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Financial Market Integration and World Economic Stabilization toward Purchasing Power Parity

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Abstract

Purchasing power parity (PPP) is one of the most important, but empirically controversial theories in international macroeconomics. Although many researchers believe that some variant of PPP holds in the long-run, there are diverse empirical results regarding the PPP hypothesis. We examine the PPP hypothesis from an alternate point of view: we investigate the possibility of financial market integration, and world economic stabilization toward PPP, by examining the change in the persistence of PPP deviations during the last three decades. We employ a fractional integration framework, which provides a powerful tool to detect changes in the persistence for highly persistent time series. First, we test the null hypothesis of no decline in the persistence of PPP deviations. The test rejects the null at the 10% significance level for 11 out of 17 countries, thus providing strong support for financial market integration and world economic stabilization toward PPP. Second, we examine the dynamics of the persistence of PPP deviations during the last three decades through rolling-window estimation. Our results show that the persistence of PPP deviations has decreased gradually, and that many real exchange rates have experienced a sharp drop in their persistence once samples starting in the mid-1980s are used. Interestingly, this timing almost coincides with the timing of U.S./world economic stabilization reported by other studies. We also examine the relation between the persistence of PPP deviations and de facto measures of financial integration by Lane and Milesi-Ferretti (2006). We confirm that they are strongly correlated for all countries. This finding suggests that the recent promotion of financial integration is one of the main sources of the decline in the persistence of PPP deviations.

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Key Words: Fractional Integration; PPP; Real Exchange Rate; Financial Integration

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1 Introduction

Purchasing power parity (PPP) is one of the most important, but empirically controversial elements in international macroeconomics. PPP simply advocates that the equilibrium exchange rate of two currencies should equalize their purchasing power. The idea behind PPP is very intuitive: once converted to a common currency, national price levels should be equal. Although many researchers believe that some variant of PPP holds in the long-run, there are diverse empirical results regarding the PPP hypothesis, in particular for the recent floating rate period.

In this paper, we examine the PPP hypothesis from a different point of view than previous studies. Specifically, we investigate the possibility of financial market integration and world economic stabilization toward PPP by testing the null hypothesis of no decline in the persistence of PPP deviations in the last 30 years. Furthermore, we examine the dynamics of the persistence of PPP deviations during the last three decades. To our best knowledge, none of the previous research investigates changes in the persistence of real exchange rates systematically. There are, however, several interests to examine the dynamics of the persistence of PPP deviations. The first relates to financial market integration. According to IMF's *Annual Report on Exchange Arrangements and Exchange Restrictions* (AREAER), many industrial countries experienced a rapid increase in the degree of financial openness since mid-1980s.[1] Likewise, the de facto measures recently constructed by Lane and Milesi-Ferretti (2006) indicate that financial integration in industrial countries has promoted gradually in 1970s and 1980s, and accelerated in mid-1990s.[2] From these observations, we can imagine that PPP should hold more naturally in recent periods. It is, therefore, instructive to examine whether we can find a stronger evidence of PPP in more recent integrated real exchange rates.

Another interest comes from the U.S. and world economic stabilization. Following Kim and Nelson (1999) and McConnell and Perez-Quiros (2000), who point out a sharp decline in the variance of the U.S. economic growth rate in the mid-1980s, several studies provide evidence of commensurate changes toward U.S./world economic stabilizations. For instance, Clarida, Galí, Gertler (2000) estimate a forward-looking monetary policy function, and show that the U.S. monetary policy has been more stabilizing after 1980. Stock and Watson (2002) and Sensier and van Dijk (2004) find declines in the volatility in a number of U.S. economic time series around mid-1980s, including series such as employment growth, consumption growth, wage, and price inflation. Following these studies, Kim, Nelson, and Piger (2004) and Herrera and Pesavento (2005) provide further supports for the U.S. economic stabilization by identifying

[1] AREAER reports a set of *de jure* measures of legal restrictions on cross-border capital flows, and is widely used to measure financial openness.

[2] See Kose et al. (2006) for details of financial integration and related measures.

possible explanations for the reduction of the variance in U.S. GDP growth. Regarding world economic stabilization, Stock and Watson (2005) find a reduction in the magnitude of the common international shocks contributing to a substantial moderation in the volatility of the GDP growth rates over the past 40 years in the G7 countries (except for Japan). In addition, recent literature finds a corresponding decline in inflation persistence in the U.S. and other industrial countries. For instance, Kumar and Okimoto (2007) find a marked decline in the U.S. inflation persistence around the early 1980s.[3] Furthermore, they find similar declines in the inflation persistence of other G7 countries, except for Italy, suggesting the possibility of world economic stabilization. A natural question raised from these studies is whether we can observe commensurate changes toward world economic stabilization for other economic variables. This paper provides an answer to this question for real exchange rates, or deviations from PPP. If there is a decline in the persistence of real exchange rates, as we will show in this paper, this indicates a new evidence of world economic stabilization toward PPP.

The null hypothesis to be investigated formally in this paper is that there has been no significant decline in the persistence of deviations from PPP over the past three decades for industrial countries. This hypothesis is tested against the alternative that there has indeed been a marked and sustained decline in the persistence of PPP deviations. To this end, we employ a fractional integration framework, which provides a powerful tool to detect changes in the persistence for highly persistent time series, here real exchange rates. In the fractional integration framework, our null hypothesis is formulated as no change in the order of fractional integration, d, and alternative as a decline in d. This paper conducts two analyses to examine this hypothesis for major industrial countries using U.S. dollar-based real exchange rates.

First, we conduct a formal statistical test of the null of no change in d using two 15-year subsamples. In this analysis, we do not try to specify the correct timing nor transition process of possible declines in PPP deviation persistence, since it is a most formidable task. Rather, we simply use the first and last 15 years of the data, and test the difference in d between the two subsamples. This may not be the most powerful way to detect a decline of persistence, since it does not specify the possible timing and type of structural changes. However, if there has been a significant decline, the test should detect it. In fact, this is the case. The tests of the null hypothesis $d_1 = d_2$ against the alternative $d_1 > d_2$, where d_1 and d_2 are orders of integration of real exchange rates for the first and second subsample, rejects the null at the 10% significance level for 11 out of 17 countries. In particular, for the G7 countries, we successfully reject the null for 5 out of 6 countries. This result provides strong support for financial market integration and world economic stabilization toward PPP in recent years.

[3]See Taylor (2000) and Cogley and Sargent (2001, 2005) for other studies which find a similar decline in U.S. inflation persistence.

Second, we employ 15-year rolling-window estimation to examine the dynamics of persistence of real exchange rates.[4] This rolling window estimation is simple, but can still provide very useful information regarding the timing and transition process of declines in the persistence of PPP deviations. Our 15-year rolling window results indicate that many real exchange rates have experienced a sharp drop in persistence once samples starting mid-1980s are used. Interestingly, this timing almost coincides with the timing of U.S./world economic stabilization reported by other studies including Kim and Nelson (1999) and Kumar and Okimoto (2007). Hence, this result provides further support of the hypothesis of world economic stabilization. We also examine the relation between the persistence of PPP deviations and de facto measures of financial integration by Lane and Milesi-Ferretti (2006), and confirm that they are strongly correlated for all countries. This finding suggests that the recent promotion of financial integration is one of the main sources of the decline of the persistence in PPP deviations.

The rest of the paper is organized as follows. Section 2 reviews the related literature. Section 3 discusses the methodology we use to obtain estimates for order of fractional integration, or a measure of persistence. Section 4 presents our main estimation results and their implications. Section 5 concludes.

2 Review of the related literature

The notion of PPP has attracted great attention among theorists, empirical researchers, and policy makers, since most of industrial countries adopted a flexible exchange rate in the early 1970s. One consequence for this attention is that there is an enormous empirical literature on PPP. In this section, we review related research, and clarify our contribution.

Most empirical studies employ unit root tests or cointegration analysis, and fail to find evidence in favor of PPP. For instance, Patel (1990) conducts cointegration tests between producer price indices and exchange rates for several countries over the flexible exchange rate period, and finds no strong evidence supporting PPP. Among others, Corbae and Ouliaris (1988), Enders (1988), Meese and Rogoff (1988), Mark (1990), and Edison and Pauls (1993) perform analogous analyses, and reach similar conclusions.

Many researchers, however, consider that those negative results obtained in previous research reflect poor performance of the econometric methodologies rather than evidence against PPP. In particular, the low power of unit root and cointegration tests has been often pointed out. For instance, Hakkio (1986) provides a simulation study to show that unit root tests often fail to reject the null hypothesis of unit root if the real exchange rate has a near unit root. To overcome this problem, several approaches have been developed. The first approach uses a longer time horizon. Since PPP is a long-run equilibrium concept, it is expected that PPP tends to hold more naturally over a longer time period. Several studies find

[4] Other studies which use rolling-window estimation include Stock (2001) and Kumar and Okimoto (2007).

stronger evidence for PPP using this more stable relationship over a longer time period. Those examples include Abuaf and Jorion (1990), Kim (1990), Ardeni and Lubian (1991), Glen (1992), and Lothian and Taylor (1996). As indicated by Engel (2000), however, using longer-span data may not completely solve the problems associated with testing PPP. In addition, it is questionable whether the exact same PPP relationship holds in such a long period. Even if PPP remains true for the entire period, the convergence speed of PPP deviations can be very different under different exchange rate regimes, such as the Bretton Woods era and the flexible exchange rate period. Furthermore, the degree of financial market integration may affect the persistence of PPP deviations.

An alternative approach employs panel unit root tests to improve the power of standard unit root tests. The panel data approach has an advantage over the long-span data approach in that it can be useful in testing PPP in the recent floating rate period. Although the PPP hypothesis is of interest in any extent, it is more instructive to examine whether it holds under the recent flexible exchange rate system. Along this vein, studies such as Wei and Parsley (1995), Oh (1996), Wu (1996), and Papell (1997) apply panel unit root tests to real exchange rate data of several countries in the flexible exchange rate period, and find evidence in favor of PPP. One concern with these panel studies is their ignorance of cross-sectional dependence. Indeed, O'Connell (1988) finds no evidence of PPP once cross-sectional dependence is controlled. Another concern is their use of the null of joint nonstationarity. It is possible that joint nonstationarity of a group of real exchange rates may be rejected when only one of these series is stationary, as indicated by Taylor and Sarno (1998). Thus, it is hard to say that these results from panel unit root tests demonstrate strong evidence of PPP.

Another approach that has been considered is the fractional integration approach, which extends the standard unit root framework. Offering a generalization of the classical dichotomy between $I(0)$ and $I(1)$ processes, fractionally integrated processes can provide a more powerful framework to detect mean reversion than the standard unit root tests. Diebold, Husted, and Rush (1991) and Cheung and Lai (1993) find evidence of long-memory, but mean reversion, in long historical series of real exchange rates, while Cheung and Lai (2001) and Achy (2003) find similar results in the recent floating rate period. On the other hand, Baum, Barkoulas, and Caglayan (1999) fail to reject the unit root hypothesis against the fractional integration alternative for the post-Bretton Woods era.

In sum, there is growing evidence supporting PPP, but the evidence is not sufficient to conclude that PPP holds. In particular, none of above studies considers the possibility of a movement toward PPP in recent periods, even though there are several reasons to expect such a trend, as emphasized in the introduction. It is, therefore, worth investigating whether we can find empirical evidence for this possibility, which is the main purpose of this paper. To this end, we semiparametrically estimate the order of fractional integration in real exchange rate, and use it to measure the persistence of real exchange rate. This framework has several advantages over the standard unit root tests or ARFIMA framework, as discussed in detail in the next section.

3 Methodology

3.1 Fractionally integrated ($I(d)$) processes

In this paper, we propose to use the order of (fractional) integration to assess the persistence of real exchange rate. Fractionally integrated ($I(d)$) processes encompass both short-memory ($I(0)$) and unit root ($I(1)$) processes as limiting cases when the order of integration, d, takes on the values zero and unity. They can accommodate temporal dependence that is intermediate in form between an $I(0)$ and an $I(1)$ process. As such, $I(d)$ processes provide a more flexible way to model long-run dynamics than $I(0)$ and $I(1)$ processes, giving some liberation from the $I(0)/I(1)$ dichotomy.

A process X_t is said to be an $I(d)$ process if its fractional difference, $(1-L)^d X_t$, is an $I(0)$ process. The fractional difference operator $(1-L)^d$ is defined by means of the gamma function

$$(1-L)^d = \sum_{k=0}^{\infty} \frac{\Gamma(k-d)L^k}{\Gamma(-d)\Gamma(k+1)},$$

where the parameter d is allowed to take any real value. When d is a nonnegative integer, the infinite-order summation terminates, giving the standard integrated processes. An $I(d)$ process is stationary and invertible when $-\frac{1}{2} < d < \frac{1}{2}$. An $I(d)$ process with $d \geq 1/2$ is nonstationary, but is still mean reverting if $1/2 \leq d < 1$. Importantly, an $I(d)$ process with $0 < d < 1$ can accommodate slowly decaying autocorrelations (when stationary) and slowly decaying impulse response function that are inconsistent with either an $I(0)$ or an $I(1)$ process. Specifically, the impulse response function of an $I(0)$ process decays exponentially, while the impulse response function of an $I(1)$ process approaches to a positive constant in the long run and never dies out. In contrast, the impulse response function of an $I(d)$ process with $0 < d < 1$ decays at a slow hyperbolic rate, k^{d-1}. The order of integration, d, determines the decay rate of autocorrelations and the impulse response function.

The long-run dynamics of an $I(d)$ process is governed by the parameter d. Using the value of d as a measure of persistence has several attractive features. First, $I(d)$ processes allow us to model persistence that is not consistent with either an $I(1)$ process or an $I(0)$ process. Empirical evidence suggests that the deviation from PPP is very persistent. On the one hand, an $I(1)$ process is not acceptable as a model of real exchange rate in light of the theory of PPP. On the other hand, using an $I(0)$ process to model the real exchange rate forces it to have exponentially decaying impulse response function, for which there is little underlying economic justification. Second, the integration parameter d has little to do with the short-run dynamics of the data. The largest autoregressive root, which is commonly used as a measure of long-run dynamics, is intimately related with the first-order autocorrelation of the data when the root is close to unity. As such, it is affected by both short and long-run dynamics. Third, the integration parameter, d, can be estimated consistently from the data. One popular way of modeling a highly persistent data is a local-to-unity model

$$X_t = \left(1 + \frac{c}{T} L\right) X_{t-1} + u_t, \quad u_t \sim I(0), \quad t = 1, \ldots, T,$$

with an initialization of X_0. In this model, the long-run dynamics of X_t is summarized by c, but one cannot obtain a point estimate of c.

3.2 Estimation of order of fractional integration

The order of integration, d, plays a central role in the definition of fractionally integrated processes, and has often been the focus of previous studies. We use the 2-step feasible exact local Whittle (FELW) estimator by Shimotsu (2006) that extends the exact local Whittle (ELW) estimator by Shimotsu and Phillips (2005). The FELW estimator is a semiparametric estimator, which is agnostic about, and robust to misspecification of, the short-run dynamics of the process. This feature is attractive for our paper, because our interest is in the long-run dynamics of real exchange rate, and we want to impose as little assumptions as possible on the short-run dynamics. Another useful feature of the FELW estimator is that it accommodates both stationary ($d < 1/2$) and nonstationary ($d \geq 1/2$) fractionally integrated processes. We do not want to impose *a priori* restrictions on whether $d \gtrless 1/2$, because the theory of PPP itself implies no restriction on the value of d.

The ELW estimator assumes that the fractionally integrated process X_t is generated by the model

$$(1 - L)^d X_t = u_t \mathbf{1}\{t \geq 1\}, \quad t = 0, \pm 1, \ldots \tag{1}$$

where $\mathbf{1}\{\cdot\}$ denotes the indicator function. The error, u_t, is a mean-zero $I(0)$ process with spectral density $f_u(\lambda)$ satisfying $f_u(\lambda) \sim G$ for $\lambda \sim 0$. We model $u_t = \sum_{j=1}^{\infty} c_j \varepsilon_{t-j}$ where ε_t has mean zero and variance 1, and is serially uncorrelated. Inverting and expanding the binomial in (1) gives a representation of X_t in terms of $u_1, \ldots, u_n$:

$$X_t = (1 - L)^{-d} u_t \mathbf{1}\{t \geq 1\} = \sum_{k=0}^{t-1} \frac{\Gamma(d+k)}{\Gamma(d)\Gamma(k+1)} u_{t-k}, \quad t = 0, \pm 1, \ldots$$

This model accommodates both $d < 1/2$ and $d \geq 1/2$ because of the initialization at $t = 0$ induced by $\mathbf{1}\{t \geq 1\}$. Define the discrete Fourier transform (dft) and the periodogram of a time series a_t, $t = 1, \ldots, n$, evaluated at the fundamental frequencies as

$$\begin{aligned} w_a(\lambda_j) &= (2\pi n)^{-1/2} \sum_{t=1}^{n} a_t e^{it\lambda_j}, \quad \lambda_j = \frac{2\pi j}{n}, \quad j = 1, \ldots, n, \\ I_a(\lambda_j) &= |w_a(\lambda_j)|^2. \end{aligned} \tag{2}$$

Shimotsu and Phillips (2005) propose to estimate (d, G) by minimizing the objective function

$$Q_m(G, d) = \frac{1}{m} \sum_{j=1}^{m} \left[\log \left(G \lambda_j^{-2d} \right) + \frac{1}{G} I_{(1-L)^d x}(\lambda_j) \right], \tag{3}$$

where n is the sample size, and m is the bandwidth that satisfies $m \to \infty$ and $m/n \to 0$. $Q_m(G, d)$ is derived from the (negative) Whittle likelihood function of X_t localized to the neighborhood of the origin. The assumption $m/n \to 0$ localizes the likelihood function to the neighborhood to the origin. The frequencies in the neighborhood of the origin correspond to the long-run dynamics of the data, and this localization makes the estimator agnostic to the short-run dynamics of the data.

Concentrating $Q_m(G, d)$ with respect to G, Shimotsu and Phillips (2005) define the ELW estimator as

$$\hat{d} = \underset{d \in [\Delta_1, \Delta_2]}{\arg\min}\, R(d), \tag{4}$$

where Δ_1 and Δ_2 are the lower and upper bounds of the admissible values of d and

$$R(d) = \log \widehat{G}(d) - 2d \frac{1}{m} \sum_{j=1}^{m} \log \lambda_j, \quad \widehat{G}(d) = \frac{1}{m} \sum_{j=1}^{m} I_{(1-L)^d x}(\lambda_j).$$

In what follows, we distinguish the true value of d and G by d_0 and G_0.

Shimotsu and Phillips (2005) show that, under some conditions, in particular, for $d_0 \in (\Delta_1, \Delta_2)$ with $\Delta_2 - \Delta_1 \leq \frac{9}{2}$,

$$m^{1/2}(\hat{d} - d_0) \to_d N\left(0, \frac{1}{4}\right), \quad \text{as } n \to \infty,$$

where m is chosen so that $1/m + m^{1+2\beta}(\log m)^2 n^{-2\beta} + m^{-\gamma} \log n \to 0$ for any $\gamma > 0$. Here β represents the degree of approximation of the spectral density of u, $f_u(\lambda)$, around the origin by G.

Shimotsu (2006) develops the 2-step feasible ELW (FELW) estimator that extends the ELW estimator to accommodate an unknown mean, so that the model that generates the data is

$$X_t = (1-L)^{-d} u_t \mathbf{1}\{t \geq 1\} + \mu_0, \quad t = 0, \pm 1, \ldots \tag{5}$$

The FELW estimator estimates the unknown mean, μ_0, by a weighted average of the sample mean, $\overline{X} = n^{-1} \sum_{t=1}^{n} X_t$ and the initial observation X_1 :

$$\hat{\mu}(d) = w(d)\overline{X} + (1 - w(d))X_1,$$

where $w(d)$ is a smooth twice continuously differentiable weight function, such that $w(d) = 1$ for $d \leq \frac{1}{2}$ and $w(d) = 0$ for $d \geq \frac{3}{4}$.[5] Shimotsu (2006) shows that the FELW estimator has the same asymptotic distribution as the ELW estimator.

The value of m is chosen by the researcher. The choice involves a bias-variance tradeoff; using a too small m increases the variance of the estimator, while using a too large m induces bias in estimation because of the effect from short-run dynamics. The value of β is known to be 2 for many probable models of u_t. Hence, the largest possible choice of m is slightly smaller than $n^{4/5}$. In practice, more conservative choices such as $m^{0.65}$ or $m^{0.75}$ are often used.

4 Empirical Analysis

We use monthly U.S. dollar-based real exchange rates for 17 industrial countries with the sample period from January 1974 to December 2006. The data are collected from IMF's International Financial Statistics (IFS). We use the CPI (IFS line 64) as the measure of prices, and the end-of-period domestic currency units per U.S. dollar (IFS line ae) as the exchange rate. We follow Papell (1997) in selecting countries whose exchange rate is examined. These countries consists of those classified as industrial by the IMF not including Australia, Iceland, Ireland, Luxembourg, and New Zealand. We exclude Luxembourg because it maintained a currency union with Belgium. Australia, Iceland, Ireland, and New Zealand do not have monthly CPI data for the entire sample period. For Euro-countries, their exchange rate after 1997 is calculated from the U.S. dollar-Euro exchange rate and the conversion rate between Euro and each national currency.

We hypothesize that there has been no significant decline in the persistence of these real exchange rates, or deviations from PPP, over the past three decades. We use two methods; the first compares two 15-year subsamples, and the second is 15-year rolling-window estimation. While the former provides us a way to test the hypothesis statistically, the later allows us to examine the dynamics of the persistence of PPP deviations more informatively. In what follows, we first present the estimates of d from the whole sample to justify the use of the fractional integration framework, and then discuss the outcome of the two analysis.

4.1 Whole sample analysis

For the first analysis, we report the estimates of the orders of fractional integration for real exchange rates, or deviations from PPP, for 17 industrial countries using the whole sample. Throughout this subsection, we do not consider the possibility of changes in the persistence of real exchange rates. This is because

[5] See Shimotsu (2006) for technical details.

we want to confirm that the order of fractional integration is a suitable measure of persistence before conducting formal tests of declines in the persistence of PPP deviations. The results from the whole sample analysis support the nonstationarity of PPP deviations, and give us a solid reason to use the fractional integration framework to detect declines in their persistence.

The second column of Table 1 reports the FELW estimates of the orders of fractional integration for real exchange rates. We set the bandwidth to $m \approx n^{0.65}$, namely $m = 48$ for this analysis and $m = 29$ for the 15-year subsample analysis, respectively. The asymptotic standard error of each estimate is $1/\sqrt{4 \times 48} = 0.072$, and the asymptotic 95% confidence interval is shown in the third column of Table 1. As can be seen, all estimates are close to one. From the 95% confidence interval, we reject the stationarity hypothesis, i.e. $d < 1/2$, at the 5% significance level. The p-value for the tests of the hypothesis $d < 1/2$ (not reported here) is smaller than 0.1%, providing strong evidence of nonstationarity of PPP deviations for all countries. The fourth column of Table 1 reports the Phillips-Perron Z_t-statistic for the null hypothesis that each real exchange rate has a unit root. The lag length is chosen to be 10. The 5% and 10% critical values of the Z_t-statistic is -2.874 and -2.570, respectively. Corroborating most previous studies, we cannot reject the null of unit root for any of the real exchange rate series at the usual significance level, indicating the nonstationarity of PPP deviations.

The fifth column of Table 1 reports the 95% confidence interval of the half-life of deviations from PPP. These intervals are computed from the 95% confidence interval of d using the relation

$$\frac{\partial X_{t+k}}{\partial u_t} \sim \frac{k^{d-1}}{\Gamma(d)} \quad \text{as } k \to \infty.$$

Since all the 95% upper bounds of d are larger than one, the 95% upper bound of the PPP deviation half-life is infinity for all countries. This finding is consistent with the conclusion from the previous studies such as Murray and Papell (2002) and Rossi (2005): the data are not sufficiently informative to pin down the half-life. The lower bound of the half-life is larger than the typical estimates based on Dickey-Fuller type regressions (Murray and Papell, 2002, Rossi, 2005). This is due to the shape of the impulse response function of fractionally integrated models. The impulse response function of the autoregressive model has an exponential decay, whereas that of the fractionally integrated model has a geometric decay. Consequently, fractionally integrated models produce larger half-life estimates, in particular when it involves long-run dynamics.

The half-life is not an informative measure to investigate changes in persistence, since an unbounded confidence interval does not allow us to conduct formal hypothesis tests of changes in persistence. To the contrary, the confidence intervals of the order of fractional integration are sufficiently tight, and we can use d as a measure of persistence to test the null hypothesis of no decline in the persistence of PPP deviations.

We also estimate d using the local Whittle estimator (Robinson, 1995) to check the robustness of our results. Note that the differenced series of an $I(d)$ process is $I(d-1)$. These estimates are calculated as follows. First, we take the difference of a real exchange rate series. Then, we estimate the order of integration of the differenced series by the local Whittle estimator. Finally, we add one to the estimate to get the estimate of d of the original series. Since the local Whittle estimator has a normal asymptotic distribution only when $-1/2 < d < 3/4$, this procedure implicitly assumes $d-1$ is larger than $-1/2$, namely $d > 1/2$. The last column of Table 1 reports the estimates. Not surprisingly, the estimates are very close to the FELW estimates based on the original series.

To sum, the results of the whole sample analysis clearly indicate the nonstationarity of real exchange rates and the usefulness of fractional integration framework to detect possible declines in the persistence of PPP deviations. Given these results we examine the possibility of financial market integration and world economic stabilization toward PPP in the next subsections.

4.2 Results of subsample analysis

In this section, we conduct formal statistical tests using two 15-year subsamples. The first subsample starts from January 1974, and ends in December 1988, while the second subsample is from January 1992 to December 2006. In this analysis, we do not pursue identifying the probable timing, nor the type of declines in the persistence of PPP deviations. Ideally, we can increase the power of the tests if we can correctly specify the timing and type of the transition process. However, it is very difficult to identify the type of structural changes, such as instantaneous breaks or gradual changes, and using a misspecified model may lead to erroneous conclusions. Therefore, we simply use two equally lengthed subsamples, but mitigate the problems associated with the timing, and type of structural changes, by making a three-year interval between the two subsamples. As a result, our tests are conservative in the sense that they may not detect declines in PPP deviation persistence most powerfully. If we can reject the null of no decline with these conservative tests, this constitutes strong evidence for financial market integration, and world economic stabilization toward PPP.

To conduct a formal test, we need to derive the joint distribution of the two estimates of the integration parameter from the two subsamples. This can be done as follows. Suppose $X_t, t = 1, \ldots, n_1$ is generated by model (5) with the integration parameter d_1, and suppose $X_t, t = n_2, \ldots, n$ with $n_1 < n_2$, is generated by model (5) with the integration parameter d_2. Let m be the bandwidth parameter, and let $\hat{d}_1$ and $\hat{d}_2$

be the FELW estimates of d from $X_t, t = 1, \ldots, n_1$ and $X_t, t = n_2, \ldots, n$, respectively. Then, we have

$$m^{1/2} \begin{pmatrix} \hat{d}_1 - d_1 \\ \hat{d}_2 - d_2 \end{pmatrix} \rightarrow_d N \left(0, \begin{bmatrix} 1/4 & 0 \\ 0 & 1/4 \end{bmatrix} \right), \quad \text{as } n \rightarrow \infty. \tag{6}$$

A sketch of the proof can be found in the Appendix. Thus, the two estimates of the integration parameter from the two subsamples are asymptotically independent. Based on this result, we can formally test the hypothesis that there has been no significant decline in the persistence of the deviations from PPP over the past three decades.

The first two columns of Table 2 report the FELW estimates of the orders of fractional integration (or persistence parameter values) of the U.S. dollar-based real exchange rate for each country and subsample. The asymptotic standard error of each estimate is $1/\sqrt{4 \times 29} = 0.093$. As can be seen, all the estimates for the first subsample are greater than 1, indicating highly persistent behavior of deviations from PPP. Since an $I(d)$ time series is not mean-reverting if $d \geq 1$, there is no indication of PPP in the first subsample. On the other hand, all the estimates from the second subsample are smaller than those from the first subsample. Furthermore, many of the estimates are less than 1. In particular, none of the estimates for the G7 countries are greater than 1, although they are not significantly different from 1.

These results suggest that deviations from PPP are less persistent in the second subsample, implying that PPP is more likely to hold in recent years. To examine this point more rigorously, we test the hypothesis that there is no decline in the persistence of the deviations from PPP. This amounts to testing the null hypothesis of $d_1 = d_2$ against the alternative hypothesis of $d_1 > d_2$, where d_1 and d_2 are orders of fractional integration of the first and second subsample, respectively. The fifth column of Table 2 shows the p-values of this test using the asymptotic distribution (6). The null hypothesis is rejected at the 5% significance level for Denmark, France, Japan, and Spain, and at the 10% level for Austria, Belgium, Canada, Germany, Italy, Netherlands, and Switzerland. Thus, the equality of the persistence of PPP deviations between two subsamples are rejected at the 10% significance level for 11 out of 17 countries, in particular, for 5 out of 6 for the G7 countries. For Finland, Greece, Norway, Portugal, and the United Kingdom, the estimates of d for the first subsample are relatively low, which is the main reason why the test cannot reject the null of no decline in PPP deviation persistence. These results provide strong evidence for the decline in the persistence of the deviations from PPP, implying the speed of adjustment to PPP has become faster in more recent years.

Note that our estimates fail to find mean reverting behavior in real exchange rates for some countries, even if we use the less persistent second subsample. This result is consistent with the previous studies using unit root tests, and supports the view that deviations from PPP are very persistent. Nevertheless, this does not necessarily imply that our evidence supporting PPP is weak. The following factors may contribute to this phenomenon. First, we have a small sample problem. Since our semi-parametric framework localizes the likelihood function to the neighborhood of the origin, the effective sample size is relatively small. As a consequence, the standard errors of the order of fractional integration estimates become somewhat large. Although they are sufficiently small to reject the null of $d_1 = d_2$ for many countries, the confidence intervals are not tight enough to reject the null of $d_2 = 1$. Another reason could be our ignorance of the nonlinear behavior in real exchange rates. As many studies suggest, the existence of transaction costs including transportation cost and trade barriers implies nonlinear real exchange rate adjustment toward PPP. Once this nonlinearity is considered, we can most likely find mean-reverting behavior in PPP deviations as Michael, Nobay, and Peel (1997) and Taylor, Peel, Sarno (2001) have previously found. Modeling the nonlinearity in semi-parametric fractional integration framework is beyond the scope of the present paper. Overall, our results indicate the failure of rejecting the unit root hypothesis, but provide solid evidence for the purpose of the paper: detecting the possibility of financial market integration, and world economic stabilization toward PPP.

We also examine the robustness of the results in Table 2 with respect to the bandwidth, m, by repeating the estimation for $m = 25, 27, 29, 31$, and 33. The first panel of Table 3 reports the difference of the estimates, $d_1 - d_2$. In general, the estimates are stable over the range of m, although the estimates exhibit some variability, and $d_1 - d_2$ tends to take small values when $m = 25$.

The second panel of Table 3 reports the asymptotic p-value of the one-sided test of the null hypothesis of $d_1 = d_2$ against the alternative of $d_1 > d_2$ for each bandwidth. Note that, for the same value of $\hat{d}_1 - \hat{d}_2$, the p-values are larger when m is smaller because of the larger asymptotic standard error estimate. For many countries, the p-values do not show large changes apart from an (inevitable) increase in its value for small m. The increase in the p-values is noticeable, in particular, for $m = 25$. For most countries, the conclusion with $m = 29$ remains valid for $m = 27, 31$, and 33.

In Table 2, we split the sample in the middle, each sample (1974-1988 and 1992-2006) having 180 observations. We examine how the results in Table 2 are affected by changing the point where the sample is split. Table 4 shows the estimates of $d_1 - d_2$ and their associated p-value when we change the break point. Considering the fact that the decline in the persistence and volatility in other macroeconomic variables occur in mid-1980s, we move the end of the first subsample between 1984 and 1988, but keep the interval between the two subsamples to three years. When we move the break point, the evidence in favor of the decline in d becomes stronger for some countries and weaker for other countries. Overall, the results in Table 4 are similar to Table 2, suggesting the decline in d in mid-1980s.

4.3 Rolling-window estimation

To obtain additional insight, and further support for our empirical findings of declines in the persistence of the deviations from PPP, we apply 15-year rolling-window estimation to the entire sample. First, we estimate the order of fractional integration, d, or the persistence parameter, using the first 15 years of the data (specifically, from January 1974 to December 1988). The data are then updated by 1 year increments, and d is re-estimated for the updated window (that is, for the period from January 1975 to December 1989). This procedure is repeated until the end of the sample period. Thus, the last estimate of d is based on the period from January 1992 to December 2006. The rolling-window estimation is easy to implement, and provides a significant amount of information about the underlying dynamics of the persistence of PPP deviations. In particular, this analysis can help highlight the periods over which there would likely have been a pronounced decline in the persistence of PPP deviations. Further, it gives useful observations about whether an instantaneous break, or a gradual change, better describes the transition process of d.

Figure 1 depicts the 15-year rolling-window estimates of the persistence parameter of the real exchange rate, along with the end year of the sample period, for the G7 countries. The figure shows remarkable similarities among the dynamics of the persistence of the G7 real exchange rates. For the first decade ending in 1998, the persistence of each real exchange rate decreased only slightly. Then, all the countries experienced a rapid decline in the persistence of PPP deviations between 1999 and 2002. Note that Figure 1 is drawn against the end year of estimated samples. In other words, the persistence of PPP deviations for the G7 countries declined notably once we start using samples starting mid-1980s. Interestingly, this period roughly coincides with previous studies' findings on the timing of a possible structural change toward stability in the U.S./world economy, such as Kim and Nelson (1999) and Kumar and Okimoto (2007). The persistence estimates for the G7 countries rebounded a little in 2003, and after that remained almost unchanged until 2006. The magnitude of the rebound for the UK is slightly larger than for other countries. This larger rebound, along with the initial low persistence of the UK real exchange rate, seems to be the reasons why the test based on two subsamples was insignificant for the UK. Figure 1, however, reveals that the decline in the persistence of PPP deviations for the UK is essentially the same as other G7 countries.

Figure 2 plots the 15-year rolling-window estimates for non-G7 countries, which have a significantly different PPP deviation persistence between two subsamples. The results are quite striking; all graphs behave practically same. In addition, they share analogous patterns with the G7 countries. In particular, all countries underwent sharp declines in PPP deviation persistence between 1999 and 2002. This timing coincides with the introduction of Euro, suggesting Euro may have played a significant role for the world financial integration toward PPP.

Figure 3 shows the 15-year rolling-window estimates for non-G7 countries whose differences in PPP deviation persistence between two subsamples are not significant. Despite the fact that the changes in the persistence of PPP deviations are less remarkable in these countries, their dynamics are still similar to those observed in Figures 1 and 2. In particular, the commensurate decline in PPP deviation persistence between 1999 and 2002 can also be seen in Figure 3, although the magnitude is not as striking as that of Figures 1 and 2.

As reported by Lane and Milesi-Ferretti (2006), financial integration in industrial countries has promoted gradually in 1970s and 1980s, and accelerated in mid-1990s. We examine the relation between financial integration and the persistence of deviations from PPP using two quantitative measures recently constructed by Lane and Milesi-Ferretti (2006), which are recommended by Kose et al. (2006). The first measure, IFIDGP, is the ratio of sum of gross stocks of foreign assets and liabilities to GDP:

$$IFIGDP = \frac{FA + FL}{GDP},$$

where FA (FL) denotes the stock of external assets (liabilities). The second measure, GEQGDP, focuses exclusively on portfolio equity and FDI holdings:

$$GEQGDP = \frac{PEQA + FDIA + PEQL + FDIL}{GDP},$$

where PEQA (PEQL) denotes the stock of portfolio equity assets (liabilities) and FDIA (FDIL) denotes the stock of direct investment assets (liabilities).

Table 5 reports the correlation between these two measures of financial integration and the rolling-window estimates for each country.[6] These measures and d are clearly negatively correlated, and the correlation coefficient is smaller than -0.5 in most cases. Although correlation does not necessarily imply causation, the results in Table 5 show that the decline in the persistence of PPP deviations occurred concurrent with the increase in financial integration.

5 Conclusions

Purchasing power parity (PPP) is one of the most important, but empirically controversial theories in international macroeconomics. A number of empirical studies regarding the PPP hypothesis have reached diverse results, and could not find decisive evidence, in particular for the recent floating rate period. In this paper, we examined the PPP hypothesis from a different point of view than previous studies to

[6]Both IFIGDP and GEQGDP are available only up to 2004.

provide new evidence supporting PPP. Specifically, this paper investigated the possibility of financial market integration and world economic stabilization toward PPP by testing the null hypothesis of no decline in the persistence of 17 industrial countries' U.S. dollar based real exchange rates in the last 30 years. To this end, we employed a fractional integration framework, and used the order of fractional integration as a measure of persistence of real exchange rates, or deviations from PPP. Confirming the appropriateness of our method by the whole sample analysis, we conducted formal statistical tests using two 15-year subsamples, and comparing estimates of the persistence parameter. We found marked and significant declines in PPP deviation persistence for 11 out of 17 countries. In particular, we rejected the null for 5 countries out of 6 for the G7 countries. These finding clearly indicate strong support for financial market integration and world economic stabilization toward PPP in recent years.

To obtain additional insight on declines in the persistence of the real exchange rates, the paper provided the dynamics of PPP deviation persistence by applying the 15-year rolling-window estimation. The results demonstrated remarkable similarities in dynamics of each real exchange rate's persistence. In particular, most countries experienced a rapid decline in the persistence of PPP deviations once we start using samples starting mid-1980s. Interestingly, this period roughly coincides with previous studies' findings on the timing of a possible structural change toward stability in the U.S./world economy. Hence, this result further supports the hypothesis of world economic stabilization. The paper also examined the relation between the persistence of PPP deviations and two de facto measures of financial integration, confirming that they are strongly correlated for all countries. This finding suggested that the recent promotion of financial integration is one of the main sources of decline in PPP deviation persistence.

These conclusions raise the obvious question regarding factors behind the decline in the persistence of PPP deviations. Our results indicated possibility that financial market integration and world economic stabilization have played an important role, but this does not answer the question completely. The decline in PPP deviation persistence may also reflect such factors as increase of world economic relation, competition and globalization, development of world transportation system, reduction of trade barriers, evolution in information technology, and improvement of monetary policy design and implementation. However, investigating which factors are more important remains an open question.

As a final contribution, the paper also opens up an interesting econometric issue. If the conclusions of this study are regarded as robust, and we believe they are, investigating the dynamics of PPP deviation persistence more carefully would be a conceivable agenda for further research. Our results strongly suggest that the order of fractional integration is changing over time. Therefore, examining which kind of model can describe the dynamics well could be a fruitful endeavor. Obviously, one-time permanent structural change is a one way, while gradual change can be another possibility. Accommodating both models, smooth transition parameter model by Lin and Teräsvirta (1994) may be one attractive way to proceed.

Appendix: sketch of the proof of (6)

We show that (6) holds for the ELW estimator of Shimotsu and Phillips (2005) when the data are generated by

$$\begin{aligned}(1-L)^{d_1} X_t &= u_t \mathbf{1}\{t \geq 1\}, \quad t = 1, \ldots, n_1, \\ (1-L)^{d_2} X_t &= u_t \mathbf{1}\{t \geq 1\}, \quad t = n_2, \ldots, n, \quad n_2 > n_1,\end{aligned}$$

i.e., the initial value of the processes is zero. Then, the asymptotic distribution of the two-step FELW estimator follows from repeating the argument of Shimotsu (2006).[7]

Let $R_1(d)$ and $R_2(d)$ be the objective function defined analogously to $R(d)$ in (4) but using $X_1, \ldots, X_{n_1}$ and $X_{n_2}, \ldots, X_n$, respectively. It follows from a Taylor expansion

$$m^{1/2}(\hat{d}_1 - d_1) = -\left[\frac{\partial^2}{\partial d^2} R_1(\bar{d})\right]^{-1} m^{1/2} \frac{\partial}{\partial d} R_1(d_1), \quad \bar{d} \in [d_1, \hat{d}_1].$$

It follows from Shimotsu and Phillips (2005, p.1916 and p.1918) that $(\partial^2/\partial d^2) R_1(\bar{d}) = 4 + o_p(1)$ and

$$m^{1/2} \frac{\partial}{\partial d} R_1(d_1) = \frac{2m^{-1/2} \sum_{j=1}^m \nu_j \left[2\pi I_{1\varepsilon}(\lambda_j) - 1\right] + o_p(1)}{1 + o_p(1)} \to_d N(0, 4),$$

where $I_{1\varepsilon}(\lambda_j)$ is the periodogram of $\varepsilon_1, \ldots, \varepsilon_{n_1}$, and $\nu_j = \log j - m^{-1} \sum_{j=1}^m \log j$. Therefore,

$$m^{1/2}(\hat{d}_1 - d_1) = -\frac{1}{2} m^{-1/2} \sum_{j=1}^m \nu_j \left[2\pi I_{1\varepsilon}(\lambda_j) - 1\right] + o_p(1).$$

Similarly, we obtain

$$m^{1/2}(\hat{d}_2 - d_2) = -\frac{1}{2} m^{-1/2} \sum_{j=1}^m \nu_j \left[2\pi I_{2\varepsilon}(\lambda_j) - 1\right] + o_p(1),$$

where $I_{2\varepsilon}(\lambda_j)$ is the periodogram of $\varepsilon_{n_2}, \ldots, \varepsilon_n$.

First, consider a special case in which ε_t is iid. Then, $\hat{d}_1$ and $\hat{d}_2$ are asymptotically independent because $\sum_{j=1}^m \nu_j[2\pi I_{1\varepsilon}(\lambda_j) - 1]$ and $\sum_{j=1}^m \nu_j[2\pi I_{2\varepsilon}(\lambda_j) - 1]$ are independent from the independence between $\varepsilon_1, \ldots, \varepsilon_{n_1}$ and $\varepsilon_{n_2}, \ldots, \varepsilon_n$. Thus (6) follows.

For a general case where ε_t is a martingale difference sequence, as assumed in Shimotsu and Phillips (2005), a more tedious argument is required. We only provide an outline of the proof. The required result follows if we show

$$\begin{pmatrix} m^{-1/2} \sum_{j=1}^m \nu_j[2\pi I_{1\varepsilon}(\lambda_j) - 1] \\ m^{-1/2} \sum_{j=1}^m \nu_j[2\pi I_{2\varepsilon}(\lambda_j) - 1] \end{pmatrix} \to_d N(0, I_2),$$

[7] Shimotsu (2006) shows that the FELW estimator accommodates non-zero initial condition, and has the same asymptotic distribution as the ELW estimator. See Shimotsu (2006).

where I_2 is a 2×2 identity matrix. As in Robinson (1995, p.1644), write down the left hand side as

$$\begin{pmatrix} \sum_{t=1}^{n_1} z_{1t} \\ \sum_{t=n_2}^{n} z_{2t} \end{pmatrix},$$

where z_{1t} and z_{2t} are martingale difference sequences, and defined analogously to z_t in Robinson (1995, p.1644). Then, applying a martingale CLT to this, as in Robinson (1995, pp.1644-47), shows that this converges to $N(0, I_2)$ in distribution. $\square$

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Table 1: Estimates of d: $m = 396^{0.65} = 48$

Country	ELW	95% CI	Z_t	half-life	LW
Austria	1.042	[0.900, 1.183]	-1.842	[529, ∞]	1.039
Belgium	1.052	[0.910, 1.193]	-2.005	[1175, ∞]	1.050
Canada	0.981	[0.840, 1.123]	-1.390	[37, ∞]	0.977
Denmark	1.034	[0.893, 1.176]	-1.830	[330, ∞]	1.031
Finland	1.017	[0.875, 1.158]	-2.354	[129, ∞]	1.016
France	1.075	[0.933, 1.216]	-2.112	[17434, ∞]	1.072
Germany	1.038	[0.896, 1.179]	-1.852	[412, ∞]	1.033
Greece	0.985	[0.844, 1.127]	-1.076	[41, ∞]	0.977
Italy	1.022	[0.880, 1.163]	-1.806	[164, ∞]	1.019
Japan	0.999	[0.858, 1.141]	-1.916	[65, ∞]	0.988
Netherlands	1.030	[0.888, 1.171]	-2.038	[249, ∞]	1.028
Norway	0.967	[0.825, 1.108]	-2.111	[25, ∞]	0.965
Portugal	0.977	[0.835, 1.118]	-1.237	[33, ∞]	0.976
Spain	1.087	[0.945, 1.228]	-1.685	[168917, ∞]	1.085
Sweden	1.033	[0.891, 1.174]	-2.128	[301, ∞]	1.030
Switzerland	0.991	[0.850, 1.133]	-2.184	[49, ∞]	0.985
United Kingdom	0.916	[0.774, 1.057]	-1.587	[10, ∞]	0.925

Table 2: Estimates of d from two subsamples

Country	d_1	d_2	$d_1 - d_2$	p-value
Austria	1.154	0.961	0.193	7.1%
Belgium	1.173	1.001	0.172	9.6%
Canada	1.188	1.004	0.184	8.0%
Denmark	1.219	0.995	0.224	4.4%
Finland	1.098	1.071	0.028	41.7%
France	1.235	1.001	0.235	3.7%
Germany	1.159	0.972	0.188	7.7%
Greece	1.014	0.991	0.023	43.1%
Italy	1.195	1.016	0.179	8.7%
Japan	1.258	0.959	0.298	1.2%
Netherlands	1.183	0.978	0.205	5.9%
Norway	1.080	0.961	0.119	18.2%
Portugal	1.038	1.019	0.019	44.1%
Spain	1.218	0.967	0.252	2.8%
Sweden	1.142	1.046	0.096	23.3%
Switzerland	1.131	0.920	0.212	5.4%
United Kingdom	1.070	1.008	0.062	32.0%

Table 3: Estimates of d and asy. p-values for varying m

m	25	27	29	31	33
Country			Estimate		
Austria	-0.129	-0.153	-0.193	-0.214	-0.166
Belgium	-0.122	-0.145	-0.172	-0.187	-0.142
Canada	-0.143	-0.155	-0.184	-0.161	-0.168
Denmark	-0.183	-0.200	-0.224	-0.181	-0.142
Finland	-0.076	-0.077	-0.028	-0.046	-0.037
France	-0.173	-0.199	-0.235	-0.243	-0.225
Germany	-0.114	-0.148	-0.188	-0.207	-0.156
Greece	-0.112	-0.030	-0.023	-0.019	0.006
Italy	-0.198	-0.215	-0.179	-0.204	-0.208
Japan	-0.291	-0.317	-0.298	-0.199	-0.126
Netherlands	-0.126	-0.160	-0.205	-0.178	-0.142
Norway	-0.087	-0.135	-0.119	-0.159	-0.176
Portugal	0.025	-0.009	-0.019	-0.012	-0.007
Spain	-0.233	-0.246	-0.252	-0.220	-0.198
Sweden	-0.172	-0.134	-0.096	-0.137	-0.146
Switzerland	-0.188	-0.201	-0.212	-0.262	-0.213
United Kingdom	-0.113	-0.109	-0.062	0.015	0.031
Country			Asy. p-value		
Austria	18.1%	13.1%	7.1%	4.6%	8.9%
Belgium	19.5%	14.4%	9.6%	7.1%	12.4%
Canada	15.6%	12.8%	8.0%	10.3%	8.6%
Denmark	9.8%	7.1%	4.4%	7.7%	12.5%
Finland	29.5%	28.5%	41.7%	35.9%	38.3%
France	11.1%	7.2%	3.7%	2.8%	3.4%
Germany	20.9%	13.8%	7.7%	5.1%	10.2%
Greece	21.5%	41.4%	43.1%	44.2%	51.8%
Italy	8.0%	5.7%	8.7%	5.4%	4.6%
Japan	2.0%	1.0%	1.2%	5.9%	15.4%
Netherlands	18.6%	12.0%	5.9%	8.1%	12.5%
Norway	26.9%	16.1%	18.2%	10.5%	7.6%
Portugal	56.9%	47.3%	44.1%	46.3%	47.8%
Spain	5.0%	3.5%	2.8%	4.2%	5.4%
Sweden	11.2%	16.3%	23.3%	14.0%	11.8%
Switzerland	9.2%	7.0%	5.4%	2.0%	4.2%
United Kingdom	21.2%	21.1%	32.0%	54.5%	59.8%

Table 4: Estimates of d and asy. p-values for varying sample period

First period	1974-1984		1974-1985		1974-1986		1974-1987		1974-1988	
Second period	1988-2006		1989-2006		1990-2006		1991-2006		1992-2006	
Country	$d_1 - d_2$	p-value	$d_1 - d_2$	p-value	$d_1 - d_2$	p-value	$d_1 - d_2$	p-value	$d_1 - d_2$	p-value
Austria	0.148	12.9%	0.142	14.0%	0.167	10.2%	0.224	4.4%	0.193	7.1%
Belgium	0.197	6.6%	0.185	8.0%	0.188	7.6%	0.224	4.4%	0.172	9.6%
Canada	0.018	44.5%	0.107	20.8%	0.139	14.4%	0.111	19.9%	0.184	8.0%
Denmark	0.179	8.7%	0.182	8.3%	0.240	3.4%	0.266	2.2%	0.224	4.4%
Finland	0.047	36.0%	0.083	26.4%	0.103	21.6%	0.110	20.0%	0.028	41.7%
France	0.191	7.3%	0.237	3.6%	0.221	4.6%	0.257	2.5%	0.235	3.7%
Germany	0.137	14.9%	0.133	15.5%	0.162	10.9%	0.221	4.6%	0.188	7.7%
Greece	-0.006	51.9%	-0.089	75.2%	-0.022	56.7%	0.014	45.8%	0.023	43.1%
Italy	0.162	10.8%	0.173	9.4%	0.226	4.3%	0.281	1.6%	0.179	8.7%
Japan	0.218	4.8%	0.247	3.0%	0.250	2.9%	0.234	3.7%	0.298	1.2%
Netherlands	0.157	11.6%	0.169	9.9%	0.178	8.7%	0.226	4.3%	0.205	5.9%
Norway	0.160	11.2%	0.150	12.6%	0.113	19.5%	0.167	10.2%	0.119	18.2%
Portugal	0.013	46.1%	-0.064	68.8%	-0.058	67.1%	-0.017	55.0%	0.019	44.1%
Spain	0.208	5.7%	0.190	7.4%	0.186	7.9%	0.281	1.6%	0.252	2.8%
Sweden	0.132	15.7%	0.210	5.5%	0.211	5.4%	0.186	7.9%	0.096	23.3%
Switzerland	0.206	5.9%	0.203	6.1%	0.200	6.4%	0.276	1.8%	0.212	5.4%
United Kingdom	0.191	7.3%	0.121	17.8%	0.102	21.9%	0.070	29.6%	0.062	32.0%

Table 5: Correlation between financial integration measures and rolling window estimates

Country	Corr(d,IFIGDP)	Corr(d,GEQGDP)
Austria	-0.882	-0.881
Belgium	-0.854	-0.876
Canada	-0.794	-0.737
Denmark	-0.869	-0.918
Finland	-0.519	-0.405
France	-0.886	-0.841
Germany	-0.871	-0.855
Greece	-0.606	-0.673
Italy	-0.838	-0.821
Japan	-0.470	-0.815
Netherlands	-0.875	-0.870
Norway	-0.835	-0.855
Portugal	-0.587	-0.559
Spain	-0.953	-0.944
Sweden	-0.862	-0.819
Switzerland	-0.837	-0.834
United Kingdom	-0.882	-0.917

Figure 1: Dynamics of PPP deviation persistence for G7 countries

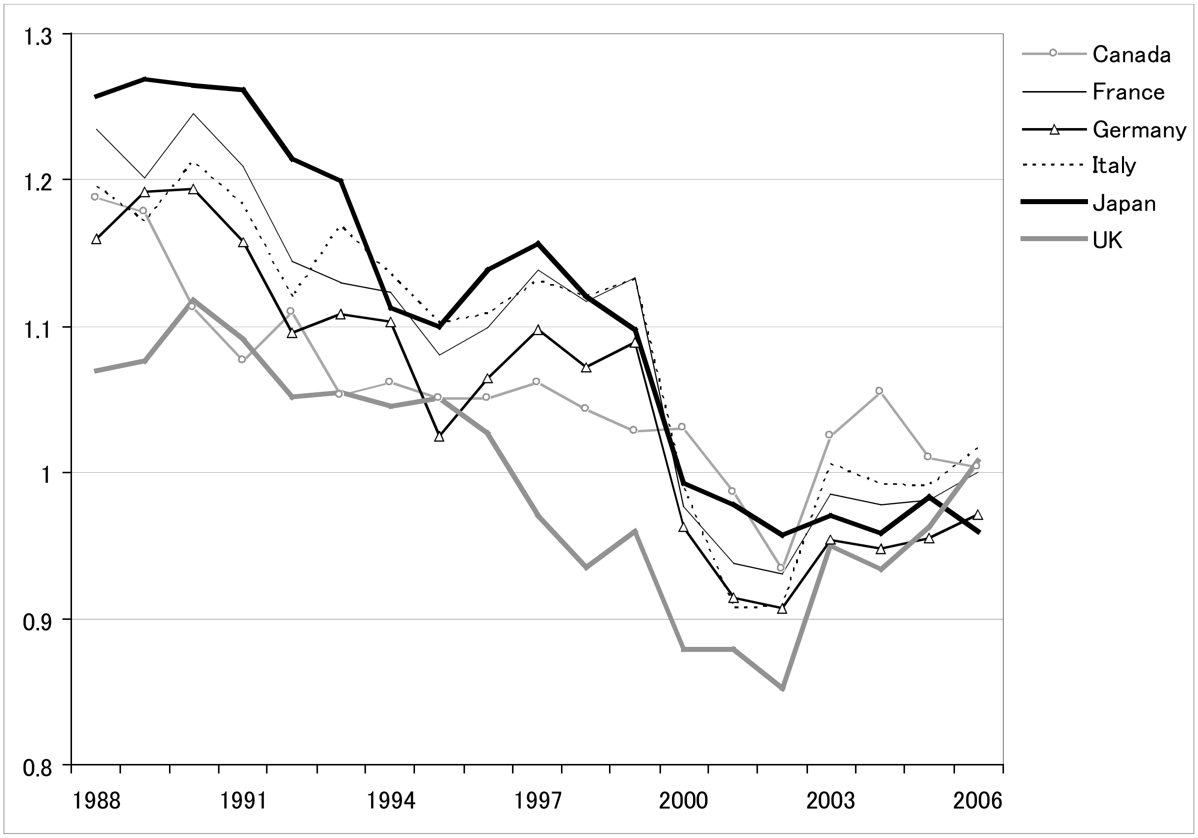

Figure 2: Dynamics of PPP deviation persistence for non-G7 countries with a significant decline

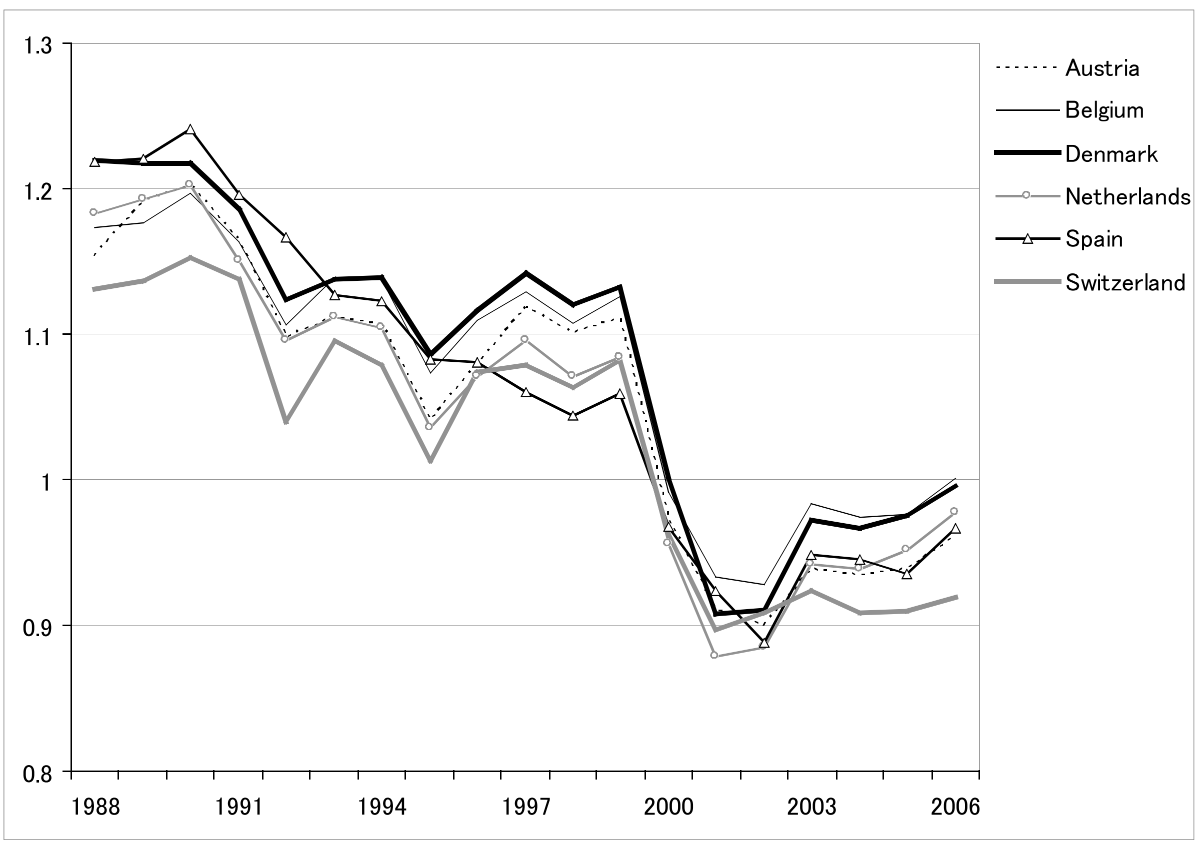

Figure 3: Dynamics of PPP deviation persistence for non-G7 countries with an insignificant decline