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EUROSYSTEM

Working Paper

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SEPTEMBER 2026

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Published by the Bank of Greece, Athens, Greece

ISSN: 2654-1912 (online)
DOI: https://doi.org/10.52903/wp2026368

ON THE TRANSMISSION OF PURCHASING POWER PARITY IN A GROUP SETTING

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ABSTRACT

Two characteristics of recent empirical studies on purchasing power parity (PPP) are that they have (1) examined whether PPP holds for individual countries and (2) investigated separately the respective roles of price adjustment and nominal exchange rate adjustment in bringing about PPP, with the magnitude of price adjustment estimated under fixed exchange rate regimes, and the magnitude of nominal exchange rate adjustment estimated under floating rates. We develop an extension of the Johansen reduced rank regression procedure, which we apply in a panel context. This extension allows us to both test the PPP hypothesis and to estimate the adjustment coefficients in a full panel more efficiently than in previous studies which use single-country data. We also extend the PPP model to allow differences between floating and fixed exchange rate regimes to test whether the adjustment process is different in the two regimes.

JEL Classification: C32; E52; E58

Keywords: Purchasing Power Parity; Exchange Rates; panel data; Johansen procedure; fixed exchange rates; floating exchange rates

Acknowledgments: We are grateful to Joscha Beckman, Mark Spiegel, and other participants at the 30th International Conference on Macroeconomic Analysis and International Finance held at the University of Crete in Rethymno, Crete, on May 27-30, 2026, for very helpful comments on an earlier version of this paper. We thank Elisavet Bosdelekidou and Maria Monopoli for research assistance. The views expressed in this paper are those of the authors and not necessarily those of the Bank of Greece or the Eurosystem.

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1. Introduction

A large empirical literature has investigated – and debated – whether the purchasing power parity (PPP) theory of exchange rates holds – that is, whether the percentage change in the exchange rate over a certain period of time just offsets the difference in inflation rates in the countries concerned over the same period.[1] According to Lothian (2016, 20), a general finding of the recent literature is the following: “as a long-run proposition, PPP is indeed a very useful approximation. This is so both with regard to the behavior of nominal exchange rates under floating rates and the behavior of price levels among countries under fixed rates.”[2] As suggested in those remarks, a characteristic of the empirical literature on PPP is that it has investigated separately the respective roles of the price and nominal exchange rate adjustment mechanisms in bringing about PPP: the price mechanism has been studied under fixed exchange rates, and the nominal exchange rate mechanism has been studied under regimes of floating rates.

In this paper, we investigate both the relative roles of the nominal exchange rate and the price adjustment mechanisms in bringing about PPP under both fixed and flexible rates. We use a panel of 15 advanced countries (in addition to the United States) covering 151 years (1870-2020), taken from the Jordà-Schularick-Taylor microhistory database (Jordà et al., 2017).[3] To conduct our investigation, we employ a reduced rank panel system – that is, a single cointegrating vector, interpreted as the PPP relationship -- which allows us to examine and test the adjustment process towards PPP. In that way, we identify the respective contributions of relative prices and nominal exchange rates

[1] Recent studies include Papell and Prodan (2020), Vo and Vo (2023), Ong (2024), and Lothian and Devereux (2026). The above definition corresponds to relative PPP, which is to be distinguished from absolute PPP. The latter maintains that identical goods should have identical prices expressed in a common currency across countries.

[2] During the past 30 years, the late James Lothian was the most prolific empirical researcher on the issue of purchasing power parity. See, for example, Lothian (1997, 2002), Lothian and Taylor (2000), and Lothian and McCarthy (2002). For a review of his findings, see Lothian (2016).

[3] We follow Papell and Prodan (2020) and Ong (2024) in using this database. The countries used in this study are: Australia, Belgium, Canada, Denmark, Finland, France, Italy, Japan, the Netherlands, Norway, Portugal, Sweden, Spain, Switzerland, the United Kingdom, and the United States. All bilateral exchange rates are against the U.S. dollar. The Jordà et al. database also includes data for Germany and Ireland, but the exchange-rate data for those two countries contain many missing observations.

in the determination of PPP for the entire set of countries -- rather than individual countries studied in previous works.

We make two contributions to the empirical PPP literature.[4] First, we set-up a system panel estimator which imposes a reduced rank restriction, and we estimate the system using full-information maximum likelihood. Our procedure is equivalent to the Johansen procedure, which has been used in previous studies. However, whereas previous empirical studies of PPP have applied the Johansen procedure to single country estimation, our study is the first to extend the Johansen's approach to a panel context.[5] This extension allows us to both test the PPP hypothesis and to estimate the adjustment coefficients in a full panel more efficiently than in previous studies. Second, we extend the model to allow differences between floating and fixed exchange rate regimes to test whether the adjustment process is different in the two regimes. In most previous studies, the effects of the exchange rate regime (fixed and floating regimes) have not been specifically taken into account. Two exceptions (to our knowledge) are studies by Sarno and Valente (2006) and Lothian (2016), both of which split their respective samples into periods arranged by the particular exchange rate regime in effect, and used an identical group of countries in each regime. Apart from those two studies, all other studies of PPP have (again to our knowledge) averaged the effects of separate exchange rate systems together.

Our findings provide evidence for the existence of PPP in the long run. Specifically, we obtain an estimated coefficient on relative inflation that is not significantly different from unity, so that changes in the nominal exchange rate are reflected one-to-one in changes in relative inflation rates. However, we also find that the adjustment toward PPP differs markedly between fixed and floating exchange rate regimes, with exchange rate adjustment playing a larger role under floating than under fixed exchange rates.

The Jordà *et al.* database records a considerable amount of detail on the exchange rate regime followed by countries over time: it records periods of (1) a fixed regime (*e.g.*, the gold standard), (2) a floating peg -- a regime under which the authorities

[4] For a recent survey of the empirical literature on PPP, see Vo and Vo (2023). Taylor and Taylor (2004) provided an earlier survey of the literature.

[5] Single country estimation has led to the result that the validation of PPP has depended on the finding that PPP holds for the majority -- but not all -- of the countries in the sample. See, for example, Papell and Prodan (2020, pp. 7-10) and Ong (2024, p. 7).

declare that they allow their currency to float but they nevertheless intervene to prevent the currency from fluctuating freely,[6] and (3) a free float (the post-1973 period), for each country for each year. On an annual basis there is typically little difference between a floating peg and a pure float. Therefore, we focus on two regimes: the strict fixed regime and a combination of the two floating variants. Over the 150 years spanned by the Jordà *et al.* data, at least one currency has been fixed around 52 per cent of the time; there are periods when many currencies were fixed simultaneously. By limiting their sample to five countries, Sarno and Valente (2006) and Lothian (2016) were able to split their samples into four exchange rate regimes by assuming that an identical regime applied to each country at each point in time. Such a restriction is not likely to be realistic in a sample comprised of sixteen countries – for example, the Canadian dollar floated throughout the 1950s while most other currencies were fixed (against the U.S. dollar). In contrast to Sarno and Valente (2006) and Lothian (2016), we estimate a panel over the entire 151 years of data, allowing each individual country to switch between a fixed and a floating regime (and vice versa), where the specific regime is identified in the Jordà *et al.* database.

In his study using the Jordà *et al.* database, Ong (2024) applied a number of estimation strategies, starting with the reduced rank Vector Error Correction Model (VECM). Based on the cross-country average of the system individual countries he considered, he confirmed that PPP holds in the long run. But given a lack of suitable software, Ong was not able to estimate the reduced rank PPP relationship based on the full panel; nor did he test whether the unit coefficient restriction was an acceptable one. Of course, given the presence of fixed exchange rate regimes in his sample it is likely that the model he estimated must have been unstable because the adjustment parameters would be expected to differ in the different regimes. The standard super consistency result for cointegrating relationships means that Ong's conclusion regarding PPP remains valid, but his results leave open the issue of how PPP comes about. In what follows, we address this issue.

The remainder of this paper is comprised of four sections. Section 2 provides a brief overview of both the recent empirical and theoretical literature on PPP. Section 3

[6] This arrangement has been called a "fear of floating." See Calvo and Reinhart (2002).

briefly discusses the determinations of the reduced rank VECM, and our generalization of it. Section 4 reports our empirical findings. Section 5 concludes.

2. Literature review

In what follows, we first briefly discuss the recent empirical literature on PPP, which, by-and-large, concludes that PPP holds in the long run; that is, it is trade in goods that dominates in the long run. We then discuss the modern theoretical literature, under which financial markets play a dominant role in influencing the exchange rate in the short and medium terms.

2.1 Empirics

PPP has been a central concept in international macroeconomics for over a century. The theory asserts that exchange rates adjust to equalize the purchasing power of different currencies, linking domestic price levels to international exchange rates (Cassel, 1918). Cassel (1918) provided the first formal articulation of PPP, arguing that exchange rates should reflect relative price levels to restore equilibrium after World War I. Modern interpretations distinguish between:

- Absolute PPP, which states that the nominal exchange rate equals the ratio of national price levels.[7]
- Relative PPP, which predicts that exchange-rate changes correspond to inflation differentials.

Empirical studies in the 1970s and 1980s cast doubt on PPP. Frenkel (1978) found exchange rates to be far more volatile than relative price levels, while Roll (1979) showed that deviations from PPP persisted for long periods. Rogoff (1996) famously described the slow rate of mean reversion of a shock to a real exchange rate to dissipate and return to equilibrium as the "PPP puzzle", noting that real exchange rates revert to parity at an implausibly slow rate. Rogoff (1996) estimated the half-life of deviations

[7] PPP is closely related to the Law of One Price (LOOP), which holds that identical tradable goods should sell for the same price globally when expressed in a common currency. Deviations from LOOP—due to transport costs, tariffs, and product differentiation—accumulate into macro-level deviations from PPP (Taylor et al., 2001).

from PPP to be 3 to 5 years, noting that it is far too long to be explained by short-term sticky prices and nominal rigidities.

Early econometric tests reinforced this pessimism. Unit-root tests often failed to reject the hypothesis that real exchange rates follow a random walk (Perron, 1989). Later cointegration approaches, based on as Johansen (1988), provided stronger evidence for long-run PPP, but the results remained sensitive to sample length and structural breaks.

Lothian and Taylor's (1996) landmark study used more than two centuries of data for the U.S. dollar–British pound and French franc–British pound exchange rates. They found strong mean reversion in real exchange rates, providing some of the most compelling evidence for long-run PPP. Their findings challenged the short-sample literature and helped resolve aspects of Rogoff's (1996) PPP puzzle. Lothian and Taylor (2000) expanded this analysis to a broader set of currencies and time periods, arguing that many earlier studies suffered from "temporal myopia" due to reliance on post-1973 floating-rate data. Specifically, they found that long-horizon regressions yield more reliable estimates of mean reversion than short-horizon tests, which often suffered from low statistical power. Their long-horizon datasets smoothed short-term volatility and revealed underlying parity relationships more clearly. A hallmark of Lothian's research was his use of datasets spanning centuries, enabling him to observe exchange-rate behaviour across multiple monetary regimes and major historical shocks. This approach allows for more accurate testing of whether mean reversion is a persistent feature of exchange-rate dynamics. Subsequently, Lothian (2016) stressed that ignoring long-run historical evidence leads to biased conclusions about PPP. He also integrated structural-break analysis—building on Perron (1989) — to show that accounting for wars, crises, and regime changes, strengthens evidence for PPP. By splitting the sample into fixed and floating periods, Lothian (2016) has been one of the few researchers to explicitly recognize the importance of regime change between floating and fixed regimes.

As noted, two recent studies, by Papell and Prodan (2020) and Ong (2024), respectively, used the Jorda et al. data base to test for PPP. Papell and Prodan used nominal U.S. dollar exchange rates for 16 countries, and found evidence of PPP based on two tests: the confidence intervals for long-run linear correlation coefficients did not contain zero and the confidence intervals for long-run regression coefficients contained unity. With a “long run” of six periods, each with a length of 24 years, and 90 percent confidence intervals, long-run PPP could not be rejected for 9 of the 16 countries. Analogous results were found if the “long-run” consisted either of nine periods of 16 years or 3 periods of 48 years. Similarly, using Augmented-Dickey-Fuller tests, the null of a unit root could not be rejected against a level stationary alternative for 9 of the 16 real exchange rates at the 5 percent level. Papell and Prodan did not provide estimates of half-life.[8]

Ong (2024) also tested for PPP for the identical panel of 16 countries used by Papell and Prodan.[9] Using the System Pooled Mean group (SPMG) model to study the adjustment of exchange rates and relative prices, Ong found that the unit long-run elasticity between the level of the nominal exchange rate and inflation differentials could not be rejected for 9 of the 16 countries at a 10 percent significance level. He also found that relative prices adjusted to maintain a constant exchange rate in all 16 countries; only Germany and Italy did not show significant adjustments of nominal exchange rates. Ong (2024, 7) estimated an average half-life of 4.87 years.

Lothian and Devereux (2026) investigated real exchange rate behavior using data for the Netherlands and Great Britain over the period 1590 to 2020 – that is, over the course of 430 years. Using consumer prices for both countries, the authors found evidence in favor of relative PPP. In addition, they found that the real exchange rate was mean reverting. However, Lothian and Devereux also found that the real exchange rate was subject to two prolonged disturbances from PPP. First, a large and prolonged real appreciation of sterling began in the late-18th century and lasted until 1931. The authors attributed the appreciation to trade disruptions, including from the British Corn laws, and the effects of the British industrial revolution. Second, the authors found that rates of convergence to PPP for the real exchange rate was faster in the early modern

[8] In an earlier paper, Papell (1997) estimated a panel half-life of 2.5 years. Subsequently, Murray and Papell (2005) utilized a median-unbiased panel method that revised that estimate to 4 years. In his work, Papell frequently underscores that, while point estimates hover between 2 to 4 years, the 95 percent confidence intervals are often very wide (sometimes ranging from 1.5 years to infinity).

[9] As mentioned, in this study we use the Jorda et al. data base to test for PPP against the U.S. dollar for 15 countries. Specifically, we exclude Germany from our sample, in contrast to Papell and Prodan (2020) and Ong (2024). We found that the data for nominal exchange rates for Germany in the Jorda et al. data base had many missing observations so that it was inappropriate to estimate a PPP relationship for that country.

period, suggesting that "the 'PPP puzzle' is not a feature of early centuries" (Lothian and Devereux, 2026, 466).

2.2 Modern theory

The Meese and Rogoff (1983) empirical finding that structural exchange rate models were not able to forecast exchange rates better than a random walk model inflicted a devastating blow to established theories of exchange rate determination, effectively leading to the cessation of theoretical research on the *macroeconomic* determinants of exchange rates. After a long lull, this research area is now witnessing a strong revival. Leading examples are papers by Engel and Wu (2023, 2024), Itskhoki and Mukhin (2021, 2025) and Koijen and Yogo (2026). This body of work aims not only at understanding specific exchange rate relations such as PPP -the subject of the present paper- or interest rate parity (IRP) but also at providing -to different degrees - a unifying explanation of several, interrelated, extant phenomena, namely, the so called exchange rate puzzles (Obstfeld and Rogoff, 2001, Itskhoki and Mukhin, 2025). We briefly summarize the main innovations of the modern literature, and then venture an assessment of how our empirical findings relate to it.

For comparative purposes, it is useful to start with the "older" theoretical work on PPP. In that literature, PPP failure was attributed either to (1) nominal rigidities; or, (2) *relative* price changes (the price between traded vs non-traded goods; or between domestic vs foreign traded goods) arising from supply or preference shocks (Stockman, 1987). The former theory has lost momentum because the commonly accepted degree of nominal stickiness is not large enough to account for the very high persistence of PPP deviations. The latter theory has fallen out of fashion because movements in exchange rates seem disconnected from the standard macroeconomic fundamentals that ought to drive relative prices movements (the Meese and Rogoff results as well as the Mussa (1986) puzzle).[10]

Modern theories of exchange rate determination and PPP have shifted focus away from goods markets to financial markets as the main focus of exchange rate

[10] The Mussa puzzle refers to the observation that large and sudden jumps in the volatility of both nominal and real exchange rates occurred after the adoption of a floating exchange rate regime by the large industrial countries in 1973.

determination, thus proposing new mechanisms responsible for the observed large and persistent changes in real exchange rates. Engel and Wu (2023, 2024) introduced liquidity demand premia – the non-monetary “convenience yield” of U.S. government liabilities because of their safety, the ease with which they can be sold, and their value as collateral - and risk premia. They found that taking these elements into account leads to mean reversion in exchange rates to their properly defined equilibria over the medium and long run. In the spirit of Meese and Rogoff, Engel and Wu (2023) found that their New Keynesian model, augmented with a model of the liquidity yield, has a significantly better out-of-sample fit than a random walk model.

Koijen and Yogo (2026) used their pathbreaking demand system approach on exchange rate determination to argue that because global demand for financial assets is highly inelastic -unlike the popular assumption of perfect elasticity- even small capital flows significantly move exchange rates and asset prices, giving rise to a disconnect between these prices and standard macroeconomic aggregates. And as in Engel and Wu, the US assets convenience yield causes deviations from UIRP. PPP failure here is the direct consequence of the large response of exchange rates to shocks because of the inelastic asset demands.

Itskhoki and Mukhin (2021, 2025) used a combination of old and new features to provide an explanation for the large volatility and persistence in PPP violations. The old features are standard: home bias in consumption, low substitutability, local currency pricing (LCP) and price stickiness. Under LCP, exporters set consumer prices in the currency of the destination market. Because local currency prices are assumed to be sticky in the short to medium run, any movement in the nominal exchange rate passes through directly into the real exchange rate (as in Engel and Wu, 2023). The novel feature is financial segmentation and risk-averse intermediaries: international currency flows must be cleared by risk-averse market makers. When financial shocks hit these intermediaries (e.g., a change in risk appetite or a change in the supply of assets), nominal exchange rates move sharply. With sticky prices, these financial shocks impact on real exchange rates, generating persistent violations of PPP. Moreover, if financial market noise is significant -as they assume- this generates large volatility in the deviations from PPP that is disconnected from macroeconomic fundamentals.

Finally, Crucini et al. (2026) introduced behavioral features, namely rational inattention to sever the relationship between fundamentals and exchange rates. In particular, firms' limited responsiveness to aggregate shocks can generate real exchange rate persistence consistent with the data.

The modern theories of exchange rate determination have been specifically built to account for the exchange rate puzzles. But it should be noted that these puzzles concern exchange rate behaviour in the short-medium term. For instance, in their discussion of the Mussa puzzle -the difference in real exchange rate behavior across fixed and flexible exchange rate regimes- Itskhoki and Mukhin use the 1960-1999 period (essentially 20 years per regime). A finding such as ours (below) or Engel and Wu's, that PPP holds well over the *long run* simply implies that in the long term, trade in goods is the main driver of exchange rates: the long run trade balance equilibrium pins down the real exchange rate equilibrium. Note also that consistent with our argument, asset demands are more inelastic in the long run relative to the short run, which reduces the impact of asset market shocks to asset prices and exchange rates in the long term and other such features may be powerful drivers of the exchange rate in the short and perhaps medium run, but they lose their potency over the long run.[11] It is the determinants of trade that matter for the long run properties of PPP. This claim is in line with Lothian and Devereaux's view that the very persistent PPP deviations from late 18-th century to the interwar period were due to the industrial revolution and trade disruption.

3. Empirical approach

3.1 Reduced Rank Systems and Panel analysis

Reduced rank regression models (RRR) have become a staple of time series analysis since the work of Johansen (1988). Following Johansen's contribution, there has been considerable work in testing stationarity and cointegration in a panel setting. Studies using panels have generally followed one of two approaches. First, that of Larsson et al. (2001) who apply the standard RRR model to each member of the panel individually and then essentially average either the cointegrating vectors or the

[11] These other factors include noisetrading, liquidity constraints, rationale inattention, and inelastic asset demands.

likelihood ratio (LR) test for cointegration to give an overall test. Second, a related testing procedure, due to Breitung (2005), pools the data before constructing the reduced rank test and then generalises the Johansen moment matrices to a pooled setting across the panel. Neither approach estimates the full panel RRR model. In a survey paper, Breitung and Pesaran (2008) discuss the extension of the full panel RRR model; they consider a range of possible approaches, although there are very few (if any) examples of a full reduced rank panel estimation.

We suspect that the reason for this scarcity of work on full reduced rank panel estimation is that the technology for estimating a system of equations with a reduced rank imposed, as set-out in Johansen's work, is not easily generalized to a panel setting. The foundational work in this area was developed by Hotelling (1936) and Anderson (1951). Hotelling introduced canonical correlation analysis which became one of the building blocks for RRR, and Anderson extended that idea into the full RRR model. Indeed, Anderson (1958) subsequently covered rank restrictions and likelihood estimation. Izenman (1975) gave a full formalization of the RRR model. Because computing was in its infancy, early researchers required an analytical solution to the RRR problem. In light of the limits imposed by computing, most of the analysis set out in Johansen (1988) and Hall (1989) focused on an analytical solution to the problem, first partialling out the dynamics of the problem, and, then, concentrating on the likelihood function: imposing an identification procedure to get down to a generalized eigenvalue problem, the solution to which gives the reduced rank regression coefficients.

The insight we provide is that an explicit analytical solution is no longer necessary because we can easily set up a panel reduced rank regression model and estimate it using full information maximum likelihood. The only requirement is that the model must be identified by some set of just identifying restrictions. The identification issue was dealt with by Pesaran and Shin (2002), who showed that if there are r cointegrating vectors in a RRR system then we need r^2 restrictions to exactly identify the model. It is then possible to calculate the Johansen likelihood ratio tests of the rank of the system by first estimating a model with no cointegrating vectors, then with one, then with two and so on and calculating likelihood ratio tests of one assumption against another. We implemented this procedure using EVIEWS. Appendix A shows some of the actual code used in this paper to calculate a panel RRR model.

3.2 The PPP relationship as a reduced rank panel model

We begin by outlining the standard reduced rank regression model following Johansen (1991), which we generalized to a panel setting using the specific specification of the PPP relationship used by Ong (2024). We define a two equation system involving y_{it} as the log of the spot exchange rate against the US dollar and $x_{it} = \log(p_{it}) - \log(p^*_t)$ that is the difference between the log of the price level (defined as the CPI) for country i (p_{it}) and the log of the US price level (p^*_t). The reduced rank system with only one cointegrating vector can then be stated as

$$\begin{aligned}\Delta y_{it} &= \gamma_{1i} + \alpha_1(y_{it-1} - \beta x_{t-1i}) + \sum_{l=1}^{p} \delta_{1l} \Delta y_{it-l} + \sum_{l=1}^{p} \gamma_{1l} \Delta x_{t-li} + \varepsilon_{1ti} \\ \Delta x_{it} &= \gamma_{2i} + \alpha_2(y_{it-1} - \beta x_{t-1i}) + \sum_{l=1}^{p} \delta_{2l} \Delta y_{it-l} + \sum_{l=1}^{p} \gamma_{2l} \Delta x_{t-li} + \varepsilon_{2ti}\end{aligned} \tag{1}$$

At this point we assume that there is only one cointegrating vector. This assumption will be tested in the next section. Most of the PPP literature has concentrated on the parameter β; under PPP this should take the value of unity. But the parameters $\alpha_1 \, and \, \alpha_2$, the loading weights, which show how the cointegrating vector affects the adjustment of each variable, are also important since they control the long run causality of the system. If $\alpha_1 < 0$ and $\alpha_2 = 0$, then the exchange rate adjusts to bring PPP about in the long run. If $\alpha_1 = 0$ and $\alpha_2 > 0$, then the domestic price level adjusts over time. If $\alpha_1 < 0$ and $\alpha_2 > 0$, both the exchange rate and the price level will adjust to bring about PPP in the long run. Hence, these coefficients determine what happens when we fix the exchange rate. Under a fixed exchange rate, $\alpha_1 = 0$ since the exchange rate does not adjust. The question then arises: does α_2 become larger so that the domestic price adjusts more quickly or does it remain unchanged?

It is worth noting that the constants in each equation vary between each country. This is equivalent to a fixed country and equation effect. This is important in this context because all the exchange rates are scaled very differently; in a log formulation, this scaling is taken up by the constants.

We now turn to the issue of fixed and floating exchange rate regimes. After investigating the model (1) above, we generalise the model in the following way to allow for the change in regime.

$$\begin{aligned}
\Delta y_{it} &= FIX_{it}(\gamma_{11i} + \alpha_{11}(y_{it-1} - \beta x_{t-1i}) + \sum_{l=1}^{p}\delta_{11l}\Delta y_{it-l} + \sum_{l=1}^{p}\gamma_{11l}\Delta x_{t-li}) \\
&(1 - FIX_{it})(\gamma_{21i} + \alpha_{21}(y_{it-1} - \beta x_{t-1i}) + \sum_{l=1}^{p}\delta_{21l}\Delta y_{it-l} + \sum_{l=1}^{p}\gamma_{21l}\Delta x_{t-li}) + \varepsilon_{1ti} \\
\Delta x_{it} &= FIX_{it}(\gamma_{12i} + \alpha_{12}(y_{it-1} - \beta x_{t-1i}) + \sum_{l=1}^{p}\delta_{12l}\Delta y_{it-l} + \sum_{l=1}^{p}\gamma_{12l}\Delta x_{t-li}) \\
&(1 - FIX_{it})(\gamma_{22i} + \alpha_{22}(y_{it-1} - \beta x_{t-1i}) + \sum_{l=1}^{p}\delta_{22l}\Delta y_{it-l} + \sum_{l=1}^{p}\gamma_{22l}\Delta x_{t-li}) + \varepsilon_{2ti}
\end{aligned} \tag{2}$$

Where FIX_{it} is a dummy variable which is 1 when country i in period t has a floating exchange rate and 0 when its fixed. This formulation allows all the parameters of the model to change between the fixed and floating exchange rate regime. In this system α_{11} controls the speed of convergence of the nominal exchange rate to PPP under a floating regime, α_{21} controls the speed of convergence of the nominal exchange rate under a fixed exchange rate regime, α_{12} controls the speed of convergence of the domestic price level under the floating regime, and α_{22} controls the speed of convergence of the domestic price level under the fixed regime.

4. Results

We begin by implementing a set of panel stationarity tests. We then test the panel of countries for cointegration. These results largely confirm the findings of Ong (2024), discussed above. We then go on to estimate the reduced rank system with an unrestricted β coefficient and test to see if this can be restricted to unity. We then estimate the full system allowing for the changes in exchange rate regime as described in Section 3.

Tables 1 and 2 report the standard panel tests for stationarity for x and y. The top half of each table shows the first-generation unit root tests which assume cross section independence. The second half of the table report the second-generation tests, which allow for cross section dependence. The overwhelming conclusion from these tests is that both x and y are non stationary processes.

[Insert Tables 1 and 2 here]

These tests, however, may not be appropriate given the general assumption that N>T - that is, the cross-section dimension is greater than the length of the time series. In this study this is not the case since N=15 and T=150. It is possible to test the series individually using conventional time series unit root tests. Table 3 then reports the number of series which are found to be non-stationary using conventional time series tests, where ADF is the Augmented Dickey Fuller test, PP is the Philips and Perron test, KPSS is the Kwiatkowski-Phillips-Schmidt-Shin test, and MZa is the Ng-Perron MZa test. The table shows the number of variables in our cross section that are stationary using standard time series tests. This test shows that, using standard time series tests all the variable are found to be non-stationary. Hence, the conclusion that we are dealing with a set of I(1) variables is confirmed.

[Insert Table 3 here]

Under the Granger representation theorem, if there are n I(1) variables there can be at most r cointegrating vectors where r is at most n-1. In this case n=2 so r can be at most 1. Therefore, we do not have to consider the case of multiple cointegrating vectors. Table 4 reports some standard panel cointegration tests: they conclusively determine that there is cointegration between x and y. We can, therefore, conclude that, since the variables are non-stationary, and cointegrate, there must be only one cointegrating vector linking them together. This justifies the formulation in (1) and (2).

[Insert Table 4 here]

To this point, we have largely confirmed the findings of Ong (2024). We now extend his analysis. Instead of continuing with the general system (1) and estimating the full system, Ong (2024) turned to simpler panel analysis[12]. Specifically, we now estimate the full reduced rank system in (1). Initially, we estimate the key PPP parameter β freely. We estimate the full model using full information maximum likelihood, OLS gives very similar results, so that we can then test the validity of restricting β to unity, which is the value expected under PPP, using the Pesaran and Shin (2002) likelihood ratio test for over-identifying restriction. The system is large (30 equations), and each equation has a sperate constant; hence, we will not report all the parameters of the system since these are not generally of interest. Instead, we focus on the three parameters which are key for the overall performance of the long run part of the system, that is $\beta, \alpha_1 \; and \; \alpha_2$, Table 5 reports the results.

[Insert Table 5 here]

β is the key parameter for PPP, which should, in theory, be unity. Clearly 0.95 is close to unity and we can test this as an overidentifying restriction. The two adjustment coefficients $\alpha_1 \; and \; \alpha_2$ are both correctly signed for stability. Strictly, the condition for stability is that the roots of the $\Pi = \alpha' \beta$ matrix should be less than unity. However, a sufficient condition to satisfy this requirement is that $\alpha_1 > 0 \, and \, \alpha_2 < 0$, which is the case here. All coefficients are significantly different from zero; α_1 is over twice the size of α_2 which implies that the exchange rate adjusts twice as fast from a departure from PPP as does to the domestic price level.

We are now in a position to formally test the PPP restriction. As Pesaran and Shin (2002) show, once the cointegrating rank of the system has been established and a just-identifying set of restrictions has been imposed, any further over-identifying restriction may be tested as a standard likelihood ratio test with a standard $\chi 2$ distribution. Our system (1) is just identified by the imposition of a unit coefficient on y_{it-1} in the cointegrating space. We can, therefore, test if $\beta = 1$ by re-estimating the system subject

[12] This likely reflected the fact that Ong did not have access to software that is able to estimate the full system (1).

to this restriction. The log likelihood of the original system is 4917.83; once we impose $\beta = 1$ the log likelihood becomes 4916.34. Therefore, the LR test is 2(4917.83-4926.34) = 2.99 and the $\chi 2$ (5%) critical value is 3.84, which means that we cannot reject the hypothesis that $\beta = 1$ at a 5% critical value. PPP is, therefore, accepted by the data as a long run property, with adjustment taking place much as we might have expected.

However, the issue of how the system reacts when the exchange rate is fixed remains open, something which is almost never addressed in the literature. To investigate, we expand the system by allowing all the parameters of the system to be different in fixed or floating exchange rate regimes, as in equation (2) above. The model is very non-standard panel, and the way we address this is to set up the whole system of 30 equations as a system (in EVIEWS), where each pair of country equations switches its parameters, depending on whether the exchange rate is fixed or floating in a particular period. Appendix A provides the EVIEWS code that we used. Again, we will not report the full estimation results because many coefficients are of no interest (60 constants, for example). We also maintain the PPP restriction that $\beta = 1$ so that we can focus on the results for the four adjustment coefficients. The results are reported in Table 6.

[Insert Table 6 here]

These results provide answers our two main questions: (1) How does the exchange rate equation change when the exchange rate is fixed? The adjustment coefficient falls substantially and becomes insignificantly different from zero. Thus, during periods of a fixed exchange rate system the exchange rate does not adjust to bring PPP into line. (2) Does the domestic price level adjust more quickly when the exchange rate is fixed? The domestic price level continues to adjust towards PPP at much the same rate regardless of the exchange rate regime.

Consider the model which did not allow for fixed exchange rate regimes; the speed of adjustment of the exchange rate is estimated to be 0.08 (Table 5). Table 6, however, shows this estimate is an average of the two regimes, and that during periods of a floating exchange rate the exchange rate adjusts much more quickly. Calculating the half-life is more complex in this system than in a single equation since both equations have an adjustment towards the long run. Combining these two effects, the half-life of the system is just under 5 years The domestic price adjustment on its own is even slower (a half-life of 10 years). If, for example, the pound sterling was 10 percent overvalued in real exchange rate terms relative to the dollar, and if the nominal exchange rate was fixed at a particular point in time (for example, as at the beginning of the Bretton Woods regime of fixed exchange rate), UK inflation relative to the US had to make up the difference. If the US inflation rate averaged 2 percent per year, and UK inflation averaged 1 percent, it would take 10 years to bring about equilibrium, assuming that nothing else had changed in the interim.

These results are broadly in line with the general PPP literature, which typically finds a long half-life in the adjustment process. Our results, however, are more efficient than other results: we have, for the first time, investigated the mechanics of PPP based on both a long time series and a panel data set applied to a group of countries --rather than individual countries, as in previous studies -- using a full reduced rank regression approach. This approach also allows us to clearly identify the adjustment processes under the differing regimes.

5. Conclusions

To sum up, we have made two extensions to the empirical literature on PPP. First, we estimated the full reduced rank panel model for PPP, which allows us to formally test PPP for a group of countries instead of individual countries within a group. Consequently, our results are more efficient than studies that rely on individual country results (within a group). We found that, in the long run, we can accept the full PPP restriction. Second, we investigated how the exchange rates and domestic prices adjust during periods of fixed and floating exchange rates. While the behavior of domestic prices does not change between the two regimes, there is a very significant difference between the adjustment process for the exchange rate in the two regimes. As expected, during a fixed regime the nominal exchange rate adjustment coefficient is effectively zero, while in the floating periods it is much larger. Ignoring the fact that exchange rate regimes were fixed for a substantial part of the last 150 years leads to a significant and empirically important downward bias in the estimated adjustment speed of the exchange rate. Finally, our finding that PPP holds in the long run suggests a clear distinction between the short and long determinants of exchange rates. Whereas financial markets play a dominant role in the short-medium term -as emphasized by the modern literature-, it is goods markets (trade in goods) that become the dominant force in the long run.

Tables

Table 1
Stationarity tests for y.

Assumption of cross section independence
Sample: 1 151
Exogenous variables: Individual effects
Automatic selection of maximum lags
Automatic lag length selection based on SIC: 0 to 2
Newey-West automatic bandwidth selection and Bartlett kernel

Method	Statistic	Prob.**	Cross-sections	Obs
Null: Unit root (assumes common unit root process)				
Levin, Lin & Chu t*	-0.93571	0.1747	15	2240
Null: Unit root (assumes individual unit root process)				
Im, Pesaran and Shin W-stat	0.63281	0.7366	15	2240
ADF - Fisher Chi-square	26.8778	0.6297	15	2240
PP - Fisher Chi-square	21.9690	0.8552	15	2250

Assumption of cross section dependence;
Bai and Ng Pooled statistic -0.64 p value 0.52
Pesaran CIPS statistic -2.45 p value< 0.05

** Probabilities for Fisher tests are computed using an asymptotic Chi-square distribution. All other tests assume asymptotic normality.

Table 2
Stationarity tests for x.

Assumption of cross section independence
Sample: 1 151
Exogenous variables: Individual effects
Automatic selection of maximum lags
Automatic lag length selection based on SIC: 0 to 4
Newey-West automatic bandwidth selection and Bartlett kernel

Method	Statistic	Prob.**	Cross-sections	Obs
Null: Unit root (assumes common unit root process)				
Levin, Lin & Chu t*	-0.62549	0.2658	15	2219
Null: Unit root (assumes individual unit root process)				
Im, Pesaran and Shin W-stat	1.80905	0.9648	15	2219
ADF - Fisher Chi-square	26.3421	0.6576	15	2219
PP - Fisher Chi-square	28.2062	0.5595	15	2250

Assumption of cross section dependence
Bai and NG Pooled statistic -2.7 p value 0.007
Pesaran CIPS test statistic -2.2 pvalu>0.05

** Probabilities for Fisher tests are computed using an asymptotic Chi-square distribution. All other tests assume asymptotic normality.

Table 3
The results from standard time series unit roots.

	ADF	PP	KPSS	MZa
x	15	15	15	15
y	15	15	15	15

Table 4
Panel tests for cointegration between x and y.

Pedroni Residual Cointegration Test
Null Hypothesis: No cointegration
Trend assumption: No deterministic trend
User-specified lag length: 1
Newey-West automatic bandwidth selection and Bartlett kernel

Alternative hypothesis: common AR coefs. (within-dimension)

	Statistic	Prob.	Weighted Statistic	Prob.
Panel v-Statistic	12.37591	0.0000	9.106494	0.0000
Panel rho-Statistic	-13.74573	0.0000	-13.57860	0.0000
Panel PP-Statistic	-8.189710	0.0000	-8.011318	0.0000
Panel ADF-Statistic	-8.774500	0.0000	-9.076352	0.0000

Alternative hypothesis: individual AR coefs. (between-dimension)

	Statistic	Prob.
Group rho-Statistic	-11.64442	0.0000
Group PP-Statistic	-8.330992	0.0000
Group ADF-Statistic	-9.846332	0.0000

Kao Residual Cointegration Test
Null Hypothesis: No cointegration
Trend assumption: No deterministic trend
User-specified lag length: 1
Newey-West automatic bandwidth selection and Bartlett kernel

	t-Statistic	Prob.
ADF	-9.730407	0.0000
Residual variance	0.020824	
HAC variance	0.013143	

Table 5
The key parameters determining the long run behaviour of system (1).

β	α_1	α_2
0.95 (17.9)	0.08 (7.4)	-0.03 (3.9)

'z' statistics in parenthesis, FIML estimation

Table 6
Adjustment coefficient for system (2).

α_{11} (fixed regime)	α_{21} (floating regime)	α_{21} (fixed regime)	α_{22} (floating regime)
0.04 (1.19)	0.13 (7.4)	-0.03 (3.3)	-0.03 (3.4)

‘z’ Statistics in Parathesis, FIML estimation.

Appendix A: The Eviews code for the reduced rank model with switching

```
dlrxusdaus=peg_strictaus*(c(111)+c(12)*(-lrxusdaus(-1)+dlcpiaus(-1))+c(13)*dlrxusdaus(-
1)+c(14)*ddlcpiaus(-1))+(1-peg_strictaus)*(c(112)+ c(121)*(-lrxusdaus(-1)+dlcpiaus(-
1))+c(131)*dlrxusdaus(-1)+c(141)*ddlcpiaus(-1)  )

ddlcpiaus=c(211)+peg_strictaus*(c(22)*(-lrxusdaus(-1)+dlcpiaus(-1))+c(23)*dlrxusdaus(-
1)+c(24)*ddlcpiaus(-1))+(1-peg_strictaus)*(c(221)*(-lrxusdaus(-1)+dlcpiaus(-1))+c(231)*dlrxusdaus(-
1)+c(241)*ddlcpiaus(-1))

dlrxusdbel=peg_strictbel*(c(114)+c(12)*(-lrxusdbel(-1)+dlcpibel(-1))+c(13)*dlrxusdbel(-
1)+c(14)*ddlcpibel(-1))+(1-peg_strictbel)*( c(115)+c(121)*(-lrxusdbel(-1)+dlcpibel(-
1))+c(131)*dlrxusdbel(-1)+c(141)*ddlcpibel(-1)  )
ddlcpibel=c(212)+peg_strictbel*(c(22)*(-lrxusdbel(-1)+dlcpibel(-1))+c(23)*dlrxusdbel(-
1)+c(24)*ddlcpibel(-1))+(1-peg_strictbel)*(c(221)*(-lrxusdbel(-1)+dlcpibel(-1))+c(231)*dlrxusdbel(-
1)+c(241)*ddlcpibel(-1))

dlrxusdcan=peg_strictcan*(c(116)+c(12)*(-lrxusdcan(-1)+dlcpican(-1))+c(13)*dlrxusdcan(-
1)+c(14)*ddlcpican(-1))+(1-peg_strictcan)*(c(117)+ c(121)*(-lrxusdcan(-1)+dlcpican(-
1))+c(131)*dlrxusdcan(-1)+c(141)*ddlcpican(-1)  )
ddlcpican=c(213)+peg_strictcan*(c(22)*(-lrxusdcan(-1)+dlcpican(-1))+c(23)*dlrxusdcan(-
1)+c(24)*ddlcpican(-1))+(1-peg_strictcan)*(c(221)*(-lrxusdcan(-1)+dlcpican(-1))+c(231)*dlrxusdcan(-
1)+c(241)*ddlcpican(-1))

dlrxusdche=peg_strictche*(c(741)+c(12)*(-lrxusdche(-1)+dlcpiche(-1))+c(13)*dlrxusdche(-
1)+c(14)*ddlcpiche(-1))+(1-peg_strictche)*( c(742)+c(121)*(-lrxusdche(-1)+dlcpiche(-
1))+c(131)*dlrxusdche(-1)+c(141)*ddlcpiche(-1)  )
ddlcpiche=c(214)+peg_strictche*(c(22)*(-lrxusdche(-1)+dlcpiche(-1))+c(23)*dlrxusdche(-
1)+c(24)*ddlcpiche(-1))+(1-peg_strictche)*(c(221)*(-lrxusdche(-1)+dlcpiche(-1))+c(231)*dlrxusdche(-
1)+c(241)*ddlcpiche(-1))

dlrxusddnk=peg_strictdnk*(c(118)+c(12)*(-lrxusddnk(-1)+dlcpidnk(-1))+c(13)*dlrxusddnk(-
1)+c(14)*ddlcpidnk(-1))+(1-peg_strictdnk)*( c(119)+c(121)*(-lrxusddnk(-1)+dlcpidnk(-
1))+c(131)*dlrxusddnk(-1)+c(141)*ddlcpidnk(-1)  )
ddlcpidnk=c(215)+peg_strictdnk*(c(22)*(-lrxusddnk(-1)+dlcpidnk(-1))+c(23)*dlrxusddnk(-
1)+c(24)*ddlcpidnk(-1))+(1-peg_strictdnk)*(c(221)*(-lrxusddnk(-1)+dlcpidnk(-1))+c(231)*dlrxusddnk(-
1)+c(241)*ddlcpidnk(-1))

dlrxusdesp=peg_strictesp*(c(711)+c(12)*(-lrxusdesp(-1)+dlcpiesp(-1))+c(13)*dlrxusdesp(-
1)+c(14)*ddlcpiesp(-1))+(1-peg_strictesp)*( c(712)+c(121)*(-lrxusdesp(-1)+dlcpiesp(-
1))+c(131)*dlrxusdesp(-1)+c(141)*ddlcpiesp(-1)  )
ddlcpiesp=c(216)+peg_strictesp*(c(22)*(-lrxusdesp(-1)+dlcpiesp(-1))+c(23)*dlrxusdesp(-
1)+c(24)*ddlcpiesp(-1))+(1-peg_strictesp)*(c(221)*(-lrxusdesp(-1)+dlcpiesp(-1))+c(231)*dlrxusdesp(-
1)+c(241)*ddlcpiesp(-1))


dlrxusdfin=peg_strictfin*(c(713)+c(12)*(-lrxusdfin(-1)+dlcpifin(-1))+c(13)*dlrxusdfin(-1)+c(14)*ddlcpifin(-
1))+(1-peg_strictfin)*(c(714)+ c(121)*(-lrxusdfin(-1)+dlcpifin(-1))+c(131)*dlrxusdfin(-1)+c(141)*ddlcpifin(-
1)  )
ddlcpifin=c(217)+peg_strictfin*(c(22)*(-lrxusdfin(-1)+dlcpifin(-1))+c(23)*dlrxusdfin(-1)+c(24)*ddlcpifin(-
1))+(1-peg_strictfin)*(c(221)*(-lrxusdfin(-1)+dlcpifin(-1))+c(231)*dlrxusdfin(-1)+c(241)*ddlcpifin(-1))


dlrxusdfra=peg_strictfra*(c(715)+c(12)*(-lrxusdfra(-1)+dlcpifra(-1))+c(13)*dlrxusdfra(-1)+c(14)*ddlcpifra(-
1))+(1-peg_strictfra)*(c(716)+ c(121)*(-lrxusdfra(-1)+dlcpifra(-1))+c(131)*dlrxusdfra(-
1)+c(141)*ddlcpifra(-1)  )
ddlcpifra=c(218)+peg_strictfra*(c(22)*(-lrxusdfra(-1)+dlcpifra(-1))+c(23)*dlrxusdfra(-1)+c(24)*ddlcpifra(-
1))+(1-peg_strictfra)*(c(221)*(-lrxusdfra(-1)+dlcpifra(-1))+c(231)*dlrxusdfra(-1)+c(241)*ddlcpifra(-1))

dlrxusdgbr=peg_strictgbr*(c(717)+c(12)*(-lrxusdgbr(-1)+dlcpigbr(-1))+c(13)*dlrxusdgbr(-
1)+c(14)*ddlcpigbr(-1))+(1-peg_strictgbr)*( c(718)+c(121)*(-lrxusdgbr(-1)+dlcpigbr(-
1))+c(131)*dlrxusdgbr(-1)+c(141)*ddlcpigbr(-1)  )
ddlcpigbr=c(219)+peg_strictgbr*(c(22)*(-lrxusdgbr(-1)+dlcpigbr(-1))+c(23)*dlrxusdgbr(-
1)+c(24)*ddlcpigbr(-1))+(1-peg_strictgbr)*(c(221)*(-lrxusdgbr(-1)+dlcpigbr(-1))+c(231)*dlrxusdgbr(-
1)+c(241)*ddlcpigbr(-1))
```

dlrxusdita=peg_strictita*(c(719)+c(12)*(-lrxusdita(-1)+dlcpiita(-1))+c(13)*dlrxusdita(-1)+c(14)*ddlcpiita(-
1))+(1-peg_strictita)*(c(720)+c(121)*(-lrxusdita(-1)+dlcpiita(-1))+c(131)*dlrxusdita(-1)+c(141)*ddlcpiita(-
1))
ddlcpiita=c(210)+peg_strictita*(c(22)*(-lrxusdita(-1)+dlcpiita(-1))+c(23)*dlrxusdita(-1)+c(24)*ddlcpiita(-
1))+(1-peg_strictita)*(c(221)*(-lrxusdita(-1)+dlcpiita(-1))+c(231)*dlrxusdita(-1)+c(241)*ddlcpiita(-1))

dlrxusdjpn=peg_strictjpn*(c(721)+c(12)*(-lrxusdjpn(-1)+dlcpijpn(-1))+c(13)*dlrxusdjpn(-
1)+c(14)*ddlcpijpn(-1))+(1-peg_strictjpn)*(c(722)+ c(121)*(-lrxusdjpn(-1)+dlcpijpn(-
1))+c(131)*dlrxusdjpn(-1)+c(141)*ddlcpijpn(-1))
ddlcpijpn=c(252)+peg_strictjpn*(c(22)*(-lrxusdjpn(-1)+dlcpijpn(-1))+c(23)*dlrxusdjpn(-
1)+c(24)*ddlcpijpn(-1))+(1-peg_strictjpn)*(c(221)*(-lrxusdjpn(-1)+dlcpijpn(-1))+c(231)*dlrxusdjpn(-
1)+c(241)*ddlcpijpn(-1))

dlrxusdnld=peg_strictnld*(c(723)+c(12)*(-lrxusdnld(-1)+dlcpinld(-1))+c(13)*dlrxusdnld(-
1)+c(14)*ddlcpinld(-1))+(1-peg_strictnld)*(c(724)+c(121)*(-lrxusdnld(-1)+dlcpinld(-
1))+c(131)*dlrxusdnld(-1)+c(141)*ddlcpinld(-1))
ddlcpinld=c(222)+peg_strictnld*(c(22)*(-lrxusdnld(-1)+dlcpinld(-1))+c(23)*dlrxusdnld(-
1)+c(24)*ddlcpinld(-1))+(1-peg_strictnld)*(c(221)*(-lrxusdnld(-1)+dlcpinld(-1))+c(231)*dlrxusdnld(-
1)+c(241)*ddlcpinld(-1))

dlrxusdnor=peg_strictnor*(c(725)+c(12)*(-lrxusdnor(-1)+dlcpinor(-1))+c(13)*dlrxusdnor(-
1)+c(14)*ddlcpinor(-1))+(1-peg_strictnor)*(c(726)+ c(121)*(-lrxusdnor(-1)+dlcpinor(-
1))+c(131)*dlrxusdnor(-1)+c(141)*ddlcpinor(-1))
ddlcpinor=c(223)+peg_strictnor*(c(22)*(-lrxusdnor(-1)+dlcpinor(-1))+c(23)*dlrxusdnor(-
1)+c(24)*ddlcpinor(-1))+(1-peg_strictnor)*(c(221)*(-lrxusdnor(-1)+dlcpinor(-1))+c(231)*dlrxusdnor(-
1)+c(241)*ddlcpinor(-1))

dlrxusdprt=peg_strictprt*(c(727)+c(12)*(-lrxusdprt(-1)+dlcpiprt(-1))+c(13)*dlrxusdprt(-1)+c(14)*ddlcpiprt(-
1))+(1-peg_strictprt)*(c(728)+ c(121)*(-lrxusdprt(-1)+dlcpiprt(-1))+c(131)*dlrxusdprt(-
1)+c(141)*ddlcpiprt(-1))
ddlcpiprt=c(224)+peg_strictprt*(c(22)*(-lrxusdprt(-1)+dlcpiprt(-1))+c(23)*dlrxusdprt(-1)+c(24)*ddlcpiprt(-
1))+(1-peg_strictprt)*(c(221)*(-lrxusdprt(-1)+dlcpiprt(-1))+c(231)*dlrxusdprt(-1)+c(241)*ddlcpiprt(-1))

dlrxusdswe=peg_strictswe*(c(729)+c(12)*(-lrxusdswe(-1)+dlcpiswe(-1))+c(13)*dlrxusdswe(-
1)+c(14)*ddlcpiswe(-1))+(1-peg_strictswe)*(c(730)+ c(121)*(-lrxusdswe(-1)+dlcpiswe(-
1))+c(131)*dlrxusdswe(-1)+c(141)*ddlcpiswe(-1))
ddlcpiswe=c(225)+peg_strictswe*(c(22)*(-lrxusdswe(-1)+dlcpiswe(-1))+c(23)*dlrxusdswe(-
1)+c(24)*ddlcpiswe(-1))+(1-peg_strictswe)*(c(221)*(-lrxusdswe(-1)+dlcpiswe(-1))+c(231)*dlrxusdswe(-
1)+c(241)*ddlcpiswe(-1))

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