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The Behavior of Real Exchange Rates: The Case of Japan

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ABSTRACT

The study examines the convergence rate of mean reversion by contrasting the estimated half-life of real exchange rate (RER). We employ an extensive monthly consumer price index (CPI)-based product price's panel for Japan (the U.S. as the numéraire). We find that the disaggregated RERs are persistent due to the cross-sectional dependence problems. By controlling common correlated effects, the estimated half-life for all goods may fall to as low as 2.54 years, below the consensus view of 3 to 5 years summarized by Rogoff (1996). After correcting the small-sample bias, the estimated half-life of deviations from purchasing power parity (PPP) increase by 1.03 year. Our findings also support that the half-life of mean reversion of RER is about 3.55 years for traded goods, about 0.11 year lower than non-traded goods. We also show that traded goods and non-traded goods perform distinct distributions of persistence.

JEL classification: C33, F31

Keywords: Common correlated effect, cross-sectional dependence, purchasing power parity, real exchange rate, traded and non-traded goods

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1 Introduction

The purchasing power parity (henceforth PPP) is a fundamental empirical hypothesis that in the absence of transaction costs and other distribution costs, national price levels should be equal if converted to a common currency. By definition, the real exchange rate (RER) can be expressed as the nominal exchange rate adjusted for relative national price levels, and hence variations in the RER imply deviations from PPP. In other words, the RER must be stationary if PPP holds. If RERs are stationary, it means that deviations from the parity are temporary and ultimately self-reverting. Rogoff (1996) argues that the aggregate RERs are stationary, but pretty persistent, with the estimated half-life of 3-5 years, so-called PPP puzzle.¹ The presence of nominal rigidities, such as, prices or wages, cannot explain so slow rate to disappear, consensus estimated by most empirical evidences.

Though PPP performs poorly in the most aggregate series, many economists still believe that national relative sectoral or product level prices may move in proportion to the adjustment in the nominal exchange rate so that the RER will revert to its parity soon (e.g., Imbs, Mumtaz, Ravn and Rey, 2005; Cheung and Fujii, 2008; Robertson, Kumar and Dutkowsky, 2009). Recently, Imbs et al. (2005) offer a possible way to resolve this puzzle. Imbs et al. (2005) investigate how sectoral heterogeneity in convergence rates to the law-of-one-price (LOP) may lead to upward (positive) bias in the estimation of half-life. They claim that the PPP puzzle can be successfully resolved if the heterogeneity and aggregation bias are considered. However, Chen and Engel (2005) argue that the half-life estimations are still similar to Rogoff's consensus after corrections of small sample bias, entry errors and missing nominal exchange rate data.

Macroeconomists therefore have become interested in investigating for PPP as a sectoral or product level price data. Recently, Robertson et al. (2009) use a panel disaggregated price data between Mexico and the U.S. to investigate the parity and find that PPP holds.

¹By definition, the half-life is used to measure the time required for deviation from PPP to dissipate by a half.

Cheung and Fujii (2008), for example, examine the individual retail prices of intra-Japan for the deviations of product-specific LOP. Cheung and Fujii find that the deviations from individual LOP are considerable persistence and traded products, compared with non-traded, have different distributions of LOP deviations across cities. Burstein, Neves and Rebelo (2003) argue that the distribution costs play a major role in PPP. Burstein et al. (2009) find that the fraction of the retail price accounted for by distribution costs is quite large. Engel (1999) finds that the relative prices of traded goods explain the most variability of the U.S. RERs. On the contrary, Imbs et al. (2005) find that the degree of sectoral heterogeneity is lower in non-traded goods rather than traded goods. Crucini and Shintani (2008) investigate LOP persistence using the worldwide retail prices from micro-data and find traded good have less persistence than non-traded goods in all locations. In an interesting study, Parsley and Wei (2007) decompose the Big Mac's RER using a subset of hamburger inputs data and find that traded ingredients are less persistence than non-traded ingredients.

In this paper, we examine the seasonally-adjusted Japanese consumer price index (CPI)-based product level's RER (the numéraire is the U.S.) for two distinct categories, traded and non-traded goods. Our price data are made up of specific products, such as ham, bananas and postage, etc. The levels of disaggregation are comparable to Crucini and Shintani (2008) and Parsley and Wei (2007). In order to study the average persistence of these two categories, we are not only taking account of the estimation issues in the conventional PPP literature, such as heterogeneity, cross-sectional dependence, and small-sample bias, but also consider the possible miss-specification of the optimal lags of the autoregression (AR) in our model.²

The main findings of the study are summarized as follows. We find that the magnitude of product-level aggregation bias plays a central role for our estimations. Particularly dramatic decreases in turn are obtained after accounting for common correlated effects (CCE). Importantly, if the cross-sectional dependence is not considered into our estimation, they may reach 8.20 years for all product level goods prices (7.96 years for traded goods and 8.86

²According to Chen and Engel (2005), we report the bias-corrected half-life by a bootstrap approach. Regarding the optimal lags, see, Rossi (2005).

years for non-traded goods). By the comparison of AR specifications, we conclude that the mis-specification bias is important for examining the RER persistence (see also, Rossi, 2005).

We divide whole items into two categories and find that traded goods are less persistence than non-traded goods in all tests. The point estimation by the common correlated effects mean group (CCEMG) with small-sample bias correction for traded goods' half-life is about 3.55 years, 0.11 year lower than that for non-traded goods. Our findings also indicate slightly difference in estimated half-lives between traded and non-traded goods. Our results support Crucini and Shintani's (2008) arguments that the traded goods are less persistence than non-traded goods. Not surprisingly, because the traded goods require an important fraction of distribution services, these services are intensive in local labor and land hence are non-traded (see, Burstein et al., 2003). The findings are confirmed by alternative nonparametric approaches. Similarly, Parsley and Wei (2007) conclude that the traded inputs display less price dispersion than the non-traded inputs for Big Mac production. We also find that the correction of small-sample bias bears essential but does not change our conclusions (e.g., Chen and Engel, 2005).

The remaining parts of this paper are organized as follows. Section 2 presents half-life measurement and the data used. The empirical results are reported in Section 3. Section 4 concludes the paper.

2 Methodology and Data

To explore the dynamics of Japanese RERs, a simple and conventional way is to estimate their half-lives. We employ some standard panel approaches to investigate the RER persistence. For instance, we use disaggregated data estimations to deal with aggregation bias and calculate the average half-lives. Additionally, we will briefly express the methods used to cure small sample bias and to control for cross-sectional dependence, respectively.

2.1 Methodology

Suppose that there are N kinds of products in the economies of Japan and the U.S. Let P_{it} and P_{it}^* be the prices of goods i in Japan and the U.S., respectively, for $i = 1, \dots, N$, at time t , $i = 1, \dots, T$. Let S_t denote the Japan-U.S. nominal exchange rate. Suppose the logarithm of RER of the i th product at time t , $\ln Q_{it} = \ln S_t + \ln P_{it}^* - \ln P_{it}$, follows an AR process:

$$\begin{aligned} q_{it} &= \alpha_i + \theta_i t + \sum_{\kappa=1}^{\kappa_i} \rho_{i\kappa} q_{i,t-\kappa} + e_{it}, \\ &= \alpha_i + \theta_i t + \phi_i q_{i,t-1} + \sum_{\kappa=1}^{\kappa_i-1} \delta_{i\kappa} \Delta q_{i,t-\kappa} + e_{it}, \end{aligned} \quad (1)$$

where $q_{it} \equiv \ln Q_{it}$. α_i and t are the fixed effects (FE) and incidental time trend, $\phi_i = \sum_{\kappa=1}^{\kappa_i} \rho_{i\kappa}$, $\delta_{i\kappa} = -\sum_{j=\kappa+1}^{\kappa_i} \rho_{ij}$, $\Delta q_{i,t-\kappa} = q_{i,t-\kappa} - q_{i,t-\kappa-1}$, and $e_{it} \sim (0, \sigma_i^2)$. When $\kappa_i = 1$, $\phi_i = \rho_{i1}$ and $\delta_{i\kappa} = 0$. There are several features of this specification. First, we allow for the coefficient heterogeneity. It is well-known that imposing the parameter homogeneity in the panel data with slope heterogeneity and heteroskedasticity will potentially result in a bias in the slope coefficients, which is referred of as the aggregation bias by Imbs et. al. (2005).³ For comparison, we will further investigate the empirical results obtained from FE estimation and the generalized method of moments estimator (GMM) of Arellano and Bond (1991) to study the effect of imposing the assumption of slope homogeneity on half-life estimation.

Second, we do not restrict the individual AR lag order κ_i to be the same across distinct products. Rossi (2005) point out that the measurements of half-life should consider the optimal lag length. By Monte Carlo simulations, Rossi (2005) show that the bias of half-lives tends to have a substantial downward bias when the regression model is AR(1) but the true data generating process (DGP) is AR(κ_i). In this study the optimal lag length, κ_i , is selected based on the Akaike Information Criterion (AIC) for each i . To explore this possible bias, we report the estimated results from both AR(1) and AR(κ_i) models.

³Imbs et. al. (2005) used the AR(1) model to illustrate that there would exit an upward bias in aggregated half-life estimation when ρ_{i1} is positively correlated with σ_i^2 . On the other hand, Chen and Engel (2005) argue that the bias is not the main source of RERs' persistence and the aggregate bias may be positive or negative. To prevent from the potential bias, however, we still consider the possibility of coefficient heterogeneity in this work.

It is well-known that the bias of the least squares AR estimator diminishes at rate of $1/T$. Even when T is large, it is still crucial to correct this bias because for many products $\phi_i = \sum_{\kappa=1}^{\kappa_i} \rho_{i\kappa}$ are very close to unity and a tiny change might result in a huge difference in the half life estimates. Chen and Engel (2005) argue that the half-life estimates can be close to Rogoff's consensus view after small sample bias is corrected for each items. Therefore, we apply a bootstrap method to reduce this bias:

1. Randomly draw residuals \hat{e}_{it} from (1) to generate $\{e_{it}^{(r)}\}_{t=1}^{T+100}$ and generate:

$$q_{it}^{(r)} = \hat{\alpha}_i + \hat{\theta}_i t + \sum_{\kappa=1}^{\kappa_i} \hat{\rho}_{i\kappa} q_{i,t-\kappa} + e_{it}^{(r)}, \quad (2)$$

2. Drop the first 100 observations of $\{q_{it}^{(r)}\}_{t=1}^{T+100}$ and regress $q_{it}^{(r)}$ on $q_{it-1}^{(r)}, \dots, q_{it-\hat{\kappa}_i}^{(r)}$ to get $\hat{\rho}_{i1}^{(r)}, \dots, \hat{\rho}_{i\hat{\kappa}_i}^{(r)}$, by using the rest of observations, $r = 1, 2, \dots, R$.
3. Repeat $R = 1,000$ times to get the empirical $\hat{\rho}_{i\kappa}^*$, $\kappa = 1, \dots, \hat{\kappa}_i$, and the bias corrected estimator is obtained as:

$$\hat{\rho}_i^c = 2\hat{\rho}_{i\kappa} - \hat{\rho}_{i\kappa}^*. \quad (3)$$

In order to measure the persistence of the RERs, we estimate the half-life of RER for each item by:

$$\tau_{\text{HL},i} = \frac{\ln(0.5)}{\ln(\hat{\phi}_i^c)}.$$

However, the conventional average half-life measure, which is obtained by using the cross-sectional average AR coefficients in the above transformation, can be highly distorted. Instead, we follow the way of Gadea and Mayoral (2009) to evaluate the average half-life in the presence of cross-sectional heterogeneity in the AR coefficients by using:

$$\bar{\tau}_{\text{HL}} = \frac{1}{N} \sum_{i=1}^N \tau_{\text{HL},i} = \frac{1}{N} \sum_{i=1}^N \frac{\ln(0.5)}{\ln(\hat{\phi}_i^c)}, \quad (4)$$

where N denotes the number of items in the data.

In addition, Boivin, Giannoni and Mihov (2009) show that the macroeconomic fluctuations explain 15% of the variation on individual prices, which indicates that the prices of goods and services are usually simultaneously affected by macro policies and global shocks. The macro policies and global shocks can be regarded as the monetary policies and the fluctuation of crude oil prices. Due to these common factors, the conventional estimation, inference and (panel) units test will be invalid. To control for the dependence caused by a common factor, we also estimate the AR coefficients by using the method of Persaran (2006):

$$\begin{aligned} q_{it} &= \alpha_i + \theta_i t + \sum_{\kappa=1}^{\kappa_i} \rho_{i\kappa} q_{i,t-\kappa} + \sum_{\varsigma=1}^{\varsigma_i} \gamma_{i\varsigma} \bar{q}_{t-\kappa} + e_{it}, \\ &= \alpha_i + \theta_i t + \phi_i q_{i,t-1} + \sum_{\kappa=1}^{\kappa_i-1} \delta_{i\kappa} \Delta q_{i,t-\kappa} + \sum_{\varsigma=1}^{\varsigma_i} \gamma_{i\varsigma} \bar{q}_{t-\kappa} + e_{it}, \end{aligned}$$

where $\bar{q}_{t-\varsigma} = \frac{1}{N} \sum_{i=1}^N q_{i,t-\varsigma}$ and ς_i are various across i and are selected based on AIC.

2.2 Data

The Ministry of Internal Affairs and Communications (MIAC) posts product-level price indices for goods and services that compose the Japanese CPI. The product price data of Japan, consisting of more than 500 specific products, are collected from the Statistics Bureau, Director-General for Policy Planning & Statistical Research and Training Institute, MIAC, and are available at <http://www.stat.go.jp/>. The U.S. price data are obtained from the all urban consumers price index of U.S. Bureau of Labor Statistics (BLS) from the website <http://www.bls.gov/>.

We consider a balanced panel seasonally-adjusted monthly data from 1985:1 to 2009:6. Using the items descriptions, we match the Japanese price data with available the U.S. prices as closely as possible (for instance, we match wheat flour with flour and prepared flour mixes). The matched prices contain 304 items (224 for traded goods & 80 for non-traded goods). However, some specific Japanese products, for example, mochi and women's kimono, still

fail to match the similar U.S. products. The consumption expenditure weights (these are the national expenditure weights used to construct the Japanese CPI) for all price indices we included are 67.86% of Japanese consumption basket. The U.S. prices are converted to Japanese *yen* via the nominal Japanese-U.S. exchange rate at the end of the month. The nominal exchange rate is obtained from the International Monetary Fund's *International Financial Statistics*.

The classification of traded and non-traded goods in this study follows the spirit of Esaka (2003). For example, foods and apparel, are regarded as traded goods; in contrast, other sectors, such as housing, education, medical care, transportation and communication, fuel, light and water charges are regarded as non-traded goods. This classification is similar to Boivin et al. (2009) and Nagayasu and Inakura (2009). After carefully matching, we obtain prices of 304 items, including 204 traded goods & 80 non-traded goods.

3 Empirical Results

In this section, we show the cross-sectional mean of $\hat{\rho}_i$'s from various estimation methods, including the FE estimator, GMM estimator, mean group (MG) estimator of Pesaran and Smith (1995) and CCEMG estimator of Pesaran (2006), and the associated half-life estimates in Table 1.⁴ Tables 1-2 report the results of estimated AR coefficients and half-lives without and with the small sample bias correction by distinct ways, respectively.⁵

Consider the results without small sample bias correction first. Notice that the MG and CCEMG allow for heterogeneous coefficients, while the parameter homogeneity is imposed on the FE and GMM. If there is upward aggregation bias, the FE or GMM will result in higher estimated half-lives than those of the MG or CCEMG. The estimated half-lives obtained

⁴In our empirical results, we let $\hat{\rho}_i$ are equal to 0.995 if the estimates above 0.995. The half-life point estimates are the average of individual series and their standard errors are calculated by the *delta* method.

⁵Panel unit root tests without and with a time trend are all rejected the hypothesis that q_{it} 's have unit roots.

by the FE and GMM for whole sample are 28.39 and 8.20 years, respectively. Similarly, the confidence intervals are also quite large, (21.52, 35.26) for FE and (6.73, 9.66) for GMM. However, once slope heterogeneity is allowed by the MG in AR(1) regressions, the estimated half-life declines dramatically to 5.04 years. This result implies that the dynamic processes of Japan-U.S. RERs across distinct products are more likely to be heterogeneous.

[Insert Tables 1-2 about here]

Of interest, we compare the results from the MG estimation with different specification in the AR orders. Obviously, the average estimated half-lives obtained from the AR(1) are considerably less than those from the AR(κ_i).⁶ Our findings are consistent with the empirical results of Choi, Mark and Sul (2006) and Murray and Papell (2005) and support the arguments of Rossi (2005) that the half-lives tend to be downward biased when the true DGP is not an AR(1) process but estimated by an AR(1) model. Due to the potential heterogeneity in κ_i , hereafter, we will focus on the results obtained from the AR(κ_i) model only.

After controlling for the cross-sectional dependence, we find that the average half-life for whole products, incorporated traded and non-traded goods, is about 2.54 years with 95% confidence intervals between 2.19 and 2.89 years, which are considerably less than the average half-lives obtained by other methods and indicates that ignoring cross-sectional common effects may notably distort the half-life calculation.

As pointed by Chen and Engel (2005), the effect of small-sample bias on half-life estimation may be severe. We use the bootstrap procedure to correct the possible small-sample bias. Due to the heterogeneity in coefficients and the numbers of AR lagged order, we therefore only compare the bias-corrected MG and CCEMG estimates with optimal lag length (κ_i) in Table 2. The estimated half-lives by the MG and CCEMG for whole samples are 8.20

⁶We also consider the model with $\kappa_i = 12$ for all i 's. However, the results from AR(12) and from AR(κ_i) are similar. We therefore remove the results of AR(12) for simplicity. A complete description of the results is available on request from the authors.

and 3.58 years. They both are higher than those obtained from the same methods without small-sample bias corrections, but these don't change the main conclusions described above. In addition, the estimated half-lives of Japan-U.S. RERs by MG and CCEMG without/with small-sample bias corrections are considerably higher than those for Mexican-U.S. RERs (see, Robertson et al., 2009). The finding supports that higher transportation costs with international trade for Japan-U.S. rather than that for Mexico-U.S. lead to higher disaggregated RER persistence. Furthermore, the estimated half-life for all goods by CCEMG is the same as the remarkable consensus view of Rogoff (1996).

The cross-sectional dependence test findings are reported in columns 2 and 3. They are proposed by Pesaran (2006) and Frees (1995), respectively. The both results reject the null hypothesis of zero cross-sectional heterogeneity at the 1% significant level for models with whole categories. These findings show the importance for taking CCE into account and reaffirm why the methodology we adopt CCEMG.

Importantly, the bias-corrected half-lives by CCEMG, see Table 2, for non-traded goods is about 3.66 only about 0.11 year higher than that for traded goods. Similar to the findings of Crucini and Shintani (2008) and Parsley and Wei (2007), the estimated speed of mean reversion (or, equivalently, deviations from PPP) for non-traded goods appear to be more persistent than that for traded goods. However, difference of the estimated half-lives between traded and non-traded goods are not significant.

The high distribution costs for the retailed prices of consumption goods might interpret the phenomenon that the insignificant convergence gap between the traded and non-traded goods. The distribution services are an important component for many traded goods, while these inputs are regarded as the non-traded. The costs may create a natural wedge of the traded goods prices between Japan and the U.S. (see, Burstein et al., 2003). Another reason is that the mean of half-lives cannot fully describe the behaviors of various traded and non-traded goods. Below we will further investigate the difference between the distributions of the traded and non-traded goods.

[Insert Table 3 about here]

Figure 1 plots the kernel-based probability density estimates of the AR coefficients ρ_i for the traded and non-traded goods, respectively. The vertical lines present the mean of the RERs persistence. Although the traded and non-traded goods appear to have similar average convergence rates to PPP, their distributions are quite different. It is obvious that the distribution of non-traded goods has a higher peak than that of the traded goods, which indicates that the non-traded goods are less dispersive than the traded goods. Furthermore, we apply the nonparametric tests to re-examine whether these two categories' population distributions are identical or not. The results from the Wilcoxon rank sum test, the Cramér-von-Mises test and the Kolmogorov-Smirnov two-sample test are summarized in Table 4 and all of them indicate that the null hypothesis of same population is rejected.⁷ This result confirms the similar findings of Crucini and Shintani (2008) and Parsley and Wei (2007) that traded goods tend to exhibit shorter half-lives than non-traded goods with marginal difference.

[Insert Table 4 and Figure 1 about here]

4 Concluding Remarks

The study has examined the product level CPI-based Japanese RER (the U.S. served as the numéraire) for investigating the convergence rate of deviations from PPP. The estimated half-lives of disaggregated RER, for whole goods via CCEMG (MG) estimation take about 2.54(6.06) years for mean reversion. The estimated half-lives are similar to the consensus half-lives of 3 - 5 years (e.g., Rogoff, 1996). After correcting the small-sample bias, the estimated half-lives of disaggregated RER via CCEMG (MG) increase by 40%(35%). The

⁷Anderson (1962) shows that in the case of the two sample test, the Cramér-von-Mises test is more powerful than the Kolmogorov-Smirnov two-sample test. It is a two-sided rank sum statistic used to test the null that data are independent samples from identical continuous distributions with equal medians, against the alternative that they do not have equal medians.

correction of small-sample bias bears crucial but does not change our main conclusions (e.g., Chen and Engel, 2005). The estimated half-lives of Japanese RER for whole goods are higher than those results for Mexico (e.g., Robertson et al., 2009) even though examine the similar data level and by similar approaches. A plausible reason is that trading costs of Japan-U.S. are higher than Mexico-U.S. potentially due to distance, free trade agreements, etc.

In addition, after controlling the CCE and correcting the small sample bias, the estimated half-life for non-traded goods deviations from PPP is about 3.66 years, about 0.11 year slightly higher than that for traded goods. The possible reason is that the prices of traded goods are heavily accounted for by distribution costs. These costs are intensive in labor or land which are non-traded good. They create a natural wedge between prices of traded goods in Japan-U.S. (see, for example, Burstein et al., 2003). Of interest, the AR coefficients of disaggregated RER for traded goods perform distinct distribution of that for non-traded goods. The results are also imply that the average half-life of traded goods is less than that of non-traded goods. Our results support the findings of Crucini and Shintani (2008) and Parsley and Wei (2007).

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Table 1: Persistence estimations without bias correction

	ρ	95%C.I. of ρ	$\bar{\tau}_{HL}$	95%C.I. of τ_{HL}
Fixed Effects				
All	0.998 (0.000)	[0.997, 0.998]	28.39 (3.506)	[21.52, 35.26]
Generalized Method of Moments				
All	0.993 (0.000)	[0.992, 0.994]	8.195 (0.748)	[6.729, 9.662]
Mean Group with AR(1)				
All	0.954 (0.003)	[0.948, 0.960]	5.043 (0.249)	[4.555, 5.531]
Mean Group with AR(κ_i)				
All	0.973 (0.003)	[0.967, 0.979]	6.063 (0.253)	[5.567, 6.559]
Traded	0.969 (0.004)	[0.961, 0.977]	5.787 (0.300)	[5.199, 6.375]
Non-traded	0.984 (0.002)	[0.980, 0.988]	6.835 (0.459)	[5.935, 7.735]
Common Correlated Effects Mean Group with AR(κ_i)				
All	0.935 (0.004)	[0.927, 0.943]	2.543 (0.179)	[2.192, 2.894]
Traded	0.926 (0.006)	[0.914, 0.938]	2.522 (0.216)	[2.099, 2.945]
Non-traded	0.959 (0.003)	[0.953, 0.965]	2.602 (0.307)	[2.000, 3.204]

¹ The average half-life, $\bar{\tau}_{HL}$, is defined as the expected years declined by half for the PPP deviations and measured as $\ln(0.5)/\ln(\rho)$. The 95% confidence interval (C.I.) for half-lives is based on the *delta* method approximation and places in square brackets.

² The reported numbers in parentheses are standard errors.

Table 2: Persistence estimations with bias correction

	ρ	95%C.I. of ρ	$\bar{\tau}_{HL}$	95%C.I. of τ_{HL}
Mean Group with AR(κ_i)				
All	0.980 (0.003)	[0.974, 0.986]	8.197 (0.244)	[7.719, 8.675]
Traded	0.977 (0.003)	[0.971, 0.983]	7.959 (0.296)	[7.379, 8.539]
Non-traded	0.989 (0.002)	[0.985, 0.993]	8.863 (0.412)	[8.055, 9.671]
Common Correlated Effects Mean Group with AR(κ_i)				
All	0.943 (0.004)	[0.935, 0.951]	3.575 (0.224)	[3.136, 4.014]
Traded	0.935 (0.005)	[0.925, 0.945]	3.547 (0.268)	[3.022, 4.072]
Non-traded	0.966 (0.003)	[0.960, 0.972]	3.656 (0.406)	[2.860, 4.452]

¹ The average half-life, $\bar{\tau}_{HL}$, is defined as the expected years declined by half for the PPP deviations and measured as $\ln(0.5)/\ln(\rho)$. The 95% confidence interval (C.I.) for half-lives is based on the *delta* method approximation and places in square brackets.

² The reported numbers in parentheses are standard errors.

Table 3: Specification tests

	Fixed effects v.s. random effects	Cross-Sectional independence	
	Hausman test	Pesaran's test	Frees' test
All	697.2 (0.000)	2319.0 (0.000)	143.4 (0.000)
Traded	657.7 (0.000)	1522.8 (0.000)	87.6 (0.000)
Non-traded	35.2 (0.000)	842.3 (0.000)	64.2 (0.000)

¹ The reported numbers in parentheses are p -value.

Table 4: Nonparametric tests

	Wilcoxon rank sum test	Cramér-von Mises test	Kolmogorov -Smirnov test
Mean Group with AR(κ_i)	1.172 (0.241)	0.242 (0.200)	0.182 (0.035)
Common Correlated Effects Mean Group with AR(κ_i)	2.160 (0.031)	0.798 (0.009)	0.199 (0.016)

¹ H₀: Traded and Non-traded are drawn from the same underlying continuous population.

² The values in the parentheses denote the p -value.

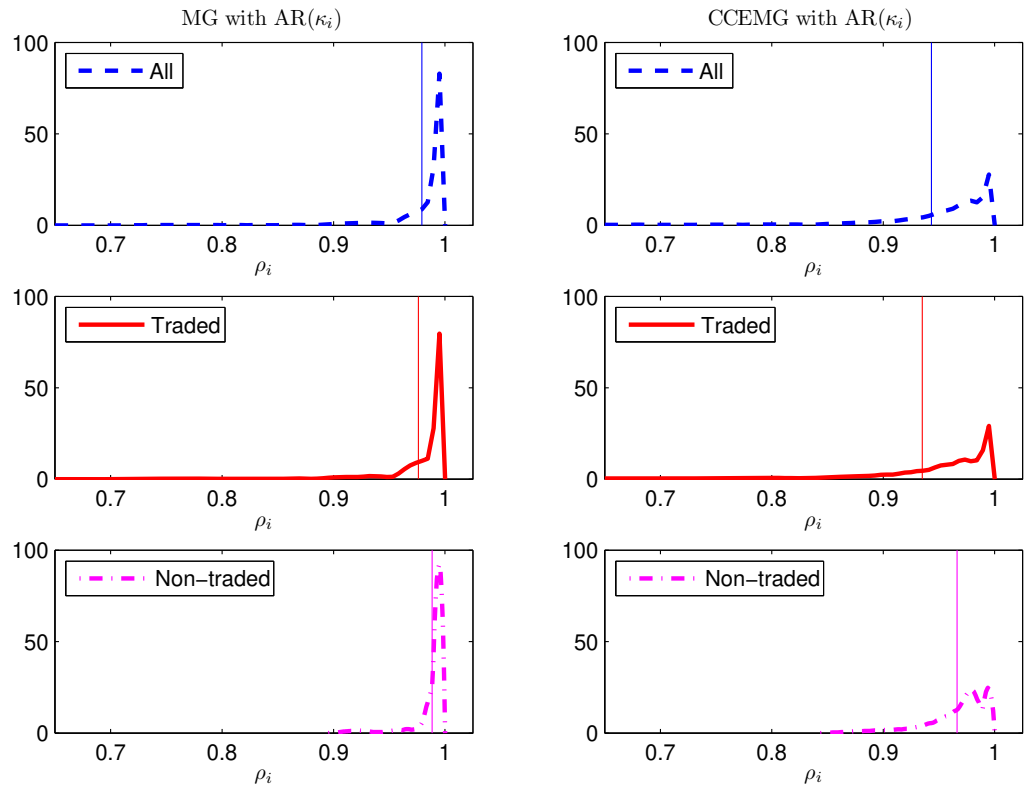


Figure 1: Kernel-based density estimates of the bias corrected MG and CCEMG estimates, $\hat{\rho}_i$, among consumption goods.