



# Inflation-targeting and real interest rate parity: A bias correction approach



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## ABSTRACT

This paper investigates whether inflation-targeting influences real interest rate parity (RIP) by a bias correction approach under cross-sectional dependence. The recursive mean adjustment (RMA) method proposed by So and Shin (1999) and Shin and So (2001) is employed to correct the downward bias in the panel unit root tests and in the half-life estimates of real interest rate differentials for traded and non-traded goods. The empirical findings differ depending on whether we apply the RMA. More importantly, the empirical results show that as more homogeneous economies become involved in terms of inflation-targeting regime, stronger mean reversion and much a tighter confidence interval are present. Thus, inflation-targeting plays an important role in providing favorable evidence for long-run RIP.

## 1. Introduction

The present paper examines whether inflation-targeting influences real interest rate parity (RIP) by a bias correction approach under cross-sectional dependence. RIP comprises uncovered interest parity (UIP) and purchasing power parity (PPP), which together imply the equalization of real rates of return in foreign exchange markets. Indeed, the assumption of the equality of real interest rates across countries characterized by a high degree of capital mobility together with high levels of technology diffusion served as an important premise in early monetary approaches to exchange rate determination.<sup>1</sup> RIP has also been used to investigate an array of key questions in open-economy macroeconomics regarding the efficiency of capital allocation, the volatility of consumptions, and economic growth. Although the theoretical importance of RIP as well as its validity for analyzing issues related to fiscal and monetary policy are important, empirical support for RIP in the literature is elusive.

A number of studies of OECD countries provide support for long-run RIP based on panel data.<sup>2</sup> One common explanation for this finding is that increasing the amount of information on real interest rates typically increases the power of unit root tests and overcome the

issue of the low power of early univariate unit root studies.<sup>3</sup> On the other hand, Rose (2014) shows that the existence of bond market under inflation-targeting is associated with stable inflation because it creates an effective safeguard for low inflation.<sup>4</sup> As shown by Svensson (2000), Mishkin and Schmidt-Hebbel (2007), and Kim (2014), the high degree of transparency and accountability of monetary policy limits not only variability in inflation but also that in the real exchange rate at a long horizon, thereby stabilizing real exchange rates to a significant amount relative to the cases under other monetary regimes.

Various industrial and emerging countries have explicitly used an inflation target as their nominal anchor since New Zealand adopted inflation-targeting in 1990.<sup>5</sup> As shown by Svensson (2000) and Mishkin and Schmidt-Hebbel (2007), what made this monetary policy regime special was the explicit public commitment to stabilizing inflation as the main policy target and the emphasis on monetary policy transparency and accountability. This new monetary policy regime is characterized by (1) explicit quantitative inflation targets, (2) a policy approach based on a forward-looking assessment, namely use of an internal conditional inflation forecast as an intermediate target variables, and (3) a high degree of transparency and accountability.<sup>6</sup>

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<sup>1</sup> See Dornbusch (1976) and Mussa (1982) for details.

<sup>2</sup> See Wu and Chen (1998) and Taylor and Taylor (2004) for details.

<sup>3</sup> However, Taylor and Sarno (1998) issue an important warning related to the spurious interpretation of findings derived from panel data.

<sup>4</sup> Inflation-targeting is a monetary policy in which a central bank has an explicit target inflation rate. For details regarding bond markets and inflation-targeting, see Rose (2014).

<sup>5</sup> Today 26 countries use IT. See Roger (2010) and Mishkin and Schmidt-Hebbel (2007) for details.

<sup>6</sup> See Svensson (2000) and Mishkin and Schmidt-Hebbel (2007) for details.

Svensson (2000) provides a theoretical framework for a small open economy with exchange rate channels for the transmission of monetary policy to inflation and shows evidence that since inflation-targeting reduces variability in relative prices, the long-run unconditional variances of real exchange rates in flexible inflation-targeting cases are smaller than those in other cases.<sup>7</sup> Further, empirical evidence including Mishkin and Schmidt-Hebbel (2007) on the link between inflation-targeting and particular measures of economic performance also shows that inflation-targeting is associated with an improvement on overall economic performance in that inflation levels, inflation volatility, and interest rates have declined after countries adopted inflation targeting. The important hypothesis in the present study is that if the theory and evidence were right and at the same time if PPP were to hold better and the bond market were correlated with low inflation in countries under inflation-targeting, inflation-targeting would play an important role to provide favorable evidence for RIP.

One fundamentally and empirically important issue to the present study is the degree to which the movements of goods and capital markets across countries can be measured by the level of economic integration. The answer to this question depends on the degree of economic integration between markets across economies. Because of the high persistence of interest rates as well as of goods' prices, least squares (LS) estimates of parity might appear to suffer from a downward bias in the persistent coefficient, implying that the parity condition is estimated spuriously to be less persistent than it actually is. In order to correct this bias, Andrews (1993), Andrews and Chen (1994) and Hansen (1999) have proposed approaches such as the median-unbiased estimator and grid bootstrap methods, respectively. However, while this potential bias has been recognized in the time series literature ever since the seminal findings of Kendall (1954), no empirical study has thus far carried out an estimation of bias-correction in order to examine the influence of inflation-targeting on RIP.

The other important point in question for understanding the parity condition is cross-sectional dependence. Panel unit root tests have been widely employed to investigate PPP and RIP, however, the results of such tests with cross-sectional dependence lend little support to PPP or RIP in contrast to tests without cross-sectional dependence.<sup>8</sup> Furthermore, Phillips and Sul (2003) show that if there exists serious cross-sectional dependence in the data and it is ignored in estimation, estimation efficiency can decrease so that the panel LS estimator may provide little efficiency gain over the single equation LS. Thus, it is interesting to examine whether inflation-targeting has a significant role for RIP in addition to other factors such as price indices and bias-corrected cross-sectional dependence in the panel data.

To test the influence of inflation-targeting in this regard and to estimate the half-life,<sup>9</sup> we use recursive mean adjustment (RMA) proposed by So and Shin (1999). According to So and Shin, the RMA estimator is computationally convenient and powerful, and has been employed in many studies. For instance, among many others, Taylor (2002) employs the RMA based seasonal unit root test and Sul et al. (2005) use RMA for heteroscedasticity and autocorrelation consistent estimation.<sup>10</sup> Further, Choi et al. (2010) develop a RMA based bias correction method for dynamic panel data and Chudik and Pesaran (2015) apply RMA to common correlated effects approach for heterogeneous panel data models with lagged dependent variable. They find that the proposed estimators have satisfactory performance to correct the bias.<sup>11</sup>

In this study, the bias-correction method is applied to the cross-sectionally augmented versions of the tests of Im et al. (2003) (IPS) and

Pesaran (2007) (CIPS) for panel data. The RMA method is also used to estimate the convergence rates to RIP for inflation-targeting and non-inflation-targeting countries correctly without bias. Moreover, in order to avoid possible aggregation bias because of heterogeneous dynamics in cross-sector aggregate prices, we use sectoral consumption data by type and implicit deflators for durable goods' and service consumptions to construct the real interest rates for durables and service consumption, respectively among seven industrialized countries.<sup>12</sup> Comparisons are made, together with durables and service consumption including producer price index (PPI) and consumer price index (CPI), between inflation-targeting and non-inflation-targeting, and with and without cross-sectional dependence.

The empirical findings based on the results of the panel unit root tests presented herein differ depending upon whether we use RMA, as do the convergence rates in terms of the half-life estimates. Despite the price indices, numeraire currencies, and cross-sectional dependence, the half-life estimates consistently show that inflation-targeting countries have shorter half-lives than non-inflation-targeting countries, while those for all countries lie in between. The empirical results further show that inflation-targeting lowers the variability in real interest rates, providing more favorable evidence for RIP, as more inflation-targeting countries become involved. Further, especially under inflation-targeting, the result is not likely to be sensitive to numeraire currencies, price indices, or cross-sectional dependence; however, correcting for bias does not increase the tendency to reject the unit root hypothesis with cross-sectional dependence in our sample.

## 2. Econometric model and estimation

RIP involves both UIP and PPP.

$$r_t - r_t^* = \epsilon_t \tag{1}$$

where  $r_t = i_t - (p_{t+1}^T - p_t^T)$ ,  $r_t^* = i_t^* - (p_{t+1}^{T*} - p_t^{T*})$ ,  $i_t$  ( $i_t^*$ ) is the domestic (foreign) nominal interest rate and  $p_t^T$  ( $p_t^{T*}$ ) is the log of the domestic (foreign) price of traded goods at time  $t$ .<sup>13</sup> Under the condition of perfect arbitrage in the traded goods and capital markets, Eq. (1) is relevant for tests of international parity. Given the fact that the composite error that arises from expectational errors in UIP, conditional on the current information set, is stationary, Eq. (1) indicates that *ex post* RIP, defined in terms of traded goods between domestic and foreign countries, holds.<sup>14</sup>

To test the long-run relationship in (1), first we consider the following regression:

$$\epsilon_t = \alpha + \beta \epsilon_{t-1} + e_t \tag{2}$$

where  $\epsilon_t$  is the real interest rate differential at time,  $t$ , and  $e_t$  is a white noise error. As mentioned above, potential downward bias exists in the LS estimator for  $\beta$  and this can become particularly severe as the true value of the parameter approaches unity. To overcome this bias, we use the RMA estimator proposed by So and Shin (1999) and Shin and So (2001). By defining the recursive mean,  $\bar{\epsilon}_{t-1} = (t-1)^{-1} \sum_{k=1}^{t-1} \epsilon_k$  and rewriting Eq. (2) we derive:

$$\epsilon_t - \bar{\epsilon}_{t-1} = \beta_{RMA} (\epsilon_{t-1} - \bar{\epsilon}_{t-1}) + e_t \tag{3}$$

<sup>12</sup> We implicitly assume that sectoral heterogeneity can induce different convergence rates in our data like traded and non-traded goods. See Imbs et al. (2005) for details.

<sup>13</sup> To see this, UIP between two countries can be shown as  $i_t - i_t^* = s_{t+1} - s_t + \epsilon_t$ , where  $s_t$  is the natural logarithm of the exchange rate between a domestic and a foreign country (domestic price of foreign currency),  $\epsilon_t = E(s_{t+1}|I_t) - s_{t+1}$  is a composite error term assumed to be white noise, and  $E(\cdot|I_t)$  is the conditional expectations operator based on the information at time  $t$ . PPP for traded goods is  $s_t = p_t^T - p_t^{T*}$ . Combining these two yields Eq. (1).

<sup>14</sup> Differential tax treatment and transactions costs may result in the existence of a neutral band for financial market speculation within which profitable trading opportunities are impossible. Thus, international financial integration will result in the stationarity of real interest rate differentials. For details, see Wu and Chen (1998).

<sup>7</sup> He considers four different IT cases. See Svensson (2000) for details.

<sup>8</sup> See O'Connell (1998) and Moon and Perron (2007) for details.

<sup>9</sup> Half-life measures the number of years for a shock to decay by 50%.

<sup>10</sup> Kim and Moh (2012) also provided with empirical evidence of more powerful RMA based unit root tests.

<sup>11</sup> In addition, Choi et al. (2010) explain why the RMA method works well when the dominant root is near unity among several bias correction methods.

According to Shin and So (2001),  $\hat{\beta}_{RMA}$  reduces this bias substantially compared with the LS estimator of  $\beta$ . Extending the RMA estimation to panel data is straightforward. For a dynamic panel model, first we consider that  $e_{it}$  in Eq. (2) is allowed to be serially correlated for country  $i$  ( $i=1, 2, \dots, N$ ) at time  $t$  and has a single common factor structure:

$$e_{it} = \gamma_i f_t + \varepsilon_{it} \tag{4}$$

where  $f_t$  is an unobserved common factor,  $\gamma_i$  is the individual factor loading, and  $\varepsilon_{it}$  is a white noise idiosyncratic error. The IPS and CIPS tests are then used together with RMA to examine the stationarity of real interest rate differentials. Following Shin et al. (2004), a test based on the  $t$  ratio of the LS estimate of  $b_i$  is considered in the cross-sectionally augmented Dickey–Fuller (CADF) regression combined with RMA for each cross-sectional unit, as suggested by Pesaran (2007),

$$\Delta \varepsilon_{it} = b_i(\varepsilon_{it-1} - \mu_i) + c_i(\bar{\varepsilon}_{t-1} - \mu_i) + \sum_{j=0}^{p_i} d_{ij} \Delta \bar{\varepsilon}_{t-j} + \sum_{j=1}^{p_i} \delta_{ij} \Delta \varepsilon_{i,t-j} + \eta_{it} \tag{5}$$

where  $\Delta \varepsilon_{it} = \varepsilon_{it} - \varepsilon_{i,t-1}$ ,  $\mu_i = \bar{\varepsilon}_{i,t-1} = (t-1)^{-1} \sum_{s=1}^{t-1} \varepsilon_{is}$ ,  $\bar{\varepsilon}_t = \frac{1}{N} \sum_{i=1}^N \varepsilon_{it}$ ,  $\Delta \bar{\varepsilon}_t = \frac{1}{N} \sum_{i=1}^N \Delta \varepsilon_{it}$ ,  $p_i$  is the lag length determined by Hall's (1994) general-to-specific method<sup>15</sup> and  $\eta_{it}$  is the idiosyncratic disturbance which is assumed to be cross-sectionally independent. According to Pesaran (2007), the cross-sectional averages of  $\Delta \varepsilon_{it}$  and  $\varepsilon_{i,t-1}$  are included in (5) as a proxy for the unobserved common factor  $f_t$ . The null hypothesis,  $H_0: \hat{b}_i = 0$ , for all  $i$  is tested against the heterogeneous alternative  $H_1: \hat{b}_1 < 0, \dots, \hat{b}_{N_0} < 0, N_0 \leq N$  in the whole panel set. In line with the findings of Im et al. (2003), Pesaran (2007) proposes the CIPS test:

$$CIPS = \frac{1}{N} \sum_{i=1}^N CADF_i \tag{6}$$

where  $CADF_i$  is the CADF statistic for the  $i$ -th cross-sectional unit in Eq. (5). The distribution of the CIPS statistic is shown to be non-standard even for large  $N$ . We also use the other panel unit root test namely the IPS test, which is based on the  $t$  ratio of the LS estimate of  $b_i$  in Eq. (5) without the cross-sectional average terms. In contrast to the CIPS test, whose distribution is shown to be non-standard even for large  $N$  as noted above, this procedure assumes that the individual time series are cross-sectionally independently distributed.

### 3. Empirical results

We use quarterly data from 1974:Q1 to 2012:Q1. Our interest rate measure is the three-month Treasury bill rate taken from the International Financial Statistics and Data Stream. To measure inflation rates, in addition to CPI and PPI, we use durable goods' and service consumption classified by type for the following seven countries: Canada, France, Japan, Italy, Sweden, the United Kingdom and the United States.<sup>17</sup> To construct the inflation rate for traded and non-traded goods, we use implicit deflators for durable goods' and service consumption, respectively. For CPI and PPI, as proxies for the prices of non-traded goods and traded goods, we examine the following 11 OECD countries: Belgium, Canada, France, Germany, Italy, Japan, New Zealand, Spain, Sweden, the United Kingdom, and the U.S.<sup>18</sup>

<sup>15</sup> Note that in the RMA scheme,  $\varepsilon_{i,t-1}$  is adjusted for  $\mu_i$  by using the recursive sample mean. See Shin et al. (2004) for details.

<sup>16</sup> Start with  $k=10$  and decrease it until the coefficient of the last included lag is significant.

<sup>17</sup> The sample countries were selected based on the availability of data. German data, for non-service and service consumption necessary to construct traded and non-traded goods' prices were unavailable. Eleven OECD countries for CPI and ten OECD countries for PPI were also studied.

<sup>18</sup> France was excluded from the non-IT group for the PPI panel due to data availability. Data on some countries were not available in the full sample. For the three month  $t$ -bill rates, they are Germany (1975Q3-2007Q3), New Zealand (1978Q1:2011Q3), and Spain (1979Q1:2011Q3). For the PPI panel, they are Italy

**Table 1**  
Inflation targeting countries.

Countries	Year of adoption	Target level
New Zealand	1990	1–3%
Canada	1991	1–3%
U.K.	1992	2%
Sweden	1993	2%
Spain	1994	NA

Notes: NA means not available. Spain adopted inflation targeting and abandoned it when they began to use Euro as their currency. All information are from central banks' websites.

Though some industrialized economies such as European Monetary Union, the U.S., Japan, and Switzerland have resembled many of the main elements of inflation targeting, the five countries which explicitly engage in inflation-targeting are only considered due to data availability.<sup>19</sup> The inflation rates used to generate the *ex post* real interest rates in our empirical study are calculated by taking the actual inflation rates from period  $t$  to period  $t+1$ . To test whether inflation-targeting affects RIP, we classify countries based on whether their central banks have adopted IT. The countries that engage in inflation-targeting in this study are therefore New Zealand (1990), Canada (1991), the United Kingdom (1992), Sweden (1993), and Spain (1994) and Table 1 summarizes these countries. Further, because many authors have noted the problem caused by choosing the U.S. as the base country,<sup>20</sup> additional countries, including Italy for durable goods' and service consumption, and Germany for CPI and PPI, are also considered to be base countries in this paper.<sup>21</sup>

Table 2 presents the results of our tests for cross-sectional dependence. The general diagnostic test proposed by Pesaran (2004) is used to test whether cross-sectional dependence exists in our data. We ran individual ADF regressions for each country and compute the pair-wise cross-sectional correlation coefficients of the residuals from these regressions in the panel.<sup>22</sup> As can be seen, the null hypothesis of no cross-sectional dependence is strongly rejected in all cases. Therefore, we should consider CIPS and CADF more reliable than IPS or ADF.

Table 3 reports the results of the panel-based tests, namely the conventional IPS and CIPS tests with or without RMA for the real interest rate differentials for durables and service consumption including CPI and PPI. The  $p$ -values are taken from the non-parametric bootstraps in order to provide a precise inference.<sup>23</sup> Overall, the results seem to indicate that correction for bias does not necessarily increase the tendency to reject the unit root hypothesis, especially for the CIPS tests in our sample.<sup>24</sup> More specifically, the empirical results derived from the IPS and CIPS tests depend on whether we use RMA because the standard IPS tests show stronger rejections of the null hypotheses than the tests with cross-sectional dependence and RMA, while the CIPS tests without RMA lie in between in terms of the rejection rates, implying that cross-sectional dependence for the parity condition still important in our cases. This empirical evidence shows that the LS estimates of the parity from the IPS and CIPS tests seem to suffer from downward bias in the coefficient. However, the same is not true for the CIPS tests with RMA, where the rejection rates are lower than those of the IPS and CIPS tests.

(footnote continued)

(1981Q1:2011Q3) and Belgium (1980Q1:2011Q3).

<sup>19</sup> See Roger (2010) for details.

<sup>20</sup> See Cumby and Mishkin (1986) and Wu and Chen (1998) for more details.

<sup>21</sup> Real interest rate differentials are also alternatively defined with respect to non-IT countries such as Japan and France, but the results are similar. These results are available upon request from the author.

<sup>22</sup> See Pesaran (2004) for details.

<sup>23</sup> We also used a parametric bootstrap for the IPS and CIPS tests. As the results do not depend on the normality assumption, only the nonparametric results are reported here.

<sup>24</sup> See Cook (2005) for details.

**Table 2**  
Tests for cross-sectional dependence.

Base countries	Durables			Services		
	CD <sub>IT</sub>	CD <sub>ALL</sub>	CD <sub>NIT</sub>	CD <sub>IT</sub>	CD <sub>ALL</sub>	CD <sub>NIT</sub>
U.S.	4.817 (0.000)	13.689 (0.000)	8.437 (0.000)	4.534 (0.000)	16.409 (0.000)	9.882 (0.000)
IT	7.037 (0.000)	21.884 (0.000)	11.439 (0.000)	7.709 (0.000)	23.299 (0.000)	12.118 (0.000)
	PPI			CPI		
U.S.	17.753 (0.000)	32.594 (0.000)	11.297 (0.000)	4.333 (0.000)	21.031 (0.000)	19.474 (0.000)
GM	10.250 (0.000)	18.255 (0.000)	5.782 (0.000)	4.882 (0.000)	19.329 (0.000)	11.710 (0.000)

Notes: All statistics are based on individual ADF regressions. U.S., IT, and GM are the United States, Italy, and Germany, respectively. Subscripts IT, ALL, and NIT represent inflation targeting (IT), all, and non-inflation targeting (NIT) countries, respectively. Numbers in parentheses are *p*-values of CD statistics.

**Table 3**  
Real interest rate differentials for traded and non-traded goods.

Base countries	IPS			CIPS			CIPS <sup>RMA</sup>		
	$\bar{u}_{IT}$	$\bar{u}_{ALL}$	$\bar{u}_{NIT}$	$\bar{u}_{IT}$	$\bar{u}_{ALL}$	$\bar{u}_{NIT}$	$\bar{u}_{IT}$	$\bar{u}_{ALL}$	$\bar{u}_{NIT}$
<i>Durables</i>									
U.S.	-2.812 (0.000)	-2.485 (0.002)	-2.260 (0.040)	-3.021 (0.039)	-2.575 (0.047)	-2.414 (0.171)	-2.048 (0.034)	-1.156 (0.086)	-0.608 (0.287)
IT	-2.575 (0.004)	-2.361 (0.008)	-2.148 (0.041)	-3.131 (0.024)	-2.604 (0.027)	-1.639 (0.712)	-2.232 (0.023)	-0.975 (0.225)	-0.963 (0.245)
<i>Services</i>									
U.S.	-2.246 (0.036)	-2.122 (0.052)	-1.717 (0.229)	-2.336 (0.192)	-1.993 (0.376)	-1.916 (0.227)	-1.705 (0.107)	-0.947 (0.216)	-0.472 (0.292)
IT	-2.079 (0.045)	-2.011 (0.059)	-2.001 (0.062)	-2.291 (0.273)	-1.952 (0.428)	-1.822 (0.536)	-1.274 (0.132)	-1.069 (0.227)	-0.849 (0.265)
<i>PPI</i>									
U.S.	-2.995 (0.000)	-2.785 (0.000)	-2.522 (0.008)	-3.248 (0.000)	-2.748 (0.005)	-2.069 (0.108)	-1.847 (0.004)	-1.517 (0.006)	-0.748 (0.196)
GM	-2.290 (0.003)	-1.942 (0.006)	-1.507 (0.173)	-2.685 (0.012)	-2.683 (0.003)	-2.027 (0.173)	-1.648 (0.013)	-1.307 (0.032)	-0.622 (0.307)
<i>CPI</i>									
U.S.	-2.118 (0.011)	-2.372 (0.000)	-2.575 (0.002)	-2.442 (0.048)	-2.410 (0.043)	-1.995 (0.101)	-1.307 (0.077)	-0.953 (0.214)	-0.791 (0.218)
GM	-1.824 (0.029)	-1.552 (0.053)	-1.335 (0.248)	-2.592 (0.034)	-2.349 (0.033)	-2.447 (0.018)	-1.109 (0.128)	-0.967 (0.170)	-0.822 (0.227)

Notes: Numbers in parentheses are *p*-values taken from 10,000 nonparametric bootstrap simulations.

According to our empirical evidence, inflation-targeting countries with durable goods provide greater support for the RIP condition compared with all and non-inflation-targeting countries with both durable goods' and service consumption as well as CPI and PPI.

Table 4 shows the estimates of half-lives and 95% confidence intervals based on the ADF, CADF, and recursive mean-adjusted CADF (RMACADF) regressions. To assess the convergence rates, we first estimated the system using seemingly unrelated regressions (SUR) and conducting likelihood ratio (LR) tests for homogeneity restrictions across the system. We found that none of the tests rejects the null hypothesis at the 5% significance level, and we therefore follow simple panel AR(p) models with homogeneous restrictions across the system for ADF, CADF, and RMACADF using SUR.<sup>25</sup> The 95% confidence intervals are computed from the nonparametric bootstrap simulations.<sup>26</sup> The half-life estimates for inflation-targeting and non-infla-

tion-targeting countries suggest two distinctive characteristics. First, compared with the half-lives from the CADF with RMA tests, the LS estimates of half-lives with or without cross-sectional dependence are seriously downward-biased. For instance, when the U.S. dollar is used as the base currency, the estimated half-lives with RMA under cross-sectional dependence for inflation-targeting countries are 0.62 and 1.41 years, and for non inflation-targeting countries they are 1.43 and 2.30 years for durable goods and PPI, respectively, while those for services and CPI range from 1.32 and 1.64 years for inflation-targeting countries and 2.32 and 2.95 years for non inflation-targeting countries. Further, the 95% confidence intervals show that the range for the estimated half-life for the inflation-targeting group is narrower than that of non-inflation-targeting countries. Moreover, the LS estimates of half-lives with or without cross-sectional dependence provide much shorter half-lives than those with RMA and cross-sectional dependence for inflation-targeting and non-inflation-targeting countries. More importantly, there exists a significant difference between inflation-targeting and non-inflation-targeting countries in terms of the convergence rates, implying that the half-life estimates for inflation-

<sup>25</sup> For instance, particularly for the RMACADF, instead of Eq. (5), we use  $\epsilon_{it} - \mu_i = b_i(\epsilon_{it-1} - \mu_i) + c_i(\bar{\epsilon}_{t-1} - \mu_i) + \sum_{j=0}^{p_i} d_{ij} \Delta \bar{\epsilon}_{t-j} + \sum_{j=1}^{p_i} \delta_{ij} \Delta \epsilon_{it-j} + \eta_{it}$ .  
<sup>26</sup> For details, see Efron and Tibshirani (1993) and So and Shin (1999).



**Table 4**  
Half-lives and 95% confidence intervals for real interest rate differentials.

Base countries	HL			HL <sub>CD</sub>			HL <sub>CD</sub> <sup>RMA</sup>		
	HL <sub>IT</sub>	HL <sub>ALL</sub>	HL <sub>NIT</sub>	HL <sub>IT</sub>	HL <sub>ALL</sub>	HL <sub>NIT</sub>	HL <sub>IT</sub>	HL <sub>ALL</sub>	HL <sub>NIT</sub>
<i>Durables</i>									
U.S.	0.69 (0.48, 0.88)	0.80 (0.59, 0.95)	0.87 (0.57, 1.12)	0.44 (0.30, 0.55)	0.65 (0.47, 0.74)	0.82 (0.52, 1.07)	0.62 (0.48, 0.85)	1.21 (1.03, 1.79)	1.43 (1.05, 2.37)
LR	0.28 (0.96)	4.12 (0.53)	5.46 (0.14)	5.74 (0.12)	0.58 (0.98)	2.51 (0.47)	7.80 (0.05)	0.99 (0.96)	4.99 (0.17)
IT	0.52 (0.40, 0.63)	0.58 (0.48, 0.67)	0.77 (0.57, 0.95)	0.54 (0.36, 0.69)	0.76 (0.53, 0.89)	1.06 (0.60, 1.41)	1.27 (1.04, 2.31)	1.72 (1.47, 2.90)	2.03 (1.22, 3.35)
LR	0.55 (0.90)	2.20 (0.81)	0.86 (0.83)	6.38 (0.09)	6.94 (0.22)	0.74 (0.86)	0.42 (0.93)	0.56 (0.99)	0.95 (0.81)
<i>Services</i>									
U.S.	0.78 (0.56, 0.97)	1.09 (0.82, 1.26)	1.62 (1.03, 2.06)	0.68 (0.44, 0.87)	0.76 (0.54, 0.88)	1.10 (0.63, 1.44)	1.32 (0.88, 2.33)	2.23 (1.54, 3.67)	2.32 (1.27, 4.14)
LR	0.33 (0.95)	3.50 (0.62)	7.15 (0.07)	4.76 (0.19)	0.79 (0.98)	7.16 (0.06)	1.91 (0.59)	0.34 (0.99)	1.67 (0.64)
IT	0.66 (0.54–0.86)	1.18 (0.94, 1.49)	1.68 (1.18–2.45)	0.53 (0.33, 0.69)	0.71 (0.47, 0.84)	0.72 (0.40, 0.97)	1.46 (1.14, 2.74)	1.90 (1.67, 3.54)	2.06 (1.37, 4.45)
LR	0.78 (0.85)	2.11 (0.83)	1.49 (0.68)	1.44 (0.69)	7.19 (0.20)	1.53 (0.67)	2.26 (0.52)	0.87 (0.97)	4.12 (0.25)
<i>PPI</i>									
U.S.	0.71 (0.50, 0.85)	0.99 (0.70, 1.12)	1.51 (0.84, 1.93)	0.77 (0.55, 0.91)	0.94 (0.69, 1.03)	1.42 (0.85, 1.77)	1.41 (1.20, 2.29)	1.83 (1.59, 2.77)	2.30 (1.47, 3.65)
LR	1.97 (0.74)	8.72 (0.36)	4.68 (0.19)	4.52 (0.34)	12.84 (0.11)	7.63 (0.05)	2.70 (0.61)	5.59 (0.69)	0.88 (0.83)
GM	1.18 (0.73, 1.49)	1.12 (0.76, 1.29)	1.73 (0.89, 2.26)	0.76 (0.53, 0.90)	0.77 (0.57, 0.84)	1.13 (0.70, 1.43)	1.18 (0.96, 1.78)	1.54 (1.38, 2.33)	2.29 (1.59, 4.22)
LR	0.69 (0.95)	5.42 (0.71)	1.59 (0.66)	1.50 (0.82)	2.90 (0.94)	6.87 (0.07)	4.92 (0.29)	2.00 (0.98)	2.19 (0.53)
<i>CPI</i>									
U.S.	0.83 (0.59, 1.00)	0.98 (0.73, 1.11)	1.24 (0.79, 1.56)	0.72 (0.50, 0.84)	1.30 (0.90, 1.42)	1.89 (1.08, 2.28)	1.64 (1.29, 2.61)	2.68 (2.05, 3.82)	2.95 (1.89, 5.04)
LR	0.99 (0.91)	8.49 (0.38)	7.33 (0.06)	2.02 (0.73)	14.37 (0.07)	5.41 (0.14)	6.24 (0.10)	5.813 (0.67)	1.24 (0.81)
GM	0.95 (0.65, 1.17)	1.19 (0.84, 1.37)	1.91 (1.03, 2.43)	0.77 (0.53, 0.90)	0.99 (0.71, 1.11)	1.47 (0.87, 1.73)	1.34 (1.13, 2.20)	2.32 (2.17, 4.06)	3.15 (2.39, 7.09)
LR	1.02 (0.90)	5.53 (0.69)	4.05 (0.25)	3.87 (0.42)	5.01 (0.75)	5.00 (0.17)	4.26 (0.23)	4.46 (0.81)	3.07 (0.55)

Notes: HL (half-lives) are number of years for a shock to decay by 50 confidence intervals by taking 2.5 and 97.5 percentiles from 10,000 nonparametric bootstrap simulations at the estimates using SUR, the estimates from CADF using SUR and at the RMA estimates from CADF using SUR. We followed Efron and Tibshirani (1993) and So and Shin (1999). LR represents likelihood ratio test for a homogeneous restriction and numbers in parenthesis are *p*-values.

targeting countries are typically shorter than those for non inflation-targeting countries, while the estimates for all countries lie in between. Once again, regardless of whether RMA is used, our empirical evidence consistently shows shorter convergence rates for inflation-targeting countries with traded goods than for all and non inflation-targeting countries at any prices.

Our main empirical findings for RIP under inflation-targeting are threefold. First, RIP for inflation-targeting countries seems to be strongly supported by the data used herein. As both inflation-targeting and non inflation-targeting countries are studied in order to examine RIP in the present paper, the results obtained from the panel unit root tests and half-life estimates are consistent and thus lend greater support to the hypothesis of RIP among inflation-targeting countries compared with all and non-inflation-targeting countries. For this latter group, the results are likely to be somewhat sensitive to durable goods' and service consumption, price indices, base countries, and cross-sectional dependence. However both the IPS and the CIPS tests with or without RMA strongly reject the null hypothesis for inflation-targeting countries at lower *p*-values compared with non-inflation-targeting countries. In particular, this result is not sensitive to price indices, especially CPI and PPI, base countries, and cross-sectional dependence. The same finding holds for the half-life estimates, implying that despite the utilization of the cross-sectional dependence and RMA method, the convergence rates are growing faster and the 95% confidence intervals are tighter for inflation-targeting countries compared with all and non-inflation-targeting countries.

Second, RIP for durable goods is broadly supported by the present results, while RIP for service consumption is not, with the exception of those inflation-targeting countries whose RIP hypotheses are supported strongly from the results of the panel unit root tests and the

half-life estimates regardless of price indices such as CPI and PPI. However, RIP for the prices of service consumption is likely to be sensitive to cross-sectional dependence. As durables and service consumption are used for all countries, we find more empirical support for RIP for durable goods compared with service consumption. The same is true for CPI and the PPI. Although the results in the tables show that the price indices are unlikely to influence RIP for inflation-targeting countries, the durable goods' measures of real interest rate differentials for these countries are superior to those constructed from service consumption for all and non-inflation-targeting countries in terms of rejection rates, low *p*-values, and convergence rates. The results are, however, somewhat sensitive to the base country and cross-sectional dependence.

Third, there is evidence of cross-sectional dependence, depending on whether they adopt IT. In Tables 2–4, for all cases and for the non-inflation-targeting cases considered, the empirical results from the CIPS test outweigh the IPS support for RIP for durable goods' and service consumption, implying that they provide broad support for the findings of Moon and Perron (2007) and are robust to the base country. For inflation-targeting countries, by contrast, we are able to find support for RIP despite the existence of cross-sectional dependence.

#### 4. Conclusion

In sharp contrast to the findings on RIP presented in previous studies, our empirical results indicate a tendency for real interest rate differentials for inflation-targeting countries to converge. All real interest rates for inflation-targeting countries consistently show that they have shorter half-lives, while the RIP conditions are more likely to hold than for all countries or for non-inflation-targeting countries.

However, the empirical findings based on the IPS and CIPS tests somewhat depend upon whether we use RMA. In particular, while correcting for bias does not tremendously increase the tendency to reject the unit root hypothesis with cross-sectional dependence in our sample, the tests without the RMA show that the LS estimates of the parity from the IPS and CIPS tests seem to suffer from downward bias in the persistent coefficient, implying that the parity condition is estimated spuriously to be less persistent than it actually is. The empirical evidence presented in this study thus seem to confirm that inflation-targeting influences RIP and that stronger mean reversion as well as much a tighter confidence interval are present as more inflation-targeting countries become involved. Moreover, the evidence in favor of RIP for inflation-targeting countries does not seem to be sensitive to the choice of price index such as CPI and PPI, or to cross-sectional dependence. Further, it seems as though a test with RMA and cross-sectional dependence provides somewhat qualitatively different results, but the empirical results depend on both inflation-targeting and the price index. The prices of traded goods are among the most likely to exhibit evidence of short-run and long-run PPP, because trade between European countries and major trading partners involves relatively low transaction costs and faces relatively stable non-tariff barriers to trade. As more homogeneous countries are involved in terms of inflation-targeting regime, a more stylized economic evidence for RIP is found. The empirical evidence in this study is interesting and in line with Svensson (2000), Mishkin and Schmidt-Hebbel (2007), and Rose (2014) in showing that inflation-targeting influences inflation-targeting and non-inflation-targeting policy regimes by helping them both create an effective safeguard for stable inflation and achieve lower variability in inflation as well as that in the real interest rate at a long horizon. In addition, it also plays an important role in providing support for RIP, implying that under inflation-targeting neither the cross-sectional dependence nor the price index are crucial for understanding the RIP condition.

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