Uncovered interest parity and policy behavior: new evidence

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Abstract

McCallum [J. Monet. Econ. 33 (1994) 105–132] introduces policy behavior to resolve previous empirical rejections of the uncovered interest parity (UIP) theory. In this note, we reexamine his policy behavior argument. First, we extend the data set used by McCallum to include the recent 8 years, and contrary to the analysis provided by McCallum, we make a thorough econometric analysis of his UIP specification. It is shown that in most cases his theory is supported by the data as well as it passes conventional econometric tests. We then take a closer look at his policy behavioral relationship, but unfortunately it turns out that this specification is inconsistent with the UIP specification suggested by McCallum. © 2000 Elsevier Science S.A. All rights reserved.

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JEL classification: F31; E58

1. Introduction

The uncovered interest parity (UIP) theory states that differences between interest rates across countries can be explained by expected changes in currencies. An expected depreciation of Deutsche Mark against the US Dollar, say, raises German interest rates compared to US interest rates in order to compensate American investors for the expected currency loss.

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; Mark and Wu, 1998), or that time-varying risk premia are present (see Domowitz and Hakkio, 1985; Nieuwland et al., 1998), respectively. 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\(^1\). 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.

In this note we analyze this hypothesis in further detail. First, we reestimate the UIP relationship for the extended sample period 1978.01m to 1999.03m and to a great extent we confirm the results obtained by McCallum for the period 1978.01m to 1990.7m. One criticism, which can be raised towards the analysis in McCallum (1994a), is, however, that McCallum did not provide a sound econometric analysis of his empirical results. In order to verify his UIP theory empirically, we provide a number of conventional econometric tests, and surprisingly it turns out that the UIP specifications estimated pass these tests, thereby lending support to the theory. To provide a final test of his UIP theory, we then estimate the policy reaction function suggested by McCallum in order to compare the estimates of this relationship and the estimates obtained from the UIP specification. Unfortunately, this test rejects the UIP theory based on policy behavior.

2. UIP and policy behavior

The UIP theory can be tested empirically by regressing the expected change in the spot exchange rate on the forward premium as:

$$E_t s_{t+1} - s_t = f_t - s_t - \xi_t$$

(1)

where $s_t$ is the natural log of the spot exchange rate, $f_t$ is the natural log of the forward rate, and $E_t s_{t+1}$ is the rational expectation at time $t$ of the log spot exchange rate to prevail at time $t + 1$ based on the information set $I_t$. The forward premium $f_t - s_t$ is assumed to equal the interest rate differential according to covered interest parity, i.e. $(f_t - s_t) = (R_t - R_t^*)$, where $R_t$ and $R_t^*$ are the nominal interest rates in the home and the foreign countries, respectively. Finally, $\xi_t$ reflects the expectational error as well as other influences that keep $s_t = E_t s_{t+1} - (f_t - s_t)$ from holding exactly.

Under the rational expectation hypothesis, McCallum derives the following regression equation:

$$s_{t+1} - s_t = \alpha + \beta (f_t - s_t) + \xi_{t+1}$$

(2)

where $\beta = 1$ is required to confirm the UIP theory, and $\alpha$ is a measure of a constant risk premium. In Table 1 we present the estimation results obtained from Eq. (2) using spot and 30-day forward exchange rates collected from the Bank for International Settlements (BIS).

The estimates in Table 1 are almost similar to the results found by McCallum (1994a), besides from the estimate of $\beta$ for $$/DM, which is insignificant here. However, in all cases the estimate of $\beta$ is significantly different from the expected value of $+1$, thereby rejecting the traditional UIP theory.

\(^1\)Kugler (2000) generalizes this result by making the policy reaction function dependent on the term structure spread.
Table 1
OLS estimation results obtained from Eq. (2) (sample period: January 1978–March 1999)

<table>
<thead>
<tr>
<th></th>
<th>$/DM</th>
<th>$/£</th>
<th>$/Yen</th>
</tr>
</thead>
<tbody>
<tr>
<td>$a$</td>
<td>0.0025</td>
<td>0.0023</td>
<td>0.0031</td>
</tr>
<tr>
<td></td>
<td>(0.0025)</td>
<td>(0.0023)</td>
<td>(0.0031)</td>
</tr>
<tr>
<td>$\beta$</td>
<td>0.5140</td>
<td>2.6662*</td>
<td>2.6940*</td>
</tr>
<tr>
<td></td>
<td>(0.9728)</td>
<td>(0.9728)</td>
<td>(0.9728)</td>
</tr>
<tr>
<td>$R^2$</td>
<td>0.002</td>
<td>0.034</td>
<td>0.036</td>
</tr>
<tr>
<td>S.E.</td>
<td>0.0340</td>
<td>0.0325</td>
<td>0.0359</td>
</tr>
<tr>
<td>DW</td>
<td>1.98</td>
<td>1.95</td>
<td>1.94</td>
</tr>
</tbody>
</table>

* Indicates significance at the 5% level. White heteroskedasticity-consistent standard errors in parentheses.

To explain this phenomenon McCallum argues that if the policy maker is interested in smoothing interest rates and uses interest rates to keep exchange rates stable, the policy reaction function would look like:

\[(R_t - R^*_t) = F(R_{t-1} - R^*_{t-1}) + \lambda (s_t - s_{t-1}) + \xi_t\]  

(3)

where $\xi_t$ is a white noise error term, and $0 < \sigma \leq 1$, $\lambda > 0$ by assumption.

McCallum argues that in order to make sure that the forward premium does not become pure white noise, the disturbance term $\xi_{t+1}$ in Eq. (2) cannot be white noise, and instead he suggests an AR(1) process given by:

\[\xi_t = \rho \xi_{t-1} + u_t, \quad |\rho| < 1.0\]  

(4)

Combining Eqs. (1), (3) and (4) using covered interest parity and assuming rational expectations, McCallum then derives the following reduced form equation for the exchange rate:

\[s_t - s_{t-1} = \alpha + \frac{\rho - \sigma}{\lambda} (f_t - s_{t-1}) - \frac{1}{\lambda} \xi_t + \frac{1}{\lambda + \sigma - \rho} u_t\]  

(5)

and on this basis McCallum concludes that Eq. (5) may be consistent with the UIP relationship in Eq. (2). In fact, McCallum suggests that $\sigma$ is close to 1, $\lambda$ is close to 0.2, and for values of $\rho \ll 1$, a negative parameter to $(f_t - s_t)$, in accordance with the empirical findings in Table 1 will appear.

To verify this result empirically, one has to test the estimated regressions in Table 1 econometrically. The test statistics ($R^2$, S.E. and DW) provided by McCallum and presented in Table 1 are not sufficient to validate the UIP theory empirically. Section 3, therefore, provides a number of more advanced test statistics. Furthermore, to confirm the $\beta$ estimates in Table 1, the estimates of $\sigma$ and $\lambda$ must be consistent with these $\beta$ estimates. As a final test of Eq. (5), we then estimate Eq. (3) in Section 4.

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1. In McCallum (1994b) and Kugler (1997) the policy reaction function is made dependent on the term structure spread, and Hsu and Kugler (1997) show empirically that the term spread has a significant influence on US short rates after 1987.
Table 2
Phillips–Perron unit roots tests* (sample period: January 1978–March 1999)

<table>
<thead>
<tr>
<th></th>
<th>$/DM</th>
<th>$/£</th>
<th>$/Yen</th>
</tr>
</thead>
<tbody>
<tr>
<td>$s_t - s_{t-1}$</td>
<td>-15.80</td>
<td>-15.08</td>
<td>-14.97</td>
</tr>
<tr>
<td>$f_t - s_t$</td>
<td>-2.12</td>
<td>-2.61</td>
<td>-3.22*</td>
</tr>
</tbody>
</table>

* The 5% critical value is -1.94.
* For $/Yen the critical 5% value is -2.87.

3. Econometric evidence

To analyze the validity of the regressions in Table 1, we first test whether the variables $s_t - s_{t-1}$ and $f_{t-1} - s_{t-1}$ are stationary. In Table 2 we present the results of the Phillips–Perron unit roots test. From Table 2 we infer that all variables seem to be stationary at the 5% level\(^1\) lending support to the OLS regressions in Table 1. To confirm this conclusion we will in addition provide a number of misspecification tests. In Table 3 we present the Ljung–Box $Q$-statistic for 12th-order serial correlation, the Jarque–Bera normality test, the ARCH test for 12th-order conditional heteroskedasticity, and two Chow tests; one testing for a structural break in October 1982, where monetary policy was changed significantly in the US and the United Kingdom\(^4\), and one testing for a structural break in July 1993, where Chairman Greenspan indicated that the Federal Reserve would no longer use monetary targets as guidelines for its monetary policy.

Table 3 indicates that the regressions in Table 1 pass all tests performed besides the normality test, which is supported only for $/DM$ at the 1% level. We can, therefore, conclude that these relations seem to be well specified econometrically being fairly stable without exhibiting serial correlation or heteroskedasticity\(^5\).

Table 3
Misspecification test probabilities (sample period: January 1978–March 1999)

<table>
<thead>
<tr>
<th></th>
<th>$/DM</th>
<th>$/£</th>
<th>$/Yen</th>
</tr>
</thead>
<tbody>
<tr>
<td>$Q(12)$</td>
<td>0.634</td>
<td>0.714</td>
<td>0.507</td>
</tr>
<tr>
<td>Jarque–Bera</td>
<td>0.018</td>
<td>0.000</td>
<td>0.000</td>
</tr>
<tr>
<td>ARCH(12)</td>
<td>0.831</td>
<td>0.127</td>
<td>0.743</td>
</tr>
<tr>
<td>Chow(_{10.82})</td>
<td>0.777</td>
<td>0.580</td>
<td>0.804</td>
</tr>
<tr>
<td>Chow(_{07.93})</td>
<td>0.768</td>
<td>0.265</td>
<td>0.687</td>
</tr>
</tbody>
</table>

\(^1\)Concerning stationarity the $f_t - s_t$ variable for the $/Yen$ is critical. Therefore, for this variable stationarity has been tested by an Augmented Phillips–Perron test including an intercept term.

\(^4\)In October 1982 monetary policy was changed from controlling money to controlling interest rates in the US and the United Kingdom.

\(^5\)This evidence does not lend support to McCallum’s prediction that $\xi_t$ follows an AR(1) process.
4. The policy reaction function

To reconcile the $\beta$ estimates in Table 1 with the underlying structural parameters of the policy reaction function given by Eq. (3), a final test of McCallum’s policy behavioral UIP theory would be to estimate these structural parameters. One would imagine that if the policy behavior argument put forward by McCallum is the solution to the usual rejections of the traditional UIP theory, Eqs. (2) and (5) would be subject to the Lucas critique. However, in the previous section the UIP relationships were found to be rather stable and independent of major changes in monetary policy during the period analyzed, i.e. they escape the Lucas critique. Obviously, this evidence is inconsistent with the theoretical argument provided by McCallum (1994a).

Estimating Eq. (3) two problems arise. One is that for all three currencies we find that a structural break can be identified in October 1982, where monetary policy was changed significantly in the US and the United Kingdom. This evidence indicates that the UIP specifications in Table 1 may be incorrect, since no structural break can be identified (see Table 3).

The other problem is due to severe heteroskedasticity in the residuals for all three currencies. Although the OLS estimates provide appropriate inference, as long as we use White’s heteroskedasticity-consistent standard errors, we will try to correct for heteroskedasticity when estimating Eq. (3). Using different GARCH models we are in fact able to obtain reasonable specifications for the policy reaction function. However, it is worth mentioning that the structural parameters and their standard errors are almost identical to the OLS estimates, and therefore the ARCH parameters are not reported.

Again we follow McCallum, assuming that covered interest parity holds, and Table 4 presents the results.

<table>
<thead>
<tr>
<th></th>
<th>$/$DM GARCH(2,2)</th>
<th>$/$£ E-GARCH</th>
<th>$/$Yen E-GARCH</th>
</tr>
</thead>
<tbody>
<tr>
<td>$\alpha$</td>
<td>0.000</td>
<td>-0.000</td>
<td>-0.000*</td>
</tr>
<tr>
<td></td>
<td>(0.000)</td>
<td>(0.000)</td>
<td>(0.000)</td>
</tr>
<tr>
<td>$\sigma$</td>
<td>0.981*</td>
<td>0.960*</td>
<td>0.968*</td>
</tr>
<tr>
<td></td>
<td>(0.014)</td>
<td>(0.027)</td>
<td>(0.013)</td>
</tr>
<tr>
<td>$\lambda$</td>
<td>-0.003*</td>
<td>-0.002**</td>
<td>-0.001</td>
</tr>
<tr>
<td></td>
<td>(0.001)</td>
<td>(0.001)</td>
<td>(0.001)</td>
</tr>
<tr>
<td>$R^2$</td>
<td>0.899</td>
<td>0.835</td>
<td>0.879</td>
</tr>
<tr>
<td>Q(12)</td>
<td>0.502</td>
<td>0.380</td>
<td>0.620</td>
</tr>
<tr>
<td>Jarque–Bera</td>
<td>0.000</td>
<td>0.000</td>
<td>0.275</td>
</tr>
<tr>
<td>ARCH(12)</td>
<td>0.996</td>
<td>0.685</td>
<td>0.447</td>
</tr>
<tr>
<td>$\beta = -\sigma/\lambda$</td>
<td>0.001</td>
<td>0.067</td>
<td>0.172</td>
</tr>
</tbody>
</table>

*, ** Indicate significance at the 5% and the 10% levels, respectively.

6The 5% probability is 0.000, 0.000 and 0.028 for $/$DM, $/$£ and $/$Yen, respectively.
7These results can be obtained from the author upon request.
8In the case of $/$Yen, a MA(12) term has been included in order to correct for serial correlation.
presents the estimates of the underlying structural parameters in the policy reaction function given by Eq. (3).

The maximum likelihood estimates are obtained using the Marquardt algorithm in E-Views, and we infer that a GARCH (2,2) process performs best for the $/DM, whereas for $/£ and $/Yen an E-GARCH process is most appropriate.

From Table 4 we infer that for each currency the policy reaction function seems to be well specified. The goodness-of-fit is high and the residuals are not exposed to serial correlation or any remaining heteroskedasticity. Concerning the structural parameters, we find that five of six are significantly different from zero, and in particular we notice that the $ and $ parameters are almost identical across the three currencies analyzed. $ is close to 1 and $ is negative and very small although significantly different from zero. Table 4 also presents a Wald test of McCallum’s hypothesis that $ = $/ in Eq. (5), but unfortunately we see that this hypothesis is rejected in all cases, thereby rejecting the policy behavioral UIP theory suggested by McCallum. McCallum (1994a) predicted correctly that $ would be close to 1, but his intuition that $ = 0.20 cannot be confirmed for any of the currencies analyzed. In fact, $ is found to be negative, which is inconsistent with Eq. (3). This conclusion holds irrespective of the estimate of $ as long as $ < 1, which McCallum assumes.

5. Conclusions

By introducing a policy reaction function, McCallum (1994a) determines a UIP relationship, which seems to be consistent with most empirical findings, and thereby McCallum has provided a promising solution to previous rejections of UIP. However, McCallum (1994a) did not provide an explicit estimate of his policy reaction function. In this note we have analyzed the modified UIP theory suggested by McCallum (1994a) both from an econometric and an economic point of view. Econometrically, we find that $/DM, $/£ and $/Yen for the period 1978.01 to 1999.03 behave amazingly well according to the modified UIP theory developed by McCallum. Unfortunately, it turns out that when we estimate the policy reaction function, we find a well-specified relationship, whose structural parameters are inconsistent with the UIP relationships estimated.

Acknowledgements

Helpful comments from an anonymous referee are gratefully acknowledged.

References


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9 A structural break can be identified in October 1982, but regressions for the pre and post October 1982 periods do not alter this conclusion. In none of these sub-periods $ becomes significantly positive.

10 If one includes an AR(1) process in Eq. (2) (Table 1), the $ parameter becomes insignificant for all three countries.


