{2t} \qquad \text{if } \delta_{t-1} \leq \kappa_1, \tag{7c}$$

where $\varepsilon_{it} \sim N(0, \sigma_i^2)$ I = 1,2, $\rho \in (0,1)$, and $\kappa_1 > \kappa_2$. Although this band-TAR model is relatively simple, it can distinguish between subtly different forms of behavior. It implies that within the band (κ_2, κ_1) deviations from CIP essentially follow a random walk and therefore show no tendency to mean-revert toward zero—which would follow if there were no arbitrage within the band. Outside of this interval, however, the deviations become mean reverting and tend to revert toward the edge of the band, which would be consistent with arbitrage activity once the deviation exceeds the band.

In the absence of prior knowledge about the bandwidth, this model cannot be estimated by simple least squares methods. The method of maximum likelihood can, however, be applied to provide estimates of all the unknown parameters, including the bandwidth. The log-likelihood function for this model, ignoring the constant term in π, may be written

$$\begin{aligned} L(\rho, \theta, \sigma_1^2, \sigma_2^2, \kappa_1, \kappa_2) = &-\tfrac{1}{2} \sum_{\delta_{t-1} \in (\kappa_2, \kappa_1)} \left[\ln(\sigma_1^2) + \varepsilon_{1t}^2/\sigma_1^2\right] \\ &-\tfrac{1}{2} \sum_{\delta_{t-1} \notin (\kappa_2, \kappa_1)} \left[\ln(\sigma_2^2) + \varepsilon_{2t}^2/\sigma_2^2\right]. \end{aligned} \tag{8}$$

If κ_1 and κ_2 were in fact known, then maximum likelihood estimation would be fully equivalent to ordinary least squares applied to the three subsets of data partitioned according to whether the CIP deviation is within, below, or above the no-arbitrage band indicated by the thresholds, with a restriction that the variance of the residuals outside the band have the same variance (which could be achieved by using a dummy variables regression applied to a pooled data set of observations outside the band). This would yield a set of fitted residuals within and outside the band as well as estimates of the two variances, that could be inserted in (8) to obtain a value of the maximized likelihood function. When κ_1 and κ_2 are unknown, however, we can derive maximum likelihood estimates of them together with estimates of the other parameters by going through a similar procedure for a wide range of combinations of values of κ_1 and κ_2, at each point calculating the likelihood function and then finally choosing the combination of values of κ_1 and κ_2 that maximizes this function. This was the approach adopted in our empirical work: we carried out a twodimensional grid search for the threshold parameters, κ_1 and κ_2, and searched for a maximum of the likelihood function (8) by applying least squares to the data partitioned according to the values of κ_1 and κ_2 at each point in the grid search.

If it is believed that there are insufficient data points showing evidence of having crossed the lower threshold—such as a lack of negative deviations from CIP, for example—then it may not be possible to identify the asymmetric band-TAR model just outlined. In this case, however, we may still be able to estimate the upper threshold by estimating a single-threshold model of the form:

$$\delta_t = \delta_{t-1} + \varepsilon_{1t} \qquad \text{if } \delta_{t-1} < \kappa_1 \text{ and } \delta_{t-1} > \kappa_2, \tag{9a}$$

$$\delta_t - \kappa_1 = \rho(\delta_{t-1} - \kappa_1) + \varepsilon_{2t} \qquad \text{if } \delta_{t-1} \geq \kappa_1, \tag{9b}$$

The likelihood function then takes the form:

$$L(\rho, \theta, \sigma_1^2, \sigma_2^2, \kappa_1) = -\tfrac{1}{2} \sum_{\delta_{t-1} < \kappa_1} \left[\ln(\sigma_1^2) + \varepsilon_{1t}^2/\sigma_1^2\right] - \tfrac{1}{2} \sum_{\delta_{t-1} \geq \kappa_1} \left[\ln(\sigma_2^2) + \varepsilon_{2t}^2/\sigma_2^2\right] \tag{10}$$

and estimation may proceed as before.

Note that the maximum likelihood estimation procedure just described assumes that the disturbances are normally distributed. Allowing for other possible forms of distribution, such as Student's *t,* or for time-variation in the residual variance—for example, autoregressive conditional heteroskedasticity (ARCH) effects—would considerably complicate the algorithm just described, since it would not allow the use of simple least squares estimation to calculate the value of the likelihood function at each point of the grid search, and would require a more sophisticated method of maximum likelihood estimation. Moreover, this would also considerably complicate the Monte Carlo procedures discussed below. We therefore leave this on the agenda for future research.[9]

Having obtained maximum likelihood estimates of the parameters, we can then test the hypotheses of global mean reversion or zero bandwidth, $\kappa_1 = \kappa_2 = 0$, by estimating the restricted model and applying a likelihood ratio test. In this case, the restricted model is simply a first-order autoregression, AR(1).[10] The likelihood ratio statistic can be constructed as twice the difference between the value of the likelihood function for the AR(1) model and the maximized value of the likelihood function for the TAR model. As noted by Obstfeld and A. M. Taylor (1997), however, the thresholds are not identified under the null hypothesis and so the resulting likelihood ratio statistic may not have a standard χ^2 distribution. Accordingly, we estimated the empirical marginal significance level of the likelihood ratio statistics through Monte Carlo simulation (i.e., the parametric bootstrap). The steps involved in this were as follows: (i) estimate the restricted AR(1) model using the actual data; (ii) use the resulting parameter estimates to calibrate an artificial AR(1)

[9]As a check on the adequacy of the normality assumption, however, we performed tests for departures from normality on the estimated residuals in each of our estimated models using the Kolmogorov-Smirnov test statistic: in no case could the null hypothesis of normality be rejected.

[10]Note that in the restricted AR(1) model we also set the error variances equal to one another.

data generating process with gaussian errors and generate 5,000 artificial data sets equal in length to the actual data set plus 100, each with an initial value of zero; (iii) for each data set, discard the first 100 data points (to avoid initial value bias), estimate the TAR model and the AR(1) model, construct the likelihood ratio statistic and save it; (iv) take the resulting 5,000 values of the likelihood ratio statistic as the empirical distribution of the statistic under the null hypothesis.[11]

V. Data

The data are daily from December 16, 1996, to August 6, 1998, and cover one-, three-, and six-month maturity U.S. and Russian Treasury (GKO) bill rates, together with matching spot and forward ruble-dollar exchange rates. The U.S. data were taken from Datastream. The exchange rate data come from Datastream, Bankers Trust, and MICEX.[12] The Russian interest rate data were supplied by Bankers Trust and MFK Renaissance, Moscow. Our sample period extends until just before the mid-August 1998 ruble crisis.

VI. Empirical Results

Deviations from Standard, Spot-Forward CIP

Table 1 gives details of the asymmetric threshold autoregressive model (6) fitted to the deviations from CIP using the spot-one-month data (i.e., with j set to zero, or spot, and k set to one month). The maximum likelihood estimation procedure yields a band spanning the interval (–0.02, 0.68). Interestingly, the lower threshold is almost indistinguishable from zero, while the upper threshold is about two-thirds of one percent on an annualized basis. This indicates that arbitrage from rubles to dollars may occur even with small deviations, while arbitrage in the opposite direction—from dollars to rubles—requires a substantial profit margin. Within the band, deviations from CIP follow a random walk, while outside they switch to a first-order autoregressive process. The degree of persistence outside the band is very low, moreover, with an estimated first-order autoregressive coefficient of 0.313, indicating a half-life of deviations outside the band of less than one day, indicating strong arbitrage once the thresholds are crossed. The coefficient of determination indicates a high degree of explanatory power of the model—over 85 percent of the variation in δ_t is explained. Testing the hypothesis that both the upper and lower thresholds are equal to zero yields a large likelihood-ratio statistic with a marginal significance level of virtually zero, providing a strong rejection of these restrictions.

Table 2 gives details of the asymmetric threshold autoregressive model (6) fitted to the deviations from CIP using the spot-three-month data (i.e., with j set to

[11]These tests are quite similar to those discussed in Coakley and Fuertes (2001), where it is also shown that they have good power and size properties for the kinds of sample sizes employed in this paper.

[12]Further details on MICEX can be obtained from its website, *www.micex.com/english/index.html*. Information on the Russian financial markets can also be obtained from the Central Bank of Russia website, *http://www.cb.ru*.

Table 1. Results for the Threshold Autoregressive Model for Spot-One-Month CIP Deviations

Estimated band: $(\kappa_2, \kappa_1) = (-0.02, 0.68)$
Goodness of fit: $R^2 = 0.86$

i) *Within the band:*

$$\delta_t = \delta_{t-1} + \hat{\varepsilon}_{1t} \qquad \text{if } \delta_{t-1} \in (-0.02, 0.68)$$
$$\hat{\sigma}_1 = 0.2133$$

ii) *Outside the band:*

$$\delta_t - 0.68 = \underset{(0.088)}{0.313}(\delta_{t-1} - 0.68) + \hat{\varepsilon}_{2t} \qquad \text{if } \delta_{t-1} \geq 0.68$$
$$\delta_t + 0.02 = \underset{(0.088)}{0.313}(\delta_{t-1} + 0.02) + \hat{\varepsilon}_{2t} \qquad \text{if } \delta_{t-1} \leq -0.02$$
$$\hat{\sigma}_2 = 0.2237$$

Likelihood ratio test for setting both thresholds to zero: LR = 51.8975 [*0.00*]

Notes: Estimation results were obtained using maximum likelihood estimation as described in the text. R^2 denotes the coefficient of determination. Figures in parentheses below coefficient estimates denote estimated standard errors. *LR* denotes a likelihood ratio test statistic with empirical marginal significance level, calculated by Monte Carlo methods, given in square brackets.

Table 2. Results for the Threshold Autoregressive Model for Spot-Three-Month CIP Deviations

Estimated band: $(\kappa_2, \kappa_1) = (-0.01, 1.143)$
Goodness of fit: $R^2 = 0.89$

i) *Within the band:*

$$\delta_t = \delta_{t-1} + \hat{\varepsilon}_{1t} \qquad \text{if } \delta_{t-1} \in (-0.01, 1.143)$$
$$\hat{\sigma}_1 = 0.1817$$

ii) *Outside the band:*

$$\delta_t - 1.143 = \underset{(0.033)}{0.912}(\delta_{t-1} - 1.143) + \hat{\varepsilon}_{2t} \qquad \text{if } \delta_{t-1} \geq 1.143$$
$$\delta_t + 0.01 = \underset{(0.033)}{0.912}(\delta_{t-1} + 0.01) + \hat{\varepsilon}_{2t} \qquad \text{if } \delta_{t-1} \leq -0.01$$
$$\hat{\sigma}_2 = 0.3229$$

Likelihood ratio test for setting both thresholds to zero: LR = 51.22 [*0.00*]

Notes: Estimation results were obtained using maximum likelihood estimation as described in the text. R^2 denotes the coefficient of determination. Figures in parentheses below coefficient estimates denote estimated standard errors. *LR* denotes a likelihood ratio test statistic with marginal significance level, calculated by Monte Carlo methods, given in square brackets.

zero, and k set to three months). The estimated band for the threshold model is now a little wider—(–0.01, 1.143)—but this is due to raising the estimated upper threshold to over 1 percent on an annualized basis, while the lower threshold is again almost indistinguishable from zero. The other key difference from the spot-one-month results is the speed of mean reversion of deviations from CIP outside the band: the estimated first-order autoregressive coefficient is now 0.912, indicating a half-life of deviations outside the band of about seven days. The coefficient of determination again indicates a high degree of explanatory power of the model, with nearly 90 percent of the variation in δ_t explained. The hypothesis that both the upper and lower thresholds are equal to zero again yields a large likelihood-ratio statistic with a marginal significance level of virtually zero, providing a strong rejection of these restrictions.

Fitting the asymmetric threshold autoregressive model (6) to the deviations from CIP using the spot-six-month data, that is, with j set to zero, and k set to six months (Table 3), yields qualitatively similar results to those for the spot-three-month data, except that the lower threshold could not be identified because there were no negative deviations. The estimated upper threshold has now risen greatly, to over 2 percentage points on an annualized basis, while the estimated first-order autoregressive coefficient is again around the 0.9 level, indicating a speed of mean reversion of about 10 percent per day. The coefficient of determination has risen again, with nearly 95 percent of the variation in δ_t explained, and the hypothesis that both the upper and lower thresholds are equal to zero is again strongly rejected.

Deviations from Generalized CIP

Table 4 gives details of the asymmetric threshold autoregressive model (6) fitted to the deviations from CIP using the one-three-month data (i.e., with j set equal to

Table 3. Results for the Threshold Autoregressive Model for Spot-Six-Month CIP Deviations

Estimated upper threshold: $\kappa_1 = 2.152$

Goodness of fit: $R^2 = 0.94$

i) *Below the upper threshold:*

$$\delta_t = \delta_{t-1} + \hat{\varepsilon}_{1t} \qquad \text{if } \delta_{t-1} < 2.152$$

$$\hat{\sigma}_1 = 0.3134$$

ii) *Above the upper threshold:*

$$\delta_t - 2.152 = \underset{(0.038)}{0.901}(\delta_{t-1} - 2.152) + \hat{\varepsilon}_{2t} \qquad \text{if } \delta_{t-1} \geq 2.152$$

$$\hat{\sigma}_2 = 0.4168$$

Likelihood ratio test for setting upper threshold to zero: LR = 62.59 [*0.00*]

Notes: Estimation results were obtained using maximum likelihood estimation as described in the text. R^2 denotes the coefficient of determination. Figures in parentheses below coefficient estimates denote estimated standard errors. *LR* denotes a likelihood ratio test statistic with marginal significance level, calculated by Monte Carlo methods, given in square brackets.

Table 4. Results for the Threshold Autoregressive Model for One-Three-Month CIP Deviations

Estimated band: $(\kappa_2, \kappa_1) = (-0.03, 0.91)$
Goodness of fit: $R^2 = 0.85$

i) *Within the band:*

$$\delta_t = \delta_{t-1} + \hat{\varepsilon}_{1t} \qquad \text{if } \delta_{t-1} \in (-0.03, 0.91)$$
$$\hat{\sigma}_1 = 0.25883$$

iii) *Outside the band:*

$$\delta_t - 0.91 = \underset{(0.092)}{0.394}(\delta_{t-1} - 0.91) + \hat{\varepsilon}_{2t} \qquad \text{if } \delta_{t-1} \geq 0.91$$
$$\delta_1 + 0.03 = \underset{(0.092)}{0.394}(\delta_{t-1} + 0.03) + \hat{\varepsilon}_{2t} \qquad \text{if } \delta_{t-1} \leq -0.03$$
$$\hat{\sigma}_2 = 0.3750$$

Likelihood ratio test for setting both thresholds to zero: LR = 60.97 [*0.00*]

Notes: Estimation results were obtained using maximum likelihood estimation as described in the text. R^2 denotes the coefficient of determination. Figures in parentheses below coefficient estimates denote estimated standard errors. *LR* denotes a likelihood ratio test statistic with empirical marginal significance level, calculated by Monte Carlo methods, given in square brackets.

one month and k set equal to three months). The maximum likelihood estimation procedure yields a band spanning the interval (–0.03, 0.91). The lower threshold is again almost indistinguishable from zero, while the upper threshold is 1 percent. The degree of persistence outside the band is very low, as in the spot-one month case, with a first-order autoregressive coefficient of about 0.394, again indicating a half-life of deviations outside the band of less than one day, suggesting strong arbitrage once the thresholds are crossed. Over 85 percent of the variation in δ_t is explained and the hypothesis that both the upper and lower thresholds are equal to zero is strongly rejected on the basis of the likelihood-ratio statistic.

Table 5 gives details of the asymmetric band-TAR model fitted to the deviations from CIP using the three-six-month data over the sample period. The results are in many ways similar to those obtained for the spot-three-month data. The maximum likelihood estimation procedure yields a band spanning the interval (0.00, 1.223): the lower threshold is now estimated at exactly zero while the upper threshold estimate is again greater than 1 percent on an annualized basis. The point estimate of the first-order autoregressive coefficient is 0.904, indicating a half-life of deviations outside the band of over seven days, indicating that even when the no-arbitrage band is breached, arbitrage for this longer maturity may be relatively sluggish. The coefficient of determination indicates a very high degree of explanatory power of the model—about 95 percent of the variation in δ_t is explained, and the hypothesis that both the upper and lower thresholds are equal to zero is again easily rejected at standard significance levels, with a very large likelihood-ratio statistic and a marginal significance level of virtually zero.

The results of fitting the nonlinear model to the deviations from CIP using the one-six-month data over the sample period are given in Table 6. As in the case of

Table 5. Results for the Threshold Autoregressive Model for Three-Six-Month CIP Deviations

Estimated band: $(\kappa_2, \kappa_1) = (0.0, 1.223)$
Goodness of fit: $R^2 = 0.95$

i) *Within the band:*

$$\delta_t = \delta_{t-1} + \hat{\varepsilon}_{1t} \qquad \text{if } \delta_{t-1} \in (0.0, 1.223)$$

$$\hat{\sigma}_1 = 0.1915$$

iii) *Outside the band:*

$$\delta_t - 1.223 = \underset{(0.029)}{0.904}(\delta_{t-1} - 1.223) + \hat{\varepsilon}_{2t} \qquad \text{if } \delta_{t-1} \geq 1.223$$

$$\delta_t = \underset{(0.029)}{0.904}\delta_{t-1} + \hat{\varepsilon}_{2t} \qquad \text{if } \delta_{t-1} \leq 0.0$$

$$\hat{\sigma}_2 = 0.3229$$

Likelihood ratio test for setting both thresholds to zero: LR = 44.19 [*0.00*]

Notes: Estimation results were obtained using maximum likelihood estimation as described in the text. R^2 denotes the coefficient of determination. Figures in parentheses below coefficient estimates denote estimated standard errors. *LR* denotes a likelihood ratio test statistic with marginal significance level, calculated by Monte Carlo methods, given in square brackets.

spot-six-month arbitrage (Table 3), given that no observations of deviations from CIP for this data over the period fell below zero, we estimate only the upper threshold, in a single threshold-TAR model. The maximum likelihood estimation procedure yields a point estimate of the upper threshold of 2.922 percent, indicating that very large margins indeed—about 3 percent on an annualized basis—are necessary in order to induce arbitrage into rubles. Below this upper threshold,

Table 6. Results for the Threshold Autoregressive Model for One-Six-Month CIP Deviations

Estimated upper threshold: $\kappa_1 = 2.922$
Goodness of fit: $R^2 = 0.98$

i) *Below the upper threshold:*

$$\delta_t = \delta_{t-1} + \hat{\varepsilon}_{1t} \qquad \text{if } \delta_{t-1} < 2.922$$

$$\hat{\sigma}_1 = 0.3242$$

iii) *Above the upper threshold:*

$$\delta_t - 2.922 = \underset{(0.025)}{0.933}(\delta_{t-1} - 2.922) + \hat{\varepsilon}_{2t} \qquad \text{if } \delta_{t-1} \geq 2.922$$

$$\hat{\sigma}_2 = 0.4912$$

Likelihood ratio test for setting upper threshold to zero: LR = 78.59 [*0.00*]

Notes: Estimation results were obtained using maximum likelihood estimation as described in the text. R^2 denotes the coefficient of determination. Figures in parentheses below coefficient estimates denote estimated standard errors. *LR* denotes a likelihood ratio test statistic with marginal significance level, calculated by Monte Carlo methods, given in square brackets.

deviations from CIP follow a random walk, while above it they switch to a first-order autoregressive process. The degree of persistence above the threshold is again quite high, with a first-order autoregressive coefficient of about 0.933, indicating a half-life of deviations outside the band of nearly ten days, again indicating the comparative weakness of arbitrage even outside the no-arbitrage band when the maturity gap being arbitraged is large. A very good fit is indicated by the coefficient of determination, which shows that about 98 percent of the variation in δ_t is explained. Testing the hypothesis that the upper threshold is equal to zero yields a large likelihood-ratio statistic with a marginal significance level of virtually zero, providing a strong rejection of these restrictions.

Discussion

The econometric results reported in this paper for the behavior of deviations from CIP, both in its standard formulation and in a more general "forward-forward" formulation, are striking. In particular, our nonlinear econometric results reveal that the estimates of the no-arbitrage band for CIP deviations, as defined in equation (6), are highly asymmetric, indicating that only a very small deviation in favor of the dollar will activate arbitrage from the ruble to the dollar, but a very much larger deviation in favor of the ruble—in the longer maturities as high as 3 percent on an annualized basis—is required to activate arbitrage from the dollar to the ruble.

The size of the estimated band also increases as the maturity of the implied arbitrage contracts increases, in line with the "maturity effect" first noted by Taylor (1987, 1989).

The estimates of the speed of adjustment of deviations from CIP outside the no-arbitrage band also display a maturity effect, in that deviations outside the band disappear more slowly, the longer the maturity gap in the yield curve that is being arbitraged.

The finding of asymmetry in the no-arbitrage bandwidth is unlikely to indicate market inefficiency, since the lower edge of the estimated band (the minimum deviation needed to trigger arbitrage from rubles to dollars) is in four of the six cases estimated to be very close to, or indeed equal to, zero, and in the other two it cases could not be identified because there were no data points at which it would have been profitable to arbitrage from rubles to dollars. On the other hand, the fact that very large deviations are apparently needed to induce arbitrage from dollars to rubles indicates that arbitrageurs may have been factoring Russian default premiums into their decisions concerning the minimum deviation from CIP required for arbitrage to be deemed worthwhile. Indeed, our analysis suggests that these premiums may have been very substantial—on the order of between 70 and 300 basis points on an annualized basis. As discussed above, given the moratorium on debt repayment announced in August 1998, this behavior turned out to be justified.

It is well known that forward exchange rates in particular may embody foreign exchange risk premiums (Taylor, 1995), and yet CIP is generally thought to hold very closely among the major currencies (Taylor, 1992, 1995). A difference between studies involving major currencies and this study, however, is that CIP is generally thought of as being *riskless* arbitrage; the usual risk premium discussed in the literature relates to taking open forward positions in a currency and, as long as there is no risk in covered interest rate arbitrage, this will not affect the CIP condition. In the present context, however, where there may have been a very real perceived risk of default, covered arbitrage itself becomes risky and this risk may not be reflected in the forward rate in the usual way. The standard foreign exchange risk premium drives a wedge between the forward exchange rate and the expected future spot rate because speculators stop short of taking open forward positions to the point where the two would be driven to equality, because of the exchange rate risk of holding the open position. The risk of default may affect covered arbitrage in a similar way, by causing arbitrageurs to stop short of arbitraging from dollars to rubles to the point where CIP would hold exactly, because of the default risk of holding the Russian currency.

Note also that while Russian interest rates in this period may have themselves incorporated a default premium, making them higher than they would otherwise have been, there would still be deviations from CIP reflecting default premiums. The *level* of interest rates is in itself beside the point; what matters is that the expected profit from covered arbitrage has to outweigh the perceived risk of default and the complete loss of funds. If this perceived risk is virtually zero, as with CIP among the major currencies, then the expected profit will be driven close to zero. If the perceived risk is substantial, however, as appears to be the case for the Russian ruble during the period considered in this paper, then the deviation from CIP—at least as far as arbitrage *into* the risky currency is concerned—will not be driven to zero. Since there is no commitment of funds at the time the arbitrage is undertaken, the expected payoff from the arbitrage, assuming market efficiency, must be zero in equilibrium. If there is a debt moratorium, then the arbitrageur from dollars to rubles will, at the maturity date, have a maturing liability in dollars but a nonperforming ruble asset and hence will suffer a loss. The expected payoff of the arbitrage is therefore the probability of default times this loss plus one minus the probability of default times the profit from the arbitrage when there is no default. Since, as noted, the total expected payoff will be zero in equilibrium, the arbitrage must yield a positive profit in the absence of default but when there is a significant probability of default.[13] This is exactly what we appear to be observing in the research reported in this paper.

VII. Summary and Conclusion

The research reported in this paper represents, to the best of the authors' knowledge, the first empirical study of covered interest rate arbitrage between the recently established Russian Treasury bill (GKO) market and the U.S. Treasury bill market. Our methodology synthesizes the two main approaches that have been

[13]Note that the implied probability of default cannot be inferred from the data because the value of the ruble asset will not generally be zero even when there is a default, since there will in general still be a non-zero probability that it will be wholly or partly redeemed at some future point, so that its expected present value will also be non-zero. The deviation from CIP then represents (even assuming market efficiency and equilibrium) a single equation in two unknowns, the implied probability of default and the expected value of the ruble asset if a default is announced, so that neither of them is identifiable.

taken in the literature for examining deviations from covered interest rate parity (CIP)—that is, regression-based analysis and "neutral" or "no-arbitrage" band analysis—by employing nonlinear threshold autoregressive econometric modeling techniques. We obtained good estimates of the nonlinear band-threshold models using daily data on deviations from standard CIP (involving arbitrage between the spot and forward markets) as well as from a generalized form of CIP (involving arbitrage between different forward markets) for the period December 16, 1996, to August 6, 1998. The estimated models accord well with our priors in the sense that a no-arbitrage band is detected in each case, within which there appears to be little arbitrage (so that changes in the deviations are purely random) and outside of which arbitrage induces mean reversion of the deviation toward the band. Moreover, the bands are found to be strongly asymmetric so that only small deviations from CIP appear to be necessary to trigger arbitrage from rubles to dollars while much larger deviations—of the order of 70 to 300 basis points on an annualized basis—are necessary to trigger arbitrage into rubles. This suggests the presence of default premiums such that the expected profit from arbitrage from rubles to dollars has to be enough to outweigh the perceived risk of default and the complete loss of funds. Given that the perceived risk of default on U.S. Treasury bills was presumably very close to zero, this would suggest the presence of an asymmetry in the nature of deviations from CIP, such that only very small deviations are necessary in order to trigger arbitrage from the ruble to the dollar—which will keep deviations in this direction low—while very much larger deviations will be needed to trigger arbitrage in the other direction, that is, from dollars to rubles. Our finding of the existence of asymmetric no-arbitrage bands in the deviations from CIP between the GKO market and the U.S. Treasury bill market is exactly in accordance with such reasoning.

Moreover, the size of the no-arbitrage band appears to rise and the strength of arbitrage appears to fall as the length of the maturity gap in the yield curve being arbitraged increases, suggesting increasing reluctance of arbitrageurs to tie up funds for increasingly longer periods of time.[14] This to some extent echoes the maturity effect first observed by Taylor for the dollar against European currencies (Taylor, 1987, 1989).

This study has employed daily data in order to obtain enough data points for reliable estimation over a roughly two-year period. Over such a span of data it seems reasonable to suppose that the default premiums were roughly constant, which is implicit in our analysis. A possible avenue for future research that is suggested by our results, therefore, would be to extend the length of the data sample and to model the Russian default premiums themselves by making them functions of other variables, perhaps in a nonlinear fashion. This might then give interesting insights into the nature of default premiums. Although this promises to be a difficult exercise, the evidence reported in this paper suggests that important default premiums do indeed exist in these markets, and so it is an avenue worth exploring.

[14]See Taylor (1987, 1989, 1992) for a discussion of the maturity effect in covered interest arbitrage. This effect may plausibly be linked to Shleifer's (2000) view that short-horizon investors arbitraging against noise traders face the risk that the latter's misperception of prices may become more extreme before mean-reverting.

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