# BLACK MARKET AND OFFICIAL EXCHANGE RATES: LONG-RUN EQUILIBRIUM AND SHORT-RUN DYNAMICS

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## Abstract

This paper presents further empirical evidence on the relationship between black market and official exchange rates in six emerging economies (Iran, India, Indonesia, Korea, Pakistan, and Thailand). First, it applies both time series techniques and heterogeneous panel methods to test for the existence of a long-run relationship between these two types of exchange rates. Second, it tests formally the validity of the proportionality restriction implying a constant black-market premium. Third, it also analyses the short-run dynamic responses of both markets to shocks. Finally, it tries to shed some light on the determinants of the market premium. Evidence of slow reversion to the long-run equilibrium is found. Further, it appears that capital controls and expected currency devaluation are the two main factors affecting the size of the premium and determining the breakdown in the proportionality relationship.



JEL Code: C23, F31.

Keywords: black market and official exchange rates, panel cointegration, impulse response functions.



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## 1. Introduction

Black markets for foreign currency develop mainly as a result of government restrictions on capital outflows, which induce domestic residents to seek alternative sources of foreign currency. Its supply generally comes from the tourist industry, while demand originates mainly from residents travelling abroad on holiday or to study. Since demand for foreign currency normally exceeds supply, suppliers are able to charge a higher price than the official rate. The difference between the black market (or parallel) exchange rate and the official rate is known as the black-market premium.

As argued in Kiguel and O`Connell (1995), a significant spread between black market and official rate may be a signal of macroeconomics misalignments, and consequently central banks will often intervene in the official market to eliminate the spread – hence the importance of investigating whether there exists a long-run relationship between black market and official exchange rates. [1]The implications are extremely important not only for policy-makers but also for financial managers investing in emerging markets and managing the exchange rate risk. In addition to the long-run equilibrium, it is also interesting to analyse the short-run speed of adjustment of the two types of exchange rates in response to external shocks.

In this paper, we use both time series and heterogeneous panel cointegration tests to test for the existence of a long-run relationship between the black market and the official exchange rate in six emerging countries (i.e., Iran, India, Indonesia, Korea, Pakistan, and Thailand). Moreover, we test by means of a panel Wald test the validity of the unity (proportionality) restriction on the cointegrating coefficient implied by portfolio-balance models (see, e.g., Dornbusch et al, 1983). Further, we investigate the short-run dynamics by estimating impulse response functions using bootstrap methods, and try to shed some light on the possible determinants of the market premium.

The layout of the paper is the following. Section 2 outlines the econometric methodology. Section 3 contains a descriptive analysis of the data and the cointegration tests. Section 4 focuses on the determinants of the black market premium. Section 5 summarises the main findings and offers some concluding remarks.

## 2. Econometric Methodology

Let $s_{it}$ and $s^*_{it}$ be the log of the black market and of the official exchange rate, respectively, in country $i$. To take into account heterogeneity, we apply the McCoskey and Kao (1998) panel cointegration test by specifying the following DOLS (Dynamic OLS – see Stock and Watson, 1993) regression equation:

$$s_{it} = \alpha_i + \beta s^*_{it} \sum_{j=-k}^{k} \phi_j \Delta s^*_{i,t-1} + u_{it} \qquad (1)$$

i =1,…N, $k_i$ = leads and lags of $\Delta s^*_i$

[1] Note that many studies (e.g., Booth and Mustafa, 1991) also consider this crucial to establishing whether or not the black market processes information efficiently (i.e. whether investors use information contained in the black market rate to predict movements in the official rate), though more

The number of leads and lags in equation (1) is chosen with the Akaike criterion. In the presence of significant autocorrelation, even with high orders of leads and lags, one should instead employ the DGLS (Dynamic Generalised Least Squares – see McCoskey and Kao, 1998) estimation method, together with the Newey and West (1994) HAC (heteroscedasticity and autocorrelation consistent) covariance estimator, as we do below (see Section 3).

The method introduced by McCoskey and Kao (1998) involves a residual-based Lagrange Multiplier test for the null hypothesis of cointegration in panel data. This test is a panel version of the Harris and Inder (1994) cointegration test for time series. McCoskey and Kao (1998) show that the standardised version of the LM statistic is given by:

$$LM^* = \frac{[\sqrt{N}(LM - u_v)]}{\sigma_v} \Rightarrow N(0,1) \qquad (2)$$

where $u_v$ and $\sigma_v$ are obtained by Monte Carlo simulation methods (see Table 1, McCoskey and Kao, 1998).

We also analyse the short-run dynamic adjustment of exchange rates to external shocks by estimating impulse response functions Let $S_t = [s_t, s_t^*]'$ be a discrete time real valued vector stochastic process, and assume it follows the following Vector Autoregressive model of order $k$; $VAR(k)$:

$$S_t = \Pi_0 + \Pi_1 S_{t-1} + \Pi_2 S_{t-2} + ... + \Pi_k S_{t-k} + \mathrm{B}u_t \qquad t = 1,2... \qquad (3)$$

where $u_t$ is a $p \times 1$ vector of error terms having a martingale specification with covariance matrix $\Omega$.

In a more compact form (3) can be written as follows:

$$\Pi(L)S_t = \mathrm{B}u_t \qquad (4)$$

where $\Pi(L) \equiv \Pi_0 - \Pi_1 L - \Pi_2 L^2 - ... - \Pi_p L^p$, with $L$ being a lag operator.

A Vector Moving Average (VMA) representation of (3) is:

$$S_t = C(L)u_t \qquad (5)$$

where $C(L) = \Pi(L)^{-1}$; $C(L) = \mathrm{I} + C_1(L) + C_2 L^2 ...$, and $C(L)$ is a polynomial matrix generally of infinite order.

Let $\phi_{kl,i}$ be the response of the variable $k$ to a standard deviation shock on the variable $l$, $i$ periods ahead, and define as $\varphi = [\Pi_{1,} \Pi_2, ..., \Pi_p]'$, a vector that stacks the parameters of the system in (3). Impulse responses are obtained by inverting the polynomial (3), therefore $\phi_{kl,i}$ is a non-linear function of $\varphi'$ such that $\phi_{kl,i}(\varphi)$. Because of its non-linearity, in small samples $\phi_{kl,i}$ can be biased, with the bias increasing with the length of the horizon. For this reason, confidence intervals should also be constructed. Different methods have been suggested. In this paper we follow

recently it has been shown that the concepts of efficient markets and cointegration are independent of one another (see, e.g., Lence and Falk, 2005).

Kilian (1999) and adopt a residual-based resampling method (i.e. a semi-parametric one)[2].

**3. Descriptive Analysis and Cointegration Tests**

We use monthly data on black market and official exchange rates for six emerging market economies, namely Iran, India, Indonesia, Korea, Pakistan and Thailand, for the period 1973M1-1998M1. The former series are obtained from *Pick`s World Currency Yearbook* (various issues). In this respect, our study differs from earlier ones focusing on individual countries (see, e.g., Booth and Mustafa, 1991), or using annual data covering a shorter period (see, e.g., Bahmani-Oskooee et al, 2002)

Table 1 shows the black-market premium for the six exchange rates being analysed. As can be seen, it is sizeable in the case of Iran. This might suggest that, de facto, Iran has a different type of exchange rate regime from the other countries in the sample.

**Black-Market Premium (ABMP %) 1973-1998**

| Country | ABMP |
|---|---|
| Iran | 662.09 |
| India | 13.96 |
| Indonesia | 5.613 |
| Korea | 4.69 |
| Pakistan | 15.07 |
| Thailand | 0.14 |

***Table 1***

*Note: the ABMP has been calculated as* $(S/S^{*}-1)*100$ *dividing the black market exchange rate by the official exchange rate.*

As pointed out by Ghei and Kamin (1996), the level of the black market premium has decreased, on average, in most of the countries with black market exchange rates. Their analysis is based on annual data and ends in the early 1990s. Similar results are found at a monthly frequency for our set of countries, over a longer time period, including also the years 1994-1998 as one of the sub-samples (see Table 2).

| | **% Premium Over Different Subsamples** | | | | |
|---|---|---|---|---|---|
| **Country/Yrs** | 1973/1978 | 1979/1983 | 1984/1988 | 1989/1993 | 1994/1998 |
| **IRAN** | 3.7 | 263.3 | 876.8 | 21.3 | 165.4 |
| **INDIA** | 16.13 | 17.71 | 18.36 | 12.5 | 4.7 |
| **INDONESIA** | 3.5 | 3.9 | 9.8 | 8.7 | 2.5 |
| **KOREA** | 5.02 | 9.9 | 5.3 | 2.4 | 0.7 |
| **PAKISTAN** | 22.03 | 29.6 | 8.4 | 7.6 | 6.4 |
| **THAILAND** | 0.3 | 0.3 | 1.9 | 2.1 | 1.2 |

***Table 2***

As can be seen the size of the premium, on average, has generally decreased, in some cases substantially (e.g., in India, Korea, and Pakistan), which is prima facie evidence

[2] Note that, in order to deal with the non-stationarity of the variables, we estimate a Vector Error Correction model, save the residuals and use a re-sampling scheme on the residuals and equation (3) to generate bootstrap samples (see Li and Maddala, 1997).

that in these countries official and black market exchange rate markets are becoming integrated.

The sample period considered in Table 2 includes both periods when exchange controls are in place, and therefore parallel markets exist, and periods when markets are unified. By contrast, Table 3 reports the size of the average premium only for the latter case. Following Ghei and Kamin (1996), we consider markets to be unified if the absolute value of the difference between the black market and official market exchange rates is less than 3 per cent for at least twelve months. We identify only one such period in Iran and India, three in Indonesia and Thailand, and two in Korea. None are found in the case of Pakistan.

| | **% Premium In Periods With Unified Markets** | | |
|---|---|---|---|
| Country | | | |
| IRAN | 0.82 (1974M1-1976M8) | | |
| INDIA | 0.86 (1994M4-1995M6) | | |
| INDONESIA | 0.98 (1979M3-1980M10) | 1.05 (1989M6-1991M1) | 0.2 (1993M12-1997M6) |
| KOREA | 0.7 (1989M7-1990M7) | 0.81 (1994M6-1997 M9) | |
| PAKISTAN | N/A | | |
| THAILAND | 1.4 (1987M5-1988M9) | 0.9 (1991M4-1992M7) | 0.6 (1994M1-1997M4) |

***Table 3***

We then apply panel unit root tests to test for stationarity of the black market and official exchange rates. Taylor and Sarno (1998) propose a panel unit root test for the joint null hypothesis of nonstationarity The test is based on the Johansen maximum likelihood method for multivariate cointegration, where the null hypothesis is that at least one of the series in the panel has a unit root, $H_0$: rank $(\Pi) < N$. The null is rejected if and only if all the series in the panel are stationary processes, i.e. $H_1$: rank$(\Pi) = N$, where the matrix $\Pi$ denotes the long-run solution of the VAR system for N variables. This is equivalent to testing the null hypothesis that the smallest eigenvalue of $\Pi$, $\lambda_{min}$, is non-zero, which is done using Johansen's likelihood ratio statistic (JLR),

$$JLR = -T \ln(1 - \lambda_{\min}) \qquad (6)$$

Taylor and Sarno (1998) show that the *JLR* statistic has a limiting $\chi^2(1)$ distribution under the null hypothesis. The results of the JLR test are reported in Table 4, and confirm that both types of exchange rates are $I(1)$ processes, as usually found in the empirical literature.

**Taylor and Sarno (1998)'s *JLR* Panel Unit Root Test**

| | ***JLR* statistic** |
|---|---|
| $s_t$ | 0.110556 |
| $s_t^*$ | 0.123574 |

***Table 4***

*Note:* The 5% critical value is 3.9712 (Taylor and Sarno, 1998, Table 3).

Next, we test for cointegration between the two exchange rates[3]. Booth and Mustafa (1991) inter alia argue that the existence of cointegration is inconsistent with the efficient market hypothesis, which implies that past information on the exchange rate cannot be exploited to forecast future values: a feedback from one market to the other would constitute evidence of weak-form informational inefficiency in the black market. However, more recently it has been shown that market efficiency and cointegration are not incompatible (see, e.g., Lence and Falk, 2005).

**McCoskey and Kao (1998)'s Panel Cointegration Test**

| **Country** | **Leads/Lags** | ***AR(ρ)*** | ***Pr[Fa]*** | **Harris and Inder CointegrationTest** |
|---|---|---|---|---|
| Iran | 1 | 1 | 12.15 | 0.0787 |
| India | 2 | 1 | 10.39 | 0.1286 |
| Indonesia | 7 | 2 | 6.77 | 0.1506 |
| Korea | 8 | 2 | 15.58 | 0.2921 |
| Pakistan | 1 | 1 | 8.88 | 0.1313 |
| Thailand | 7 | 1 | 14.35 | 0.3689 |
| **McCoskey-Kao panel test ($LM^*$)** | | | | **0.56** |
| **CV-5%** | | | | **1.64** |

***Table 5***

*Note:* (a) The critical values for the Harris and Inder cointegration test (for time series with 1 regressor) are: 1%=0.5497; 5%=0.3202 (Harris and Inder, 1994, Table 1). (b) The $LM^*$ test is one-sided with a critical value of 1.64 (i.e. $LM^*>1.64$ implies rejection of the null hypothesis of cointegration).

(c) *Pr*[*Fa*] is the probability value of an $F_0$ version of the Breusch-Godfrey test for 9th order autocorrelation. *AR(ρ)* denotes the order of the autoregressive scheme employed in the model.

Table 5 reports two cointegration tests, specifically the McCoskey and Kao (1998) panel cointegration test, and the univariate cointegration tests due to Harris and Inder (1994), which is a time series version of it. Looking at the individual cointegration tests, we note that there is evidence of cointegration in all countries. The

[3] Note that, in a high inflation country, if the two exchange rates are unrelated to each other but move proportionately to the differential between the domestic and foreign price level, they will depreciate at a similar pace, which could lead to spurious cointegration. We have addressed this issue by deflating the black market and official exchange rates by the relative price levels, and by performing cointegration tests on the real exchange rates. The results, which are available on request, were qualitatively similar to those reported in Table 3 and 4, suggesting that the cointegration we find between nominal rates is not spurious.

panel cointegration test also confirms that there exists a long-run relationship between the two exchange rates.

Table 6 presents DGLS estimates of the cointegrating vector. As can be seen, all the coefficients are statistically significant (at least at the 10% confidence level), with the exception of the intercept for Thailand. The extremely small coefficient on the $\beta$ parameter for Iran might be seen as further confirmation that the Iranian exchange rate regime is different from the others. [4]

Portfolio-balance models are frequently used to provide an economic interpretation of the relationship between the two types of exchange rates (see, e.g., Dornbusch et al, 1983). They assert that asset market conditions determine the black market rate, and the current account affects it through the stock of black-market foreign currency. Consequently, there is a proportional equilibrium relationship between the two rates (with the black-market rate depreciating at same rate as the official rate in the long run), which implies that the long-run parameter $\beta$ should be equal to unity, resulting in a constant black-market premium. Consider, for instance, the case of an anticipated devaluation of the official rate. In the short run, the black market rate overshoots, as demand for black foreign currency increases but the supply of foreign currency is fixed, i.e. the premium rises as a result of an anticipated devaluation; however, the higher premium gradually leads to a current account surplus and an excess supply of foreign currency (since flow supply increases whilst flow demand decreases owing to a higher premium). This triggers off a decline in the premium, which gradually reverts to its long-run value, implying that in the long run the change in the black market rate is proportional to that in the official rate. In the Dornbusch et al, 1983 model, portfolio preferences are assumed to be constant. However, it is conceivable that they might shift, which would widen the premium. In such a case, the cointegrating coefficient would change over time and become different from the unity value implied by standard portfolio-balance models. To investigate this issue, we test the proportionality restriction by means of a Wald test (see Table 6).

**DGLS Estimates (Equation 1)**

| Country | $\alpha$ | $\beta$ | $\beta$=1 [$\chi^2(1)$] |
|---|---|---|---|
| Iran | 13.29 | 0.02 | |
| | [1.13] | [2.06] | 20.50 |
| India | 0.312 | 0.93 | |
| | [7.54] | [65.7] | 19.45 |
| Indonesia | -0.30 | 1.01 | |
| | [-1.74] | [40.57] | 0.85 |
| Korea | 1.16 | 0.83 | |
| | [2.12] | [9.84] | 4.19 |
| Pakistan | 0.463 | 0.89 | |
| | [3.62] | [20.28] | 6.91 |
| Thailand | -0.120 | 1.03 | |
| | [-0.87] | [23.37] | 0.76 |

[4] Two types of regimes are likely to characterise the countries analysed in this paper. In the first, widespread controls on a broad range of transactions may force many current account as well as capital account transactions onto the black market; then leakages between the markets (for example, as agents take foreign exchange acquired in the official market and resell it in the black market) might result in relatively strong linkages between the official and black market exchange rates. In the second (e.g., Iran), the official exchange rate is fixed and allowed to become so overvalued that few transactions take place at that rate, with most of the transactions taking place illicitly at the black market rate; in this situation, the official exchange rate might have a smaller influence on the black market rate.

| | |
|---|---|
| **PLR** | **52.66** |
| **CV-1%** | **16.81** |
| **CV-5%** | **12.59** |

***Table 6***

*Note:* Numbers in parentheses below the regression coefficients are t-values. $\chi^2(1)$ is the Wald test for the proportionality restriction, $H_0$:β=1 (5% CV=3.84). PLR is the corresponding panel likelihood ratio statistic that follows a $\chi^2(v)$ distribution with N(=6) d.f .

As in Cerrato and Sarantis (2003), we also extend this test to a panel context. The null hypothesis of a valid restriction is accepted in the case of Thailand and Indonesia, whilst it is strongly rejected in all the other cases. The panel test confirms that the proportionality restriction cannot be accepted for the panel under investigation. This could be the result of shifts in portfolio preferences, or perhaps of the imposition of foreign exchange and/or capital controls. By contrast, Bahmani-Oskooee et al (2002) report in their study that in the long run full adjustment takes place. However, these authors base their claim simply on point estimates (close to one) of the long-run slope coefficient, and unlike us do not carry out any statistical tests of the relevant restriction. [5] On the basis of these findings, we can conclude that black market and official exchange rates are linked in the long run, though the empirical evidence seems also to suggest that the black-market premium is unlikely to disappear even in the long run.

Finally, we investigate the short-run dynamic adjustment between the two types of exchange rates. Figures 1-6 show impulse responses to a unit standard deviation shock to either exchange rate. The empirical results appear to confirm the positive significant impact of the official exchange rate on the black market exchange rate. Furthermore, they are consistent with the sign of the $\beta$ -coefficient estimated by DGLS.

We also examine the response of each exchange rate to shocks to the same market. These show that both the official and the black market exchange rates are highly persistent processes[6]. However, in general, the speed of response to a unit standard deviation shock seems to be higher for the black market exchange rate compared to the official one.

On the whole, we observe the short-run overshooting predicted by portfolio-balance models. However, we also find that, even after forty months, the effects of shocks have not died away, which is inconsistent with the long-run full adjustment implied by this class of models.[7] It would appear, therefore, that the long-run black-market premium is not constant, or, at least, that deviations from the long-run equilibrium are long-lived, with long memory characterising these processes (see Granger and Joyeux, 1980).

[5] Some single-country studies test formally the long-run proportionality restriction, and cannot reject it (see, e.g., Kouretas and Zarangas, 2001).

[6] We tested for stationarity of the individual exchange rates and found that, with the exception of Korea, they were all non-stationary. This might explain their persistence which is shown by the impulse responses.

[7] This result is in line with Cerrato and Sarantis (2004), who report point estimates of half-lives deviations of the exchange rate from PPP ranging, for some of these countries, between 0.22 years and infinity, while the upper bounds of the confidence intervals were in all cases infinite.

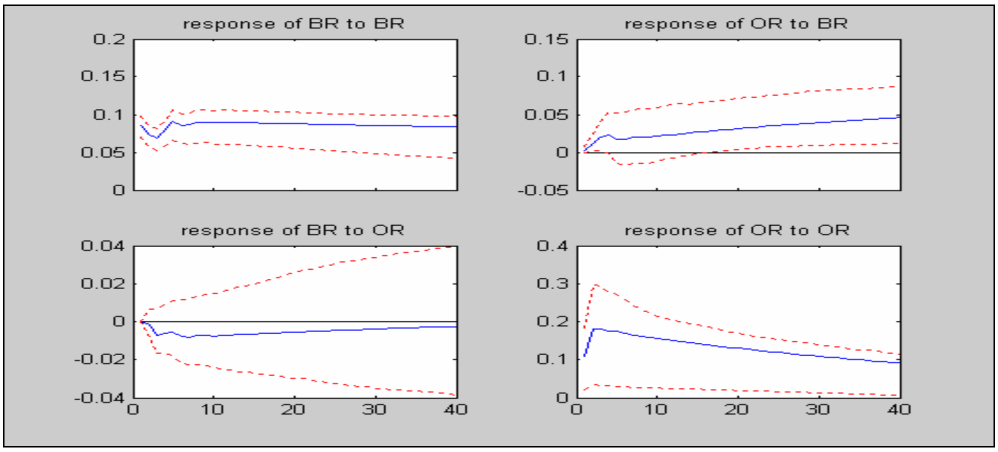


***Figure 1***: *Iran; BR and OR are respectively black market and official market exchange rates.*

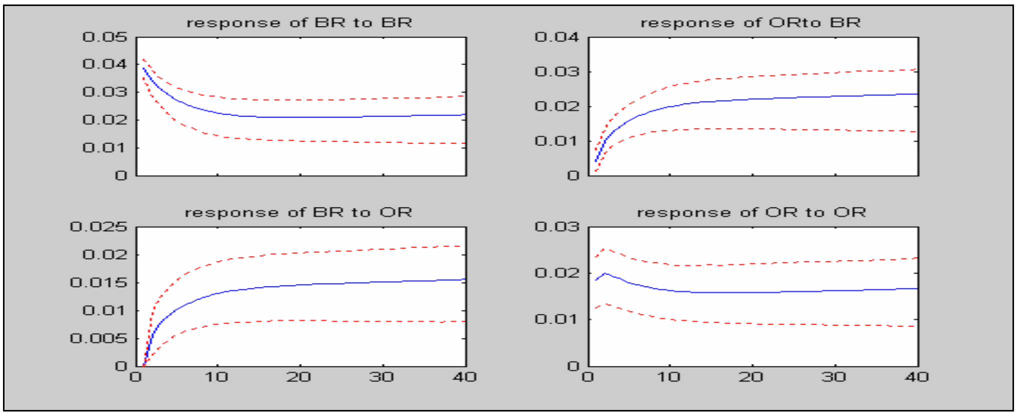


***Figure 2***: *India; BR and OR are respectively black market and official market exchange rates.*

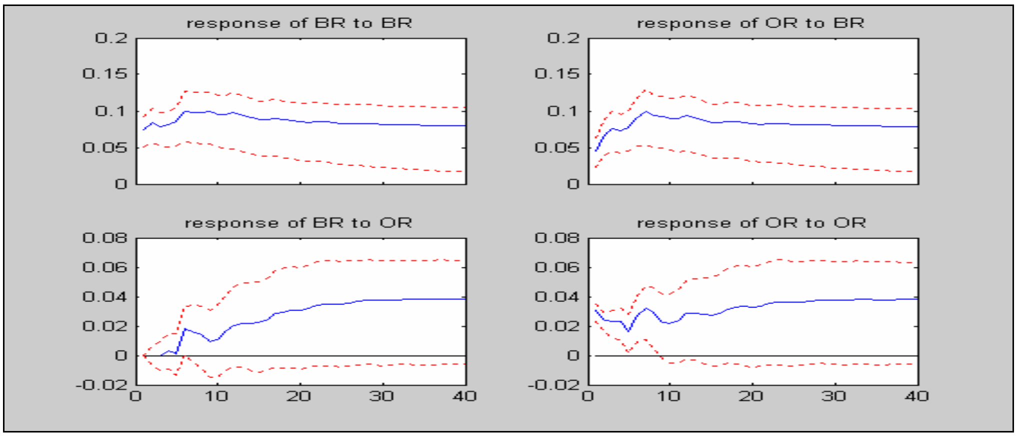


***Figure 3***: *Indonesia; BR and OR are respectively black market and official market exchange rates.*

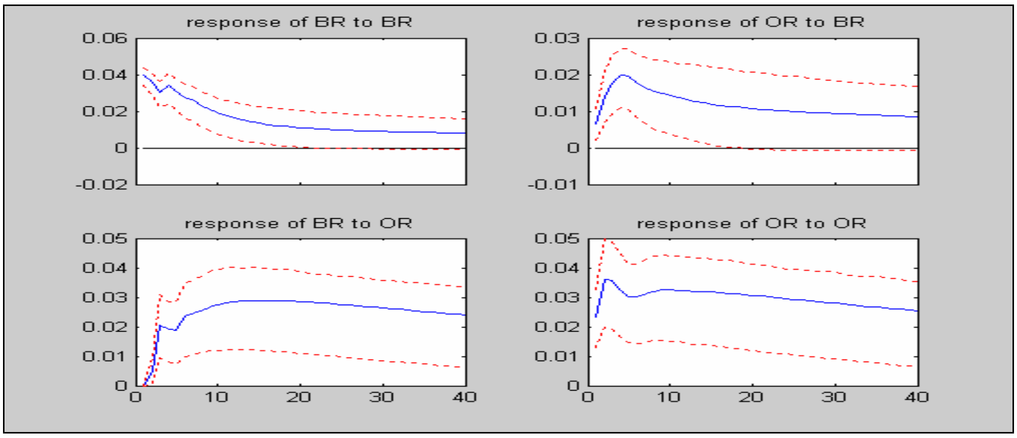


***Figure 4***: *Korea; BR and OR are respectively black market and official market exchange rates.*

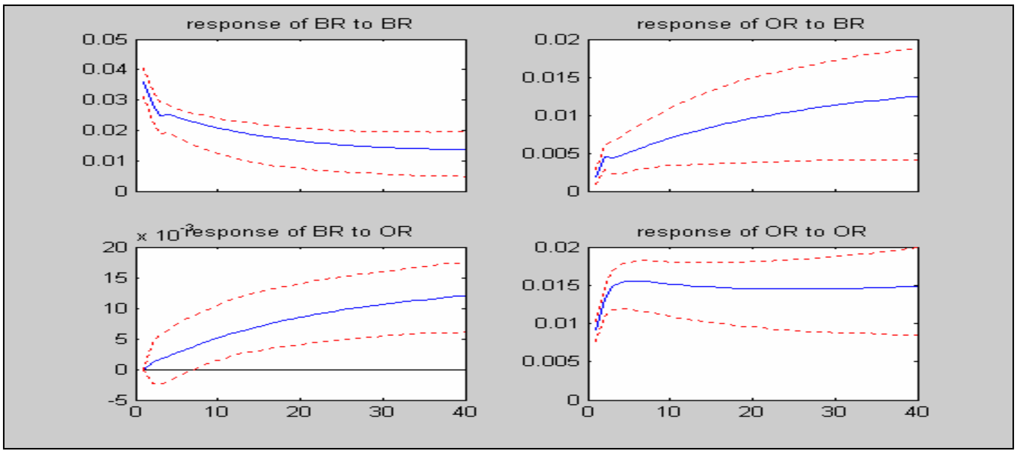


***Figure 5***: *Pakistan; BR and OR are respectively black market and official market exchange rates.*

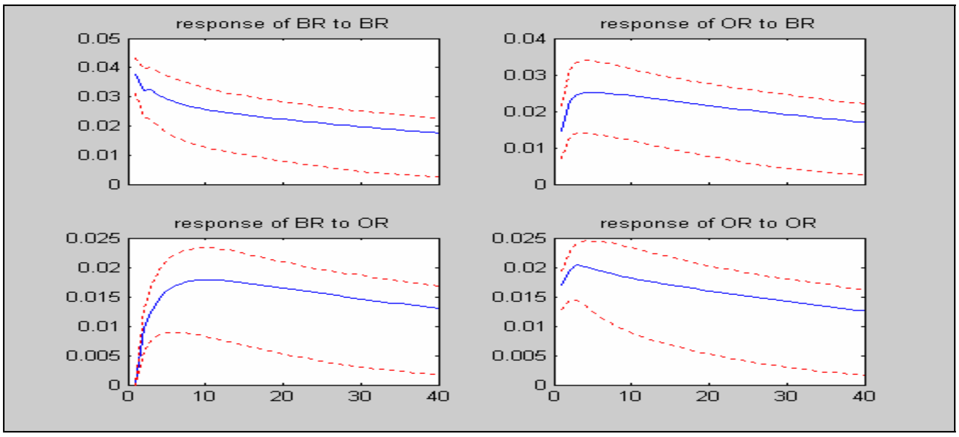


***Figure 6***: *Thailand; BR and OR are respectively black market and official market exchange rates.*

## 4. Inflation, Capital Controls, Expected Currency Devaluation and the Black Market Premium

The analysis of the previous section has shown that, although official and black market exchange rates are cointegrated, the proportionality restriction implied by the

Dornbusch et al, 1983 model is unlikely to hold. In this section we try to shed further light on this issue and to identify some reasons why this might be the case.

Following Ghei and Kamin (1996), we calculate some simple ratios of black market to official real exchange rate over different periods (i.e. with unified and not unified markets). The equilibrium real exchange rate (RER) is defined as the real official exchange rate averaged over the entire sample. We first compare the ratio of the black market exchange rate to RER over the full sample. The mean, reported at the bottom of Table 7, suggests that the black market exchange rate appreciated more than the official one over that period. This might account for the existence of a significant premium, in line with results presented in Table 1. However, this sample includes both periods when capital controls were in place and periods when they were not. Next, we focus only on periods of unified markets (i.e. when capital controls were not in place). In this case, the black market rate is found to be much closer to the long-run equilibrium exchange rate, and the market premium is now very small. Therefore, it appears that the market premium can change over time owing to factors such as the imposition of capital controls.

**Table 7: Main Ratios Indicators**

| Country/Ratios | eb/e | eb/eu |
|---|---|---|
| IRAN | 3.6 | 1.02 |
| INDIA | 1.13 | 1 |
| INDONESIA | 1.06 | 1 |
| KOREA | 1.04 | 1 |
| PAKISTAN | 1.13 | N/A |
| THAILAND | 1 | 1 |
| Mean | 1.493333 | 1.004 |

Note: eb is the black market real exchange rate for the entire period.
e is the official real exchange rate for the entire period.
eu is the official real exchange rate during periods of unified markets.

These findings might be useful to interpret the panel results on the proportionality restriction reported in the previous section: they highlight the importance of capital controls in explaining the risk premium and the appreciation of black market exchange rates relative to their long-run equilibrium level.

We also examine other factors that might affect the premium and lead to a breakdown in the proportionality relationship implied by the model due to Dornbusch et al (1983). Specifically, we model the risk premium (P) as a function of inflation (I), expected devaluation (E), and dummy variables to account for capital controls (D). As a measure of expected depreciation we use the 10-year bond yield differential between the domestic country and the US. The inflation rate is constructed using the consumer price index (CPI), and a capital controls dummy variable, which takes the value of one when markets are not unified and zero otherwise, is also included. These additional data are taken from the IMF's International Financial Statistics. The adopted model specification allows for k lags (up to eleven) of inflation and/or expected depreciation, but we only report estimates (and standard errors) for the last selected lag in the regression, as in the following equation:

$$P_t = a_0 + a_1 I_t + a_2 E_t + a_3 I_{t-k} + a_4 E_{t-k} + a_5 D_t + u_t \quad (7)$$

Table 8 presents the results. [8]

**Table 8: Regression Estimates**

| | $a_0$ | $a_1$ | $a_2$ | $a_3$ | $a_4$ | $a_5$ | Adj.R-squared |
|---|---|---|---|---|---|---|---|
| India | 12.88* | 0.12 | 0.83* | -0.12* | 1.12* | 1.4 | 0.4 |
| | [1.30] | [0.9] | [0.519] | [0.04] | [0.49] | [1.30] | |
| Korea | 0.24 | 0.33 | 0.25 | -0.11* | 0.3* | 2.68* | 0.3 |
| | [0.30] | [1.1] | [0.6] | [0.07] | [0.09] | [0.87] | |
| Thailand | 1.05* | 0.5 | -0.22* | 0.05* | 1.1* | -1.29* | 0.35 |
| | [0.025] | [0.87] | [0.08] | [0.025] | [0.7] | [0.39] | |

Note: numbers in parentheses are standard errors.
The asterisk indicates that variable is significant at the 10% level

It can be seen that the lags of inflation and expected devaluation are all significant, with the latter having the correct sign (i.e. positive, indicating that an expected devaluation increases the premium). A rise in inflation is found to lower the premium in the case of India and Korea, whilst the opposite is true in the case of Thailand. The dummy variable is significant only for Korea and Thailand, and with the expected sign only in the latter case.

## 5. Conclusions

This paper provides further empirical evidence on the relationship between black market and official exchange rates in six emerging economies (Iran, India, Indonesia, Korea, Pakistan, and Thailand). First, it applies both time series techniques and heterogeneous panel methods to test for the existence of a long-run relation between these two types of exchange rates. Second, it tests formally the validity of the proportionality restriction implying a constant black-market premium (unlike other studies, such as Bahmani-Oskooee et al, 2002, only reporting point estimates). Third, in addition to the long-run equilibrium, it also analyses the short-run dynamic responses of both markets to shocks. Finally, it investigates the determinant of the market premium. Our empirical analysis suggests that black market and official rates are linked in the long run. However, unlike other authors (e.g., Kouretas and Zarangas, 2001), we also find that the proportionality restriction, which is an essential feature of portfolio-balance models (see Dornbusch et al, 1983), is rejected, indicating that the adjustment towards equilibrium in response to short-run shocks is incomplete. Partial reversion to the long-run equilibrium (or possibly the presence of long-

[8] Owing to limited availability of data on bond yields we estimate this regression for fewer countries and over a shorter sample, specifically for India (1973M10-1990M12), Korea (1973M5-1988M12), and Thailand (1979M12-1988M12).

memory features) is confirmed by the short-run analysis, showing that the initial overshooting does not totally fade away. Further, as in Ghei and Kamin (1996), capital controls and expected currency devaluation are found to have a positive impact on the size of the premium, whilst there is weaker evidence of an impact of inflation. Therefore, it appears that the first two factors might be responsible for the breakdown in the proportionality relationship implied by portolio-balance models such as the one of Dornbusch et al (1983).

These results have important implication for both fund managers and policy-makers. The existence of international arbitrage opportunities suggests that the former can reduce exchange rate risk knowing that fluctuations in the black market rate signal corresponding adjustments in the official rate However, the exchange rate risk cannot be eliminated altogether, as full adjustment does not take place (or is extremely slow), and a widening gap between the two exchange rates is only partially closed (at least over a relevant investment horizon). Permanent (or long-lived) deviations from equilibrium also imply that monetary authorities can effectively pursue their policy objectives by imposing foreign exchange or direct controls (see Kiguel and O'Connell, 1995), not having to adjust the official rate to the market-determined parallel one in order to close the gap between the two (at least for a considerable time) – a crucial policy implication. This is in stark contrast to the study of Bahmani-Oskooee et al (2002), where proportionality is found though not formally tested.

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