THE TRIPLE-PARITY LAW

Jean-Christian Lambelet Alexander Mihailov*

University of Lausanne University of Reading

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Abstract: Scientists and epistemologists generally agree that a scientific law must be (a) relatively simple and (b) not contradicted by the available evidence. In this paper we propose and test one such law pertaining to international economics, the *triple-parity law*. It integrates three well-known equilibrium conditions: uncovered nominal interest rate parity; relative purchasing power parity; real interest rate parity. Using a cross-section of trend growth rates for 18 OECD countries in the post-Bretton-Woods/pre-EMU floating rate period (1976-1998) and employing a variety of single-equation and system estimation methods, we present robust evidence that the triple-parity law ultimately holds.

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Jean-Christian Lambelet, DEEP/HEC, University of Lausanne, BFSH1, CH-1015 Lausanne, Switzerland, + 41 (0)21 692 3481 (phone), + 41 (0)21 692 3355/65 (fax), jean-christian.lambelet@unil.ch, http://www.hec.unil.ch/jlambelet/.

* Corresponding author: Alexander Mihailov, Department of Economics, University of Reading, Whiteknights, Reading RG6 6AA, + 44 (0)118 378 5371 (phone), + 44 (0)118 975 0236 (fax), a.mihailov@reading.ac.uk, http://www.rdg.ac.uk/business/staff.asp?id=129&pid=0.

> "Human actions exhibit certain uniformities, and it is solely because of this property that they can be studied scientifically. These uniformities have another name; they are called *laws*."
>
> Vilfredo Pareto[1]

I. Introduction

Scientists and epistemologists tend to agree that, to be worthy of the name, a scientific law must be (a) general enough but compactly formulated, and (b) fully consistent, in Popper's sense (1959), with the available evidence.[2] In this paper we propose and test one such law pertaining to international economics and applicable to open market economies. We call it the *triple-parity law*. It integrates three well-known equilibrium conditions, which are shown to prevail in the long run, on average and ex post: (i) the uncovered nominal interest rate parity (UIP) condition, possibly subject to broadly interpreted cross-country financial/institutional premia; (ii) the relative purchasing power parity (PPP) condition, possibly allowing for broadly understood cross-country real/structural differentials; (iii) the real interest rate parity (RIP) condition, possibly incorporating country-specific differences of both financial and real nature. The triple-parity law thus highlights the interdependence between UIP obtaining from arbitrage in asset markets and PPP obtaining from arbitrage in goods markets, which ultimately results in a tendency towards equalization of real interest rates, i.e. RIP.

We employ a simple analytical framework and a number of straightforward yet complementary econometric techniques, to achieve robustness. With few exceptions, these methods are standard. Our contribution consists in their original application and interpretation as well as in some related novel findings. More importantly, this paper is among the few to view and test RIP as resulting –

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[1] *Manuel d'économie politique* (1966/1909, 5); emphasis in original; our translation ("regularities" might however be a better translation of *uniformités* than "uniformities").

[2] Blaug (1992, 15) characterizes the (post-)Popperian view of science as an "endless attempt to falsify existing hypotheses and to replace them by ones that successfully resist falsification".

theoretically but also empirically – from the combination of UIP and PPP. Rather than keeping to the mainstream by resorting to high-frequency time series techniques (with all their well-known imperfections and attendant controversies), or spanning an ultra long period of a century or two (at the cost of working with just a few comparable historical data sequences), we take the stand that economic laws tend to prevail in a "usual" long run, up to something like a generation. We therefore purposefully isolate our data from both short-term vagaries and too narrow sampling by relying on a cross-section of trend growth rates of relevant time series for 18 OECD countries in the post-Bretton-Woods/pre-EMU floating rate period (1976-1998). We then apply single-equation as well as system estimation methods to confront the theoretical propositions we are interested in with the statistical facts. To our knowledge, such an econometric strategy has not been pursued, in particular to formulate and test the joint determination of long-run equilibria in open market economies as embodied in the emergence of RIP out of the interaction of UIP and PPP.

The originality of our analytical and empirical approach to these classic and well-explored equilibrium conditions lies in the following *methodological features*.

(1) Most papers so far have focused on an individual international arbitrage condition – be it UIP, PPP or RIP – and on a small number of large countries taken two at a time. We use a sample of 18 industrialized countries and check – separately as well as jointly – the three basic conditions across all countries. In essence, we consider and test UIP, PPP and RIP as a system of long-run equilibrium relations. Yet whether RIP can be considered as a separate condition "in its own right" is analyzed too.

(2) The issue of the choice of a country of reference is faced squarely. This question is commonly ignored, the USA being almost always selected. By contrast, our empirical tests of the triple-parity law rely on a more intensive use of the computer involving a "shifting" cross-section of trend growth differentials whereby each economy successively plays the role of the reference country.This enables us to come up with an interpretation of the estimated intercepts in the various regressions as measuring certain comparative (dis)advantage(s) pertaining to each country.

(3) Such an econometric strategy aims at estimating long-run equilibrium conditions *directly*, filtering out what is likely to include a lot of high-frequency noise. Doing so, we sidestep a number of pitfalls linked to the usual approaches in the literature – e.g. the danger of overdifferencing, the low power of unit root tests, the difficulty of measuring country specificities in panel data methods, the somewhat arbitrary judgment commonly involved in regime-switching techniques and cointegration analyses.

(4) When the equations are initially estimated by single-equation methods, the issue of normalization/direction-of-causality as well as the error-in-data problem are addressed explicitly.

(5) When system estimation methods are applied subsequently, a number of intriguing questions arise, which have not been discussed so far in the context of joint tests of international parity relations.

The key findings can be summarized as follows.

(1) Our econometric results provide robust evidence that the triple-parity law holds in the long run, on average and ex post, up to some country-specific constants in a few cases. These findings therefore confirm – after decades of nihilistic pessimism – the direction taken by recent research to the effect that reversal to the mean does occur, eventually and when appropriately measured.

(2) The study gives proper attention to the meaning of the nonzero intercepts in the three equations comprising the law, which are interpreted as average country-specific characteristics over the sample period.

(3) Quantitative estimates of these comparative (dis)advantages are offered for all 18 countries and tested as to their statistical significance.

(4) We conclude that the joint validity of UIP and PPP implies RIP in the long run. Firstly, the

deterministic part of RIP can be analytically derived from UIP and PPP. Secondly, although it may tentatively be considered as a separate condition from an econometric standpoint, our tests show that the error term in the RIP regressions is rather a combination of the error terms in the UIP and PPP regressions.

(5) Finally, we provide estimates on how long the "long run" is, i.e. we present computations of the speed of convergence to UIP, PPP and RIP. While UIP holds even for a "medium term" of about 5 years or so, it takes – roughly – twice longer for PPP and thrice longer for RIP to emerge.

The paper is organized as follows. The underlying basic theory is summarized in section II, with references to only the most directly relevant literature; the same goes for the empirical results in section III. How our analytical approach, econometric implementation and principal findings relate to those in the closest recent research is discussed in section IV. Our conclusions are set forth in section V.

II. Analytical Framework

Consider a two-country world (A and B) where there exists "sufficient" – i.e. not necessarily perfect – mobility of capital across the border (and where capital is highly substitutable). Because of capital mobility in asset (or financial) markets, the following first arbitrage condition must hold in the long run, on average and ex post:

$$D_{A/B} = (F_A - F_B) + (I_A - I_B) + e_1. \tag{1}$$

All variables are expressed as trend growth rates (in % p.a.) over the period 1976-1998. $D_{A/B}$ is the depreciation rate of currency A with respect to currency B as measured by the spot exchange rate.[3] I_A

[3] Defined as the price of one unit of B's currency in terms of A's currency.

and I_B are long-term interest rates; their exact definition and measurement will be discussed in section III.1. e_1 is a disturbance term, due to a shock process affecting (1). F_A and F_B are country-specific financial and institutional characteristics (or factors). We interpret their differential, $F_A - F_B$, estimated as the regression intercept, in the sense of some general financial/institutional disadvantage of A relative to B (or, inversely, some advantage of B relative to A). The most straightforward and well-known interpretation in the literature is to reduce our "financial comparative disadvantage" term to a risk premium. But we argue that this interpretation is too narrow because the term in question may include considerations other than its usually proposed risk components such as the political or sovereign risk, the default one and that due to financial market regulation and capital controls. Suppose, for example, that country A is more discreet in tax matters – i.e. less inclined to cooperate internationally – than country B; and/or suppose country A has a banking secrecy law, but not country B. *Ceteris paribus*, the "financial factor" or "institutional environment" characterizing country A, F_A, will then be more favorable than that for country B, F_B. As these examples show, our financial/institutional differential could also be a safety premium, so that "comparative disadvantage/advantage" is a more general and hence better description. Section III will provide and test estimates of the average "F"-factor differentials for each of 18 industrialized economies over the 1976-1998 period. Equation (1) is thus the uncovered nominal interest rate parity (UIP) condition, in its ex post formulation and allowing for a broadly interpreted financial/institutional differential (in case the latter comes up statistically significant, as it does for a few small countries in our sample).

If, like capital, goods and services are "sufficiently" mobile (in addition to being highly substitutable), arbitrage in goods markets also ensures that the following second equilibrium condition will be fulfilled in the long run, on average and ex post:

$$D_{A/B} = (R_B - R_A) + (\Pi_A - \Pi_B) + e_2. \tag{2}$$

Π_A and Π_B are national inflation rates (in % p.a.) and e_2 denotes shocks to PPP. R_A and R_B are country-specific real or structural characteristics (or factors). As with our financial/institutional characteristics in (1), we do not propose a precise theory for their real/structural analogues in (2). We would rather interpret the differential, $R_B – R_A$, as some general real/structural advantage of A relative to B (or, inversely, disadvantage of B relative to A).[4] (2) is thus a version of the familiar purchasing power parity (PPP) condition, in its relative form and with explicit allowance for an inter-country structural or real-economy differential. The most straightforward and well-known analogy in the literature is to reduce the latter to a productivity differential, as implied by the Balassa-Samuelson effect. But we argue that such an interpretation is too restrictive, especially given some controversies in the more recent research on the link between relative national price levels and the real exchange rate (RER).[5] Our concept of "structural advantage" or "real differential" must be understood here in the broadest sense, as it may be due to other underlying causes besides the Balassa-Samuelson effect or, synonymously, the cross-country productivity differentials often mentioned in the earlier literature.[6] More precisely, later studies[7] have interpreted deviations from PPP (and hence departures of the RER from the constant level implied by relative PPP) as originating in more than technology-related productivity differentials. MacDonald and Stein (1999) and Juselius and MacDonald (2004) have notably suggested a much more complicated picture where many potential reasons could account for slow adjustment to PPP. In essence, the persistence of the deviation from PPP is due, to quote Juselius and MacDonald (2004, 4), "to the existence of important real factors working through the current account, such as productivity differentials, net foreign asset positions and fiscal imbalances". Terms of

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[4] Notice that the R factors are reversed in (2): R_B comes before R_A. This is because they are defined positively, by analogy with the productivity differential, or Balassa (1964)-Samuelson (1964) effect, argument in the earlier literature, while the F factors in (1) are defined negatively, again similarly to the risk premium term in traditional international finance; that is, a large F is a "bad", but a large R is a "good".

[5] Measured according to its PPP-based version, i.e. as the ratio of national price levels converted to a common currency.

[6] See Officer (1976), Hsieh (1982), Dornbusch (1987), Marston (1987), De Gregorio, Giovannini and Wolf (1994), and Canzoneri, Cumby and Diba (1999).

[7] See in particular Chinn and Johnston (1999), Begum (2000), MacDonald and Ricci (2001), and Lee and Tang (2003).

trade (ToT) effects should also be extant among these real factors through their influence on the current account and net foreign assets (NFA). The same goes for changes in tastes (or preferences) as they may shift demand for home products relative to foreign ones and thus affect the current account and the NFA position. We would equivalently refer to these real factors as structural factors, following Dornbusch (1987), in the sense of real disturbances that change equilibrium relative prices and thus cause systematic departures from PPP.

Combining (1) and (2), we get:

$$(F_A - F_B) + (I_A - I_B) = (R_B - R_A) + (\Pi_A - \Pi_B) + (e_2 - e_1), \tag{3}$$

or equivalently

$$(I_A - I_B) = [(R_B - R_A) - (F_A - F_B)] + (\Pi_A - \Pi_B) + (e_2 - e_1), \tag{4}$$

or still

$$(I_A - \Pi_A) = [(R_B - R_A) - (F_A - F_B)] + (I_B - \Pi_B) + (e_2 - e_1), \tag{5}$$

which are versions – less familiar in the way we have written them – of the real interest rate parity and third condition, RIP, with explicit allowance for both financial/institutional and real/structural differentials. It is not always realized that if (1) and (2) hold, (3)-(5) must too. In other words, the nominal UIP condition and the relative PPP condition imply, when taken together, that real interest rates must also be equalized internationally in the long run, on average and ex post. Whether (3)-(5) might nevertheless be considered as a separate condition "in its own right", as has been argued in part of the literature, will be examined later.

Combining (1), (2) and (5), and ignoring the error terms, we get:

$$\boxed{\underbrace{\mathbf{D_{A/B} - (F_A - F_B) - (I_A - I_B)}}_{\substack{\text{Nominal UIP allowing} \\ \text{for F-differentials}}} = \underbrace{\mathbf{D_{A/B} - (R_B - R_A) - (\Pi_A - \Pi_B)}}_{\substack{\text{Relative PPP allowing} \\ \text{for R-differentials}}} = \underbrace{\mathbf{[(R_B - R_A) - (F_A - F_B)] + (I_B - \Pi_B) - (I_A - \Pi_A)}}_{\substack{\text{RIP allowing} \\ \text{for F- and R-differentials}}} = 0, \qquad \mathbf{(6)}}$$

which is the *triple-parity law*, to be tested below both by individual equations and as a system. Note that this law is entirely specified in terms of average rates of change over time[8] and is thus compatible with any number of different combinations of depreciation rates and interest and inflation differentials.

III. Empirical Implementation

III.1 Data: Sources, Definitions and Transformations

The following data, all from the IMF's *International Financial Statistics* (IFS) or from the OECD's *Main Aggregates* national accounting publications, were collected for each of 18 industrialized countries[9] over the 1976-1998 period:[10] the average annual values of the nominal spot exchange rate

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[8] Interest rates are also rates of change since they indicate the rate at which an asset yields a return over time.

[9] I.e. all countries for which (a) complete and reasonably homogeneous time series could be obtained for all variables and (b) a "sufficiently" high degree of capital and goods mobility and substitutability could be presumed to exist over most of the sample period. These are identified (later in Figure 6) by the country tags one sees on automobiles: AUS = Australia, A = Austria, B = Belgium, CDN = Canada, DK = Denmark, SF = Finland, F = France, D = Germany, I = Italy, J = Japan, NL = Netherlands, NZ = New Zealand, N = Norway, E = Spain, S = Sweden, CH = Switzerland, GB = United Kingdom (UK), and USA = United States of America.

[10] The starting year was determined by the availability of sufficiently homogeneous series. 1976 is, of course, three years after the final breakdown of the Bretton-Woods system and, moreover, coincides with the official beginning of generalized floating. Although the triple-parity law also holds under a system of fixed exchange rates, a period of floating currencies makes for a much richer sample.

vis-à-vis the US dollar; the average annual interest rate on long-term government bonds; the average annual levels of both the CPI and the GDP deflator. When measuring the nominal interest rate our objective was to select homogeneous bonds with a long and uniform maturity (say, 10 years), but this proved unfeasible.[11] Accordingly, this series could be the one most likely to suffer from an error-in-data problem.

The difficulty here is that no international standard has yet been adopted to unify various national practices when measuring and aggregating the yields on long-term government bonds. The time series used, namely those in the IFS, therefore reflect various country-specific definitions. Broadly speaking, there are three groups of countries. A first group, such as Austria, Germany and Japan, reports the average yield on all government bonds.[12] Second, many other countries select a subset of all government bonds, but this subset is not defined everywhere in the same way. Australia, for instance, reports the assessed secondary market yield on 2- to 10-year bonds; Spain, France and Sweden report the average yield of bonds with a maturity longer than 2, 5 and 9 years, respectively;[13] the same rule applies, but for maturities in excess of 10 years, to Canada, Belgium[14] and Italy;[15] as to Switzerland, it followed a similar principle up to 1999, but with a 20-year maturity as an upper bound. Countries in a third group report the annual yield of some benchmark long-term government bond with a fixed maturity, such as 5 years for Denmark, Norway[16] and New Zealand, 10 years for Finland,[17] the

[11] Fujii and Chinn (2000) were able to employ two apparently more homogeneous series, but for the G7 countries only.

[12] In the case of Austria, bonds that are issued but not redeemed are included in a weighted average; in the case of Japan, only the bonds that are called "ordinary" enter the definition.

[13] In Sweden, the definition has been modified frequently, with the lower maturity bound set at 15 years before 1980, 10 years throughout 1980-1993, and 9 years since 1994.

[14] Before 1990, Belgium considered instead the weighted average of the yield of all government bonds that had a maturity longer than 5 years and a yield of 5-8% p.a.

[15] Italy reports end-of-month yields.

[16] Yield to maturity.

Netherlands and the USA, and 20 years for the UK.

For each country, the mean or trend depreciation/appreciation and inflation rates over the sample period were calculated by regressing the logarithm of the original series on an annual time index. Consequently, they are continuous rates. For consistency's sake, the long-term interest rate for each country was put on a continuous-compounding base too,[18] and each national series' mean value was taken (Table 1).

[Table 1 about here]

In a first step, average depreciation/appreciation, inflation and interest rate differentials were calculated with respect to the USA (Table 1, again). In a second step, such differentials were computed taking in turn each of the other economies as the reference country.

The sample is thus a *cross-section* of average long-run growth rates of *time series*. It consists of 18 (shifting) observations,[19] which may seem a rather small sample. But there is a difference between the sheer size of a sample, as measured by the number of observations, and its information content. We believe that our sample "packs" a very large amount of information, epitomizing as it does the often very different macroeconomic choices and functioning of no less than 18 industrialized countries, each over a period of no less than 23 years.

It could be argued that our procedure, i.e. taking a cross-section of trend growth rates of various annual national time series, implies that a lot of information about short-term dynamics is lost. Here we are

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[17] Since the respective time series for Finland was not available in the IFS for all years in our sample, we used instead the Finnish 10-year government bond yield, kindly provided by Erkki Kujala, Bank of Finland, to whom we owe our thanks.

[18] Applying the following formula for country i: $I_{i,t} = \log(1 + IR_{i,t})$ where $IR_{i,t}$ is the reported interest rate.

[19] "Shifting" because all 18 countries were used in turn as the reference country.

however *solely* interested in estimating a set of long-term equilibrium conditions, with a sample including as many relatively homogeneous countries as possible, and we do not want the estimation to be perturbed by short-term vagaries. This is also why only long-term interest rates were taken and why annual data had to be used. Lastly, we are focusing on *realized* outcomes and not ex ante relationships, but this is consistent with our long-run view in the study and is standard in the literature.[20]

III.2 Single-Equation Estimation Methods: OLS, WLS and ODR Results

All three parity conditions will be tested, first individually and then as a system, even though RIP was derived above from the other two. This testing is however not as straightforward as it might seem, because all conditions rest on an *arbitrage* mechanism. Taking, e.g. UIP (1),

$$D_{A/B} = (F_A - F_B) + (I_A - I_B) + e_1 ,$$

it is not clear – selecting the USA as the reference country – whether one should empirically estimate, as is most often done, an equation of the form

$$D_{i/USA} = a_1 + b_1(I_i - I_{USA}) + e_{1,i} \qquad (i = \text{country}), \tag{7}$$

or whether one should rather estimate the reverse relationship[21]

$$(I_i - I_{USA}) = \alpha_1 + \beta_1 D_{i/USA} + \varepsilon_{1,i} , \tag{8}$$

which will yield different numerical estimates.[22]

[20] See Fujii-Chinn (2000, 6), Obstfeld-Taylor (2000, 2), and Sekioua (2005, 7-8), among others.

[21] One should compare estimates of a_1 in (7) with $-\alpha_1/\beta_1$ in (8) and of b_1 in (7) with $1/\beta_1$ in (8), as we do later.

In other words, the direction of causality – and hence the choice of the dependent variable – is not a straightforward question when *arbitrage* is at work.[23] Following the discussion in Maddala (1992, 74-76, 447-472), we shall consequently estimate in all cases both an equation like (7), the *direct* regression, and one like (8), the *reverse* regression, the results to be interpreted – according to the same author – as "bounds" around the true value of the parameters. Friedman and Schwarz (1982, 173 fn. 28 and 225 fn. 18) seem to have pioneered this approach, using coefficient estimates from direct and reverse regressions as "upper and lower limits", but in another context.

Orthogonal distance regression (ODR) rather than ordinary least squares (OLS) and weighted least squares (WLS) would seem an obvious choice in such circumstances. A further reason, in addition to the direction of causality argument above, is that both variables are likely to be measured with error: in OLS and WLS, there is no symmetry in the sense that the error is minimized only in one direction, that of the dependent variable. ODR however fits the slope in a symmetrical way, so that the role of both variables in a simple regression is the same. For standardized data with dependent and independent variable of identical scale, the ODR line coincides with the first principal component. Orthogonal (distance) regression appears to be quite a popular method in other sciences, such as medicine or engineering, where it is sometimes claimed that it allows a more general treatment of the error-in-data problem; yet it does not really sidestep the problem, since the ratio of the measurement error variances must be supplied extraneously.[24] Unfortunately, orthogonal estimators have infinite higher moments[25] (at least in the case of linear models[26]), so that no hypothesis testing can be done and no confidence

[22] At least with *single-equation* estimation methods on which this section concentrates. *System* estimation results are given in section III.4.

[23] Another, separate criterion for the choice of the dependent variable is to select that variable which is most likely to suffer from an important error-in-data problem.

[24] For example, see Ammann-van Ness (1988).

[25] See Anderson (1976, 1984) as quoted in Boggs *et al.* (1988, 172).

[26] See Boggs-Rogers (1999).

intervals can be constructed. Nevertheless, we shall also supply ODR estimates, which will naturally lie between the two bounds mentioned above; moreover, as may have been expected and as shown in tables 2, 3 and 4 further down, the measure for goodness of fit of the ODRs we computed, ϕ^2, is generally higher than the adjusted r^2 for the respective OLS regressions.[27]

Table 2 lists the OLS, WLS and ODR results for the nominal UIP condition inclusive of our measure of average financial disadvantage, the reference country being the USA. Figure 1 gives an impression of the sample.

[Table 2 about here]

[Figure 1 about here]

As the table shows, the data do not reject H_0: b=1, the theoretically expected value: all confidence intervals contain this value for the slope of the UIP regressions. In other words, the nominal UIP condition stands verified on average, in the long run and ex post, except for a (country-of-reference-specific) statistically significant intercept in some cases. It was argued that the estimated intercept includes – but is not necessarily equal to – the country-specific risk premium. Since this is a relatively complex matter, we postpone further discussion to section III.3. As to the various point estimates of b, those resulting from the X-on-Y (i.e. reverse) regressions may be here preferable to those from the Y-on-X (i.e. direct) regressions, since the interest rate differentials are more likely to suffer from an error-in-data problem than the depreciation differentials. Be that as it may, it is striking that the central ODR point estimate of b is almost exactly unity. In a pure cross-section context and with a sample of

[27] We have used a simple ODR estimation *Gauss* program of our own, based on an algorithm in Malinvaud (1970, 9-13). By construction, the ODR line corresponds to the "principal component" of the scatter of points for the case of a linear relationship between two variables. This program is available on request.

18 observations, goodness of fit measures of 0.8–0.9 would seem rather comforting too.[28] Finally, note that no joint Wald test is relevant here: while theory tells us that E(b)=1, there is no a priori expectation about the value of the constant.[29]

Table 3 lists our results for the relative PPP condition, again taking the USA as the reference country. Figure 2 gives a visual impression of the sample when the GDP deflators are used to measure inflation differentials, and Figure 3 when the CPIs are taken instead. Again, the data do not contradict H_0: b=1, the theoretically expected value. The PPP condition in its relative form thus also holds empirically in the long run, on average and ex post (up to a statistically significant intercept in some cases, to be discussed in section III.3).

[Table 3 about here]

[Figure 2 about here]

[Figure 3 about here]

Table 4 lists our results for the RIP condition, still taking the USA as the reference country.[30] Figures 4 and 5 give a visual impression of the sample. Here too, the theoretically expected value for the slope, H_0: b=1, always lies within the 95 and 99% confidence intervals, so that RIP is also consistent with the data. The estimated RIP intercepts, never statistically significant at all conventional (1%, 5% and 10%) levels for the regressions with respect to the USA reported in Table 4, will be discussed in section III.3.

[28] Figure 1 also shows that the estimated coefficients are not unduly influenced by outliers. This has been confirmed by a full "influential analysis" applied to all equations (Kennedy, 2003, 373-374, 379-380; Hayashi, 2000, 21-23).

[29] The irrelevance of a Wald test in this context seems to be often overlooked in the empirical literature.

[30] Notice that the estimated equation (on top of Table 4) is our RIP specification (4), which is differently specified from the usual Fisher equation (5). In the latter case, taking the USA as the reference country would mean that, depending on the regression specification, the explanatory or explained variable ($I_{USA} - \Pi_{USA}$) is … a constant.

[Table 4 about here]

[Figure 4 about here]

[Figure 5 about here]

It was mentioned above that it is not clear from the literature whether real interest rate parity is a separate condition "in its own right", which should consequently be tested as such. This issue may hinge on which type of agents is doing the arbitrage. For an individual investor residing permanently in a given country, and hence based in that country's currency, the nominal UIP condition is clearly of the essence: she will compare the nominal rate of return on home and foreign assets allowing for the expected path of the nominal exchange rate; i.e. expected inflation in that investor's country or abroad will not affect her choice. That may be different for investors who are very mobile internationally and who may therefore be interested in getting the same real returns wherever their investments and they themselves happen to be located at any given time.[31] Alternatively, it is conceivable that multinational firms with production facilities and shareholders in many different countries will want to manage their investments, whether financial or material, in such a way that the real rate of return in the different countries is ultimately equalized.[32] We shall return to this issue in sections III.3 – III.6.

The triple-parity law therefore says that, except for inter-country financial/institutional and real/structural differentials, the real interest rate should tend to become equalized, on average, in the long run and ex post. This is illustrated in Figure 6. It is striking that average real long-term interest rates are very closely bunched around the 4% p.a. value for ten economies out of eighteen, including all the larger ones (except Australia and Canada). Yet country-specific factors seem important for some smaller countries (shaded area in Figure 6, most notably Switzerland), about which more in the

[31] On this, see Marston (1997).

[32] Obstfeld-Taylor (2000, 1) put it succinctly: "International real interest rate equality would hold in the long run in a world where capital moves freely across borders and technological diffusion tends to drive a convergence process for national production possibilities."

next subsection where the significance of these deviations will be tested.

[Figure 6 about here]

III.3 The Country-Specific Intercepts: Interpretation as Comparative (Dis)Advantages

So far, the USA has been taken as the reference country, which is an arbitrary although perhaps natural choice. If another country is selected instead, it will change the estimated intercepts, but *not* the estimated gradients and their associated statistics, which remain exactly the same as in the above tables. In other words, the choice of the reference country makes no difference for the estimated b coefficients, but it does for the constants: given the way the different variables are defined, changing the reference country is similar to changing the measurement unit.

Furthermore, including the zero differentials for the reference country in all our "shifting" regressions has an interesting econometric consequence. Take any equation where the USA is the reference country. Given that both the right-hand side (RHS) and left-hand side (LHS) variables are zero for the USA, the estimated constant must by necessity be equal to the residual for the USA, with the sign reversed.[33] The same obtains when each one of the other countries is taken in turn as the reference country. Consequently, the various estimated intercepts and the various reference-country *own* residuals are one and the same thing, and we do not need to show them separately.[34] Table 5 lists them for the first equation, i.e. UIP.

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[33] Given $y_i = \hat{a}_1 + \hat{b}_1 x_i + \hat{e}_{1,i}$, $y_i = x_i = 0$ means $\hat{a}_1 = -\hat{e}_{1,i}$.

[34] Take any one of our cross-section equations: although the estimated constant and the reference-country own residual are one and the same thing in that given equation, this leaves the residuals for the other countries. We have closely scrutinized the residuals for all equations and have found no indication of non randomness. E.g. it is never the case that large countries tend to have small residuals, and small countries large residuals, etc.

Each estimated intercept includes – but is not necessarily equal to – the average country-specific financial/institutional premium. This is so because an estimated constant can be non zero for three non mutually exclusive reasons: (1) $E(e_i)$ = some constant $\neq 0$: the measurement errors affecting the dependent variable include a systematic bias; (2) even when $E(e_i)=0$, genuinely random shocks will in general not average out to zero in any finite sample; (3) there is a non zero average financial/institutional disadvantage for the country under consideration. Given that our cross-section is made up of trend growth rates for a 23-year period, reason (2) is unlikely to be important. Reason (1) could however be important if the dependent variable is affected by measurement errors with a sizable bias. This caveat should be borne in mind when interpreting Table 5.

[Table 5 about here]

The results in Table 5 are somewhat sensitive to the direction of regression and to the estimation method (OLS vs. WLS), which was confirmed by examining the correlation matrix of the alternative intercept estimates. However, looking across the various $\hat{a}_1$'s for each country, and admitting as a rule of thumb that by and large there are "strong indications" of an economically meaningful intercept when the $\hat{a}_1$'s are statistically significant in at least three cases out of our four estimates, we find such evidence for four small countries only: Switzerland, Sweden, Finland and Denmark. A negative intercept is indicative of a risk discount differential for the country under consideration or, more generally, of a financial/institutional comparative advantage, as argued previously. This is best seen when considering the *reverse* relationship: a negative constant means that, for a given rate of depreciation/appreciation, and for a given level of foreign nominal interest rates, said country enjoys domestic interest rates that are lower than would be expected normally.[35] The results in Table 5 thus

[35] Taking the *direct* relationship, a small constant means that the country benefits from a stronger (i.e. more rapidly appreciating) currency than would be expected given its interest rate level relative to foreign interest rates. Bear in mind that we are considering a long-term equilibrium situation, so that no competitiveness problems arise due to a currency which appreciates more rapidly than one would normally expect. This means that, for a

suggest that Switzerland and Sweden have likely benefited from an important international comparative advantage whereas Finland and Denmark appear to have been at a sizable disadvantage. Why this should be so for these four countries will be discussed further on when examining the constants in the RIP equations.

[Table 6 about here]

We now turn to the estimated constants in the PPP equations, reported in Table 6. The previous caveats about the three possible reasons for non zero intercepts should be kept in mind so that not too much should be read into our results. With that proviso, a significant negative constant for a given country in Table 6 is indicative of real/structural disadvantage (in a comparative perspective), and vice-versa – be it a constant factor or an average over the sample period. Applying the same rule of thumb as previously, four countries appear to exhibit real/structural disadvantage: Australia, Canada, Norway and Sweden, all of them important producers and exporters of primary commodities and raw materials. At the other end, real/structural advantages have been enjoyed in the UK, Spain and Italy. The explanation may be that these three countries have undergone especially rapid modernization in the sample period.

Similar general comments apply to our results in Table 7, i.e. the intercepts in the RIP equations.[36] Using the same rule of thumb as above, we now find that six countries appear to be at a statistically significant *overall* comparative advantage (Switzerland, the UK) or disadvantage (Belgium, Finland, Denmark and Australia).

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given volume of exports, the country can import more cheaply from abroad without running into balance-of-payments problems.

[36] In Table 7, the sign of the estimated intercepts has been reversed (for the equations used, see top of Table 4) so as to make a small (i.e., negative) value correspond to a comparative advantage, as was the case for the nominal interest rate equations.

[Table 7 about here]

Looking back at equation (4)-(5), it is seen that the constant in the RIP equation is equal to the real/structural differential in the PPP equation minus the financial/institutional differential in the nominal UIP equation: a country will enjoy a comparative *real* interest rate advantage (or discount) if its real/structural advantage (traditionally interpreted as productivity growth differential) is larger than its financial/institutional disadvantage (traditionally interpreted as nominal interest rate risk premium) – an interesting proposition in itself. This also affords us a way to check whether the estimated country-specific constants in the RIP equations (Table 7), which we shall call the *direct* estimates of the "real interest rate (RIR) discounts", are consistent with the differences between the estimated real/structural advantages (Table 6) and the estimated financial/institutional disadvantages (Table 5), these differences to be dubbed here the *indirect* estimates of the RIR discounts. The direct and indirect estimates were actually quite close, being highly correlated. It is also worth noting that none of the differences between them was larger than half a percentage point.

It might be tempting to argue that if real interest rate equalization is due to a "special" class of arbitraging agents, be they investors or firms (as discussed above), we should rather expect the direct estimates to be different from the indirect ones because these agents' perceived RIR discounts/premia – i.e. financial/institutional and real/structural (dis)advantages – may be different from that of the other ("non special") investors or firms. But this ignores that all agents operate on the same (global) financial and goods markets where their interactions result in average market-wide discounts/premia. No matter how we look at it, the deterministic parts of equations (1) and (2) *necessarily* imply the deterministic part of equation (3)-(4)-(5): if the UIP and PPP equilibrium conditions hold, the RIP condition must hold too. It follows that real interest parity cannot possibly be a separate condition "in its own right". However, the issue takes on another meaning if we allow for the possibility that the arbitrage activities of these "special" agents may result in separate *shocks*, e_{RIP}, so that equation (3) should be rewritten as:

$$(F_A - F_B) + (I_A - I_B) = (R_B - R_A) + (\Pi_A - \Pi_B) + \underbrace{(e_2 - e_1) + e_{RIP}}_{= e_3}. \qquad (9)$$

Under these conditions, it is *econometrically* justified to test RIP as a separate condition, which we – and others – have done. Furthermore, it is also possible that the existence of this special class of arbitrageurs will reinforce and speed up the realization of the UIP and PPP conditions.

Summing up our results for both the nominal and the real interest rate parity conditions, we can conclude that only Denmark, Finland and Switzerland[37] appear to constitute significant anomalies on *both* counts. Why should that be so? E.g. why should Finland and Denmark appear to be at a significant disadvantage on both the nominal and the real interest rate count whereas Switzerland would seem to enjoy a significant comparative advantage? In the case of Switzerland, the explanation is most likely to be found in that country's reputation for political, economic and financial stability, its efficient financial sector – as well as, perhaps, its banking secrecy law and its status as an international tax heaven (although this interpretation has been rejected in at least one study[38]). In the case of Denmark, its high average RIR premium appears to be linked with high real interest rates in the 1980s and in the first half of the 1990s when the Danish economy was very inflationary and the crown under constant attack; but the problem seems to have been solved in more recent years. As to Finland, its high average RIR premium may be due mostly to disruptions in the late 1980s and early 1990s following the collapse of the Soviet Union, which used to absorb a fair share of Finish exports; but there too the problem seems to be on the mend.

[37] Switzerland being so special is not strange at all: Koedijk-Tims-Van Dijk (2005), for instance, have recently found similar conclusions in their PPP tests for the Euro area, where the only exception has been the Swiss franc.

[38] See Commission pour les questions conjoncturelles (2003). Here is not the place to go into the pros and cons of the latter two institutional factors, except maybe to point out that if there are cons (e.g. both institutional factors may be abused by non Swiss tax evaders), there are also pros (e.g. protection of the private sphere and safeguard against extortionate national tax laws).

III.4 System Estimation Methods: SUR and FIML Results

System estimation seems more appropriate in our context, since we regard the triple parity as a joint law. System methods utilize more information, namely across equations, in estimating the coefficients of interest and are more efficient. We applied two alternative and complementary such techniques, namely Zellner's (1962) seemingly unrelated regressions (SUR) estimator and full information maximum likelihood (FIML), with Marquardt's algorithm. In particular, SUR is relevant under the assumption that the error terms in the triple-parity equations are contemporaneously correlated, which is likely; or when the explanatory variables are not the same, which is the case in our system estimation by pairs of parities as reported below. If, on the contrary, the errors in the three equations are not correlated but are independent, we can replace – or rather complement – our SUR estimates with simultaneous equations models (SEM). Assuming – and also checking[39] – that these disturbances are normally distributed, a particular version of SEM that is commonly applied is FIML. Complementing SUR by FIML estimates ensures robustness for our results under alternative underlying assumptions as to the shocks to UIP, PPP and RIP. Moreover, FIML has the nice property of invariance (see, for instance, Hayashi 2000, 534). I.e. the question does not arise as to how the equations are to be normalized, and hence neither does the direction-of-causality issue. In the present case, this method however raises a number of intriguing questions. To briefly discuss them, let us simplify our notation and rewrite our three basic relationships as a system:

$$DUSA_i = c(1) + c(2)*IUSA_i + e_{1,i}, \qquad (10)$$

i.e. the UIP equation, with $DUSA_i$ = average depreciation rate of country i's currency with respect to the dollar, and $IUSA_i$ = i's average nominal interest differential vis-à-vis the USA,

$DUSA_i = c(3) + c(4)*\Pi USA_i + e_{2,i}$, (11)

i.e. the (relative) PPP equation, with ΠUSA_i = country's i's average inflation rate differential with respect to the USA,

$IUSA_i = c(5) + c(6)*\Pi USA_i + (e_{1,i} - e_{2,i}) + e_{RIP,i}$, (12)

i.e. the RIP equation.

Obviously, c(1), c(3), (c5) correspond, respectively, to the slopes a_1, a_2, a_3, and c(2), c(4), c(6) to b_1, b_2, b_3. Assuming that these three equations represent three independent arbitrage mechanisms, and in particular that the third one is not simply the result of combining the first two (but also includes $e_{RIP,i}$ so that $e_{3,i} = (e_{1,i} - e_{2,i}) + e_{RIP,i}$ as per the preceding discussion), they make up a *closed system*: all variables (= differentials) are endogenous, there are no exogenous variables which could be used as instruments, and hence no reduced form. Consequently, the FIML method cannot be applied to our full system of three equations. If we want to use FIML, we must instead take them two by two, for a total of three combinations.

Doing this however reintroduces the question of normalization, and hence that of the direction of causality. If, for example, we try to apply FIML to (10) and (11) as normalized above, i.e. with the same dependent variable, the algorithm breaks down. It is easy to see why: both equations have only one exogenous variable on their RHS, so trying to estimate them by FIML is pointless. The system (10)-(11) must therefore be renormalized, which can be done in two ways:

[39] The null of normality of the residuals from the direct and reverse UIP, PPP and RIP OLS regressions could not be rejected by the Jarque-Bera tests we performed at the 10% significance level for none of the total of 10 specifications (including PPP and RIP being alternatively estimated from CPIs or GDP deflators).

$$DUSA_i = c(1) + c(2)*IUSA_i + e_{1,i}, \tag{13}$$

$$\Pi USA_i = -[c(3)/c(4)] + [1/c(4)]*DUSA_i - [1/c(4)]*e_{2,i}, \tag{14}$$

or

$$IUSA_i = -[c(1)/c(2)] + [1/c(2)]*DUSA_i - [1/c(2)]*e_{1,i}, \tag{15}$$

$$DUSA_i = c(3) + c(4)* \Pi USA_i + e_{2,i}. \tag{16}$$

Both systems are recursive; or they are only triangular if the two e's are not independent, as might well be the case. Let us consider the two systems' deterministic parts. In the first one, for example, $IUSA_i$ is exogenous and determines $DUSA_i$ via (13), and then ΠUSA_i is determined by $DUSA_i$ via (14).[40] Since (13) comprises no endogenous variable on its RHS, the FIML point estimates of c(1) and c(2) will be exactly the same as those given by OLS (although the standard errors will be different). The FIML point estimates of c(3) and c(4) in (14) will however be different from the OLS ones. Given that the deterministic parts of (13)-(14) make up a triangular system, these FIML-estimated coefficients will also be identical to the SUR estimates (although the standard errors will again be different). The situation is reversed for the system (15)-(16). The FIML point estimates of c(3) and c(4) will now be identical to the OLS ones, while those for c(1) and c(2) will be SUR/FIML point estimates different from the OLS ones. Consequently, our SUR/FIML point estimates will be those given by (15)-(16) for c(1) and c(2), and those given by (13)-(14) for c(3) and c(4), and we need not reproduce the OLS point estimates again.

The same reasoning holds for the other two combinations of equations (10)-(12) taken two by two.[41] Table 8 gives the results for all three combinations.

[40] Notice that, by the order condition, both equations are exactly identified.

[41] Combination (10) and (12) can be estimated as it is because the system is triangular.

[Table 8 about here]

Comparing these results for any given relationship (say, UIP), it is seen that the SUR/FIML estimated coefficients are not affected by whichever combination is selected, but their standard errors are. The normalization or direction-of-causality issue is therefore irrelevant on the first count, but not on the second one. Furthermore, it is seen that the SUR method leads to t-statistics that are noticeably higher than those from FIML, which is as expected. It is true that the RHS variables in our SUR system are not strictly exogenous but rather predetermined endogenous, so that their use as regressors may lead to some bias in the reduced-form estimates. Yet Kennedy (2003, 192) notes that these estimates should nevertheless be asymptotically unbiased, assuming the errors are not autocorrelated, and that all estimators in a structural SEM context are anyway biased so that the choice among them is based on their asymptotic properties. Comparing the results in Table 8 with the OLS ones in tables 2, 3 and 4 also shows that using system estimation methods leads to a significant improvement:[42] the SUR/FIML point estimates lie between the bounds defined by the direct and reverse OLS regressions and they are actually close to the "central" ODR point estimates; on top of that, standard errors and t-statistics are now available, which was not the case for the ODR results. To sum it up: the SUR/FIML methods afford even stronger evidence in favor of the triple-parity law.

As stated above, the three-equation system (10)-(12) cannot be estimated as such by FIML, because it is a closed one, but it can be estimated by SUR if (11) is renormalized so as to determine ΠUSA_i, thus making the deterministic parts of (10)-(12) a fully triangular system – see Table 9, which also gives the OLS results for comparison purposes.

[Table 9 about here]

[42] There is however a (small) price to be paid: the adjusted r^2's are lower with SUR/FIML than with OLS; but they are only slightly lower and remain comfortingly high in a cross-section context.

It is seen that for the UIP and PPP equations, but not for the RIP one, the SUR point estimates of the slopes are closer to the theoretically expected values than the OLS point estimates. The SUR-produced t-statistics are also noticeably larger in absolute value than the OLS ones, i.e. the SUR point estimates are better assured, while the adjusted r^2's are only slightly smaller. As the standard errors show, the basic E(b)=1 hypotheses stand up under SUR too. Finally, the SUR full-system results in Table 9 are also distinctly better than the SUR/FIML two-equations-at-a-time results in Table 8, in the sense that the standard errors for the slope coefficients are smaller.

III.5 Is RIP Independent of UIP and PPP in the "Long Run"? Econometric Tests

We now return to the question whether RIP is a condition in itself, separate and independent of the arbitrage forces acting jointly through UIP and PPP. One way to answer this question, still unsettled in the literature, is to examine our cross-section regression residuals. To that end, we computed the indirect estimates of the residuals for the RIP equation, i.e. those implied jointly by the residuals of the UIP and PPP equations, and then compared them with the direct ones, i.e. the directly estimated in the RIP equation. The magnitudes of the indirect and direct measures for the RIP regression residuals were very similar. With a few exceptions, the difference between them was less than 10% of the value of the average of the direct and indirect estimates by country; and the correlation coefficient between the direct and indirect estimates was very high, 0.98: the data did not thus reject the hypothesis that the indirect and direct RIP residuals are, in essence, the same vector.

We further computed the variance of the indirect RIP residuals, using the standard formula $\mathrm{Var}(\hat{\mathbf{e}}_1 - \hat{\mathbf{e}}_2) = \mathrm{Var}(\hat{\mathbf{e}}_1) + \mathrm{Var}(\hat{\mathbf{e}}_2) - 2\mathrm{Cov}(\hat{\mathbf{e}}_1, \hat{\mathbf{e}}_2)$, where $\hat{\mathbf{e}}_1$ stands for the vector of residuals in the UIP regression and $\hat{\mathbf{e}}_2$ for that in the PPP regression. This indirectly computed variance was 1.3076, almost equal to the variance of the residuals from the direct RIP regression, $\mathrm{Var}(\hat{\mathbf{e}}_3) = 1.2591$, where $\hat{\mathbf{e}}_3$ denotes the residuals from the RIP regression. Lastly, we computed the 95% confidence interval of the variance of the residuals in the direct RIP regression, using the less well-known formula for (a two-sided test from) a χ^2 distribution,[43] RSS/h < Var($\hat{\mathbf{e}}_3$) < RSS/ℓ, where RSS stands for the regression sum of squares in RIP. For the 5% significance level and with 17 degrees of freedom the tabulated values are ℓ = 7.56 and h = 30.2 so that with RSS = 22.67, we obtain 0.7507 < Var($\hat{\mathbf{e}}_3$) < 2.9987. Our indirect estimate of the variance of the RIP residuals, Var($\hat{\mathbf{e}}_1 - \hat{\mathbf{e}}_2$) = 1.3076, is thus not only very close to the direct estimate, Var($\hat{\mathbf{e}}_3$) = 1.2591, but it is also well inside the computed 95% confidence interval. Such econometric evidence implies that the source of independent shocks in the RIP equation (9), e_{RIP}, is not important. Consequently, we conclude that our data do not reject the hypothesis that RIP results from the joint operation of UIP and PPP not just in a deterministic sense, but also in a statistical one, namely that $e_3 \approx e_1 - e_2$.

However, the analysis so far concerns a "long run" approximated by 23-year data averages (over our 1976-1998 sample). Can we say the same for shorter "long run" periods? Put otherwise, how fast is the speed of convergence, through the forces of arbitrage in asset and goods markets, for the three mutually dependent parity conditions bundled up here in the triple-parity law?

III.6 How Long is the "Long Run"? Speed of Convergence for the Triple-Parity Law

To address this question we computed the slope estimates and their 95% confidence interval in the same way as above, but now at successive cumulative sample horizons, starting with a short "long run" (or "medium term") of 5 years (1976-1980) and then iteratively moving on to progressively longer "long runs" until we reached the end of the sample (1976-1998). The results of these computations are presented in Figure 7.

[Figure 7 about here]

It is thus confirmed that in a "long run" of 23 years (1976-1998) all three conditions – UIP, PPP and

[43] See for instance Yamane (1973, 788-789).

RIP – hold convincingly: all three regressions result in slope estimates practically equal to 1 at the end of the sample horizon (LHS column of diagrams in Figure 7). The 95% confidence interval widths at the same point in time are of the order of 0.4-0.6 (RHS column in Figure 7).

This is however not the case, especially for the RIP equation, over most of the shorter cumulative horizons. Comparing the speed of convergence to the ultimate equilibrium points to some differences and, in particular, casts doubt on the (complete) dependence of the RIP on the joint action of UIP and PPP in shorter "long run" periods. More precisely, taking the 95% confidence band as indicative and looking at the three LHS diagrams, one can see that UIP holds even for periods of 5 years or so in our sample; PPP holds for longer "long runs" of about 9-10 years; it takes still longer for RIP to prevail in the data, some 14-15 years (the gray-shadowed areas in the figures visualize this failure of the RIP – and, incidentally, of PPP – in shorter "long runs").

IV. The Triple-Parity Law and the Closely Related Literature

The closest conceptual proximity to our triple-parity analysis we are aware of is in Marston (1995, 1997), Fujii and Chinn (2000), Obstfeld and Taylor (2000) and Juselius and MacDonald (2004). The latter two authors point out (p. 2) that the literature viewing RIP as jointly determined by UIP and PPP (and possibly some other equilibrium relationships, as in their study the term structure of interest rates) is only "nascent". They refer to just a few "exceptions", namely Johansen and Juselius (1992), Juselius (1995) and MacDonald and Marsh (1997, 1999). Marston (1995, 1997) however pays due credit to some earlier studies emphasizing, in essence, the interdependence we call the triple-parity law, such as Cumby and Obstfeld (1984), Mishkin (1984) and Frankel (1986). From the many papers where any joint determination, or systems approach, to the triple-parity law is neglected, the closest to our work, but only with respect to the slope coefficient estimates by individual equations, must be Lothian and Simaan (1998). In addition, none of the literature has, to our knowledge, ever measured and interpreted the intercepts of the equilibrium conditions comprising the law in the way we did, which was made possible by our more computationally-intensive strategy of shifting the reference country.

Adler-Lehmann (1983) have argued that RIP is ensured by financial arbitrage in bonds (at least indirectly, e.g. by trade flows). Yet Marston (1995, 152) is more skeptical, maintaining that, unlike for UIP, there is no sound theoretical rationale for RIP. He nevertheless was among the first to suggest, like other authors now seem to agree, that RIP will hold if the two underlying conditions, UIP and PPP, hold simultaneously. Marston (1995) makes this evident by decomposing the differential between the expected real interest rates in two countries, A and B, as

$$E_t(RR_{A,t}) - E_t(RR_{B,t}) = \{I_{A,t} - [I_{B,t} + E_t(D_{A/B,t+1})]\} - \{E_t(\Pi_{A,t+1}) - [E_t(\Pi_{B,t+1}) + E_t(D_{A/B,t+1})]\}, \qquad (17)$$

with RR denoting the real rate of return (or of interest) and E_t the (conditional) expectation operator. In his words:

> "The first term in parentheses is the uncovered (nominal) interest differential, while the second term is the expected deviation from PPP. So real interest parity must hold if (1) UIP holds and (2) ex ante PPP holds. Real interest parity involves "financial arbitrage" but of nominal, not real, returns.[44] The second condition involving ex ante PPP, however, is a condition involving goods markets, not financial markets." Marston (1995, 153-154).

Equation (1) in Obstfeld and Taylor (2000, 7) is equivalent to (17) above, but it is expressed in terms of the PPP-defined real exchange rate on its RHS. The derivation in equations (1)-(4) in Fujii and Chinn (2000, 4-5) essentially uncovers the same link: analytically, RIP thus appears to be the outcome of its key components, UIP and PPP.

On the empirical side, Marston (1995, 1997) reports an overall failure of ex ante tests of the three basic relationships in the earlier literature, where each was considered separately, as well as in his own

[44] "Strictly speaking, the term 'arbitrage' should be confined to riskless operations rather than to the risky positions required to ensure that UIP holds."; footnote as in original.

research, where their interdependence is stressed. But, in a way similar to ours, he also performs ex post tests of UIP, relative PPP and hence RIP, using the same monthly and quarterly time series for the G-5 countries[45] over the June 1973–December 1992 period as in his ex ante tests. The ex post tests confirm that all three equilibrium conditions ultimately hold, on average and for his G-5 sample.[46] This author did not however characterize his findings in an encompassing and coherent manner, neither estimated them as a system, i.e. as the manifestation of a general economic law in our sense.[47]

In assessing RIP, Fujii and Chinn (2000) use GMM estimation based on rational (unbiased) forecast errors and, alternatively, on univariate modeling of inflation forecasts. They find that RIP holds at longer rather than at short horizons, the hypothesis being decisively rejected in the latter case. Their results show that E(b)=1 in the RIP relationship is not rejected for most economies in their sample, G-7 with the USA as a reference country, when using 5- or 10-year interest rates and either consumer or wholesale price indexes.[48]

Employing unit root tests, filtering techniques and nonlinear threshold autoregressive (TAR) models within very long (more than a century) historical series at high (monthly) frequency for eight countries,[49] Obstfeld and Taylor (2000) also report that the long-run RIP condition did hold ex post,

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[45] France, Germany, Japan and the UK, with the USA as the reference country.

[46] Marston (1995, 173-175) argues that the reason for the different results for ex ante vs. ex post tests lies in the fact that some variables in the current information set of investors (such as nominal interest differentials, share yields and past inflation differentials) are systematically related to RIR differentials and that there is a significant time variation in RIR differentials.

[47] Moreover, he apparently did not allow for the structural differentials in the PPP relationship and performed a number of joint Wald tests, which, we have argued, are irrelevant.

[48] The authors also apply a joint Wald test for the a=0 and b=1 hypotheses, which we view as inappropriate. Furthermore, they do not allow for the normalization issue, although it is mentioned briefly (p. 6, footnote 2).

[49] Canada, France, Germany, Italy, Japan, the Netherlands and the UK, with the USA again being the country of reference.

most strikingly in the 1890-1914 gold standard period and somewhat less so in the post-1974 float. They argue that previous studies[50] had found negative results, or at least had not been supportive, due in part to overdifferencing and filtering problems. Moreover, they explicitly allow for the normalization and direction-of-causality issue, as we did.

The approach in Juselius and MacDonald (2004) is similar to ours in that they analyze the three parity conditions jointly, as a system of equations. However, they add a fourth relationship into their system, namely the term structure of interest rates, distinguishing a short-term (3-month treasury bill) and a long-run (government bond) interest rate. To do so, however, they rely on the expectations hypothesis of the term structure, which is not without controversy in both the theory and the data. Juselius and MacDonald (2004) also differ from us in that they look at two countries only, the USA and Germany, employing a cointegrated VAR model based on monthly data spanning the period July 1975–January 1998. Their main results are quite supportive of ours. First of all, they strongly reject the stationarity hypothesis for the parity conditions when taken separately, but stationarity is "recovered" when allowing for their interdependence. Their proposed interpretation is (p. 28) "the lack of (or very, very slow) adjustment to a stationary PPP steady state and increasing long-term bond spreads as plausible consequence of the latter." We have made a similar point here about how long the "long run" is. Moreover, another conclusion of theirs is that the short-term interest rate – neglected in our cross-section because we view economic laws as rather long-run regularities – is not a main driving force, unlike the remaining three variables in their cointegrated VAR, which are present in our analysis too.

As far as RIP is concerned when taken *independently* from UIP and PPP, a number of studies using different techniques have appeared in recent years; they are largely consistent with our findings. We shall mention only the most relevant ones.

Ferreira and Leon-Ledesma (2005) provide further support for the RIP hypothesis. They present

[50] Such as, e.g., Meese-Rogoff (1988) and Frankel (1989).

evidence for a reversion towards a zero differential for developed countries and a positive one for emerging markets, with the adjustment being highly asymmetric and markedly different for these two classes of countries. In essence, this is akin to our country-specific (dis)advantage terms. They apply unit root tests to characterize the dynamic behavior of RIR differentials, using monthly data covering only March 1995 through May 2002 and for just two countries, again the USA and Germany.

Sekioua's (2005) empirical results are supportive too of reversion toward RIP, for the UK, France and Japan relative to the USA. The methodology applied is local-to-unity asymptotics and the data are monthly, for long-term government bond yields and CPIs, going back to the end of the 19th century. His conclusions are based on confidence intervals rather than on point estimates, as we did; but he does so by supplementing unit root tests with confidence intervals for the dominant root and the half-life of shocks to RIR differentials, themselves found to be strongly persistent. Sekioua also suggests (p. 3) a reason why RIP should be tested as a separate condition: "If the residuals are non stationary, then shocks to the real interest rate differential, which incidentally represents deviations from RIP, are permanent and the validity of the RIP hypothesis is rejected".

Lopez and Reyes (2005) set as their objective to relate empirics to theory by examining RIR stationarity (when allowing for structural changes) and the stationarity of consumption growth implied by the consumption(-based) capital asset pricing model in finance. They use IFS quarterly series from 1957 to 2002 for the G-7 industrialized nations and find support for the stationarity of both the RIR and the consumption growth rate.

Finally, Bjornland and Hungnes (2005) address the PPP puzzle – but via UIP and, hence, indirectly RIP – in the special case of Norway where oil, a primary commodity, constitutes the majority of exports. They claim to have removed the PPP puzzle by controlling for the interest rate differential in the real exchange rate relationship. In other words, once the interdependence between UIP, the central parity condition in the capital market, and PPP, the central parity condition in the goods market, is accounted for so that these conditions are considered jointly, the PPP puzzle is resolved. The reason for their use of system estimation, FIML reparametrized as a vector equilibrium correction model (VEqCM), is (p. 6) similar to that in MacDonald-Marsh (1997) and Juselius-MacDonald (2004): "The balance of payments constraint implies that any imbalances in the current account have to be financed through the capital account".

All above-mentioned studies use *time series* data at relatively high frequencies (monthly, quarterly) and test the RIP relationship (or PPP for Norway) in samples of 2, 4, 5, 7 or 8 large countries. Juselius-MacDonald (2004) take the cointegration route, Bjornland-Hungnes (2005) use recent equilibrium correction techniques, Ferreira-Leon-Ledesma (2005), Sekioua (2005) and Lopez-Reyes (2005) build on unit roots and extensions, Fujii-Chinn (2000), Obstfeld-Taylor (2000) and Marston (1995, 1997) apply more traditional approaches. Our empirical strategy, based on a *cross-section* of trend growth rates for 18 OECD economies over 23 years, differs in that we traded-off a smaller number of annual observations for as large a sample of countries as possible. I.e. we have tested all three relationships for and across 18 industrialized countries. Given the data limitations and to comply with theory, these countries were purposefully selected to be as homogeneous as possible.

V. Conclusions

The triple-parity law we propose and test in the present paper is relatively simple, integrating as it does three familiar parity conditions in international economics, UIP, PPP and RIP. It has been shown to be fully consistent with the available evidence for our cross-section of trend growth rate differentials in 18 OECD economies over the post-Bretton-Woods/pre-EMU floating rate period. More precisely, our data and econometric checks could not reject (falsify) the triple-parity law hypothesis: the law can, therefore, be considered to hold in the long run, on average and ex post.

The test period (1976-1998) may well turn out to have been a unique "window of opportunity": if the Euro system does not disintegrate in the future, and all the more so if the UK and the other EMU outsiders should join it, the industrialized world might long remain dominated by three major currency areas only, the dollar, the euro and the yen zones (plus possibly a yuan one). This would mean that data such as those used here would henceforth make up a much less richer sample.

Taken individually, some of our results are not really new. What may be original is the formulation and testing of the triple-parity law as a long-run unity of UIP, PPP and RIP. The estimated and broadly interpreted average (for the sample) country-specific financial/institutional premia in UIP, real/structural differentials in PPP and resulting overall comparative (dis)advantages in RIP would constitute another novelty. Moreover, our straightforward but complementary estimation methods, extensive sensitivity analysis and various empirical findings make up a coherent whole, or so we hope.

Because of its simplicity, centrality and empirical testability, the triple-parity law should receive more attention in open-economy macroeconomics, being a manifestation of the long-run unity of three thus far separately studied and econometrically controversial equilibrium conditions. It constitutes, in essence, a succinct synthesis of basic insights on how (imperfectly unlimited) arbitrage in goods and financial markets ultimately leads the world economy to certain regularities and, thus, predictability. Finally, it illustrates a more fundamental point: if we look beyond short-term fluctuations and vagaries, economic laws do exist in the long run, just as economists used to think in the days of Marshall, Fisher, Walras and Pareto.

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Table 1: 23-Year (1976-1998) Annual Averages for the Cross-Section of 18 OECD Economies Used in the Triple-Parity Law Tests, USA as Country of Reference

Country	Nominal Depreciation % p.a.	CPI Inflation % p.a.	GDP Deflator Inflation % p.a.	Nominal Long-Term Interest % p.a.	Real (CPI) Depreciation % p.a.	Real (GDP Deflator) Depreciation % p.a.	Real (CPI) Long-Term Interest % p.a.	Real (GDP Deflator) Long-Term Interest % p.a.	CPI Inflation Differential % points	GDP Deflator Inflation Differential % points	Nominal Long-Term Interest Differential % points	Real (CPI) Long-Term Interest Differential % points	Real (GDP Deflator) Long-Term Interest Differential % points
Australia	2.61	6.14	4.70	11.00	1.03	2.00	4.86	6.30	1.58	0.61	2.31	0.73	1.70
Austria	-2.14	3.32	3.35	7.62	-0.90	-1.41	4.30	4.27	-1.25	-0.74	-1.07	0.18	-0.33
Belgium	-0.62	3.74	3.74	9.09	0.20	-0.27	5.35	5.35	-0.82	-0.35	0.40	1.22	0.75
Canada	1.03	4.81	4.15	9.91	0.78	0.98	5.10	5.77	0.25	0.05	1.22	0.97	1.17
Switzerland	-2.59	2.30	3.09	4.49	-0.33	-1.59	2.19	1.40	-2.26	-1.00	-4.20	-1.94	-3.20
Denmark	-0.22	4.87	4.82	11.38	-0.52	-0.94	6.51	6.56	0.30	0.73	2.69	2.38	1.96
Finland	0.80	5.26	5.35	11.08	0.11	-0.45	5.83	5.74	0.69	1.25	2.39	1.70	1.14
France	0.40	5.08	5.07	9.60	-0.12	-0.58	4.52	4.53	0.52	0.98	0.91	0.39	-0.07
Germany	-2.06	2.73	2.97	7.07	-0.23	-0.94	4.33	4.09	-1.83	-1.12	-1.62	0.20	-0.51
Italy	2.83	8.20	8.74	12.75	-0.81	-1.82	4.55	4.01	3.64	4.64	4.06	0.42	-0.59
Japan	-4.73	2.12	1.68	5.59	-2.28	-2.31	3.47	3.92	-2.45	-2.42	-3.10	-0.65	-0.68
Netherlands	-1.82	2.61	2.33	7.82	0.13	-0.06	5.22	5.50	-1.96	-1.77	-0.87	1.09	0.90
Norway	1.17	5.70	4.72	9.70	0.02	0.54	4.00	4.98	1.14	0.63	1.01	-0.13	0.38
New Zealand	2.54	7.85	7.20	11.06	-0.75	-0.56	3.20	3.86	3.29	3.11	2.37	-0.93	-0.74
Spain	2.69	8.10	8.18	12.31	-0.85	-1.40	4.21	4.13	3.54	4.09	3.62	0.08	-0.47
Sweden	2.30	6.33	6.24	11.14	0.53	0.15	4.81	4.90	1.77	2.15	2.45	0.68	0.30
UK	0.82	6.00	5.97	10.39	-0.61	-1.05	4.39	4.42	1.44	1.88	1.70	0.26	-0.17
USA	0.00	4.56	4.09	8.69	0.00	0.00	4.13	4.60	0.00	0.00	0.00	0.00	0.00
Average	0.17	4.99	4.80	9.48	-0.26	-0.54	4.50	4.69	0.42	0.71	0.79	0.37	0.09

Source: Calculations of the authors on the basis of the raw data and their transformations, as explained in section III of the main text.

Table 2

First Parity: The Uncovered Nominal Interest Rate Condition

$$\underbrace{D_{i/USA}}_{\mathbf{Y}} = a_1 + b_1\underbrace{(I_i - I_{USA})}_{\mathbf{X}} + e_{1,i}$$

	Regressing **Y** on **X**[a]		Orthogonal Regression[b]	Regressing **X** on **Y**[a]	
	OLS	WLS[c]		OLS	WLS[c]
Slope ($\hat{b}_1$)	**0.90**	**0.82**	**0.98**	**1.07**	**1.25**
t-stat.	9.11	5.49	-	9.00	5.32
Prob. value	0.00	0.00	-	0.00	0.00
95% conf. int.	0.69–1.11	0.50–1.14	-	0.82–1.32	0.75–1.76
99% conf. int.	0.61–1.19	0.38–1.26	-	0.72–1.42	0.56–1.95
Constant ($\hat{a}_1$)	**-0.39**	**-0.14**	**-0.44**	**-0.50**	**-0.84**
t-stat.	-1.78	-0.42	-	-2.08	-2.81
Prob. value	0.09	0.68	-	0.05	0.01
95% conf. int.	-0.87–0.09	-0.85–0.57	-	-1.01–0.01	-1.47–-0.20
99% conf. int.	-1.05–0.27	-1.12–0.84	-	-1.20–0.21	-1.72–0.04
Adj. r^2 or ϕ^2 (ODR)	**0.83**	**0.75**	**0.92**	**0.83**	**0.80**
F prob. Value	0.00	0.00	-	0.00	0.00

a/ **All values are given for the Y = a + bX relationship**. The t-statistics and confidence intervals from the X-on-Y regressions were transformed so as to be comparable to the Y-on-X results by applying the delta method.

b/ Unweighted.

c/ WLS uses as weights the 1990 values of the various countries' GDP converted into a common currency via the 1990 PPP exchange rates as calculated by the OECD.

Table 3

Second Parity: The Relative PPP Condition

$$\underbrace{D_{i/USA}}_{\mathbf{Y}} = a_2 + b_2\underbrace{(\Pi_i - \Pi_{USA})}_{\mathbf{X}} + e_{2,i}$$

	Regressing **Y** on **X**[a]		Orthogonal Regression[a,b]	Regressing **X** on **Y**[a]	
	OLS	WLS[c]		OLS	WLS[c]
A. Taking the GDP Deflators					
Slope ($\hat{b}_2$)	**0.97**	**0.94**	**1.13**	**1.28**	**1.13**
t-stat.	7.12	9.07	-	6.98	8.96
Prob. Value	0.00	0.00	-	0.00	0.00
95% conf. int.	0.68–1.26	0.72–1.16	-	0.87–1.67	0.86–1.39
99% conf. int.	0.57–1.37	0.64–1.25	-	0.74–1.82	0.76–1.50
Constant ($\hat{a}_2$)	**-0.53**	**-0.41**	**-0.64**	**-0.75**	**-0.72**
t-stat.	-1.93	-1.73	-	-2.45	-3.92
Prob. value	0.07	0.10	-	0.02	0.00
95% conf. int.	-1.13–0.07	-0.91–0.09	-	-1.40–-0.15	-1.11–-0.33
99% conf. int.	-1.36–0.30	-1.10–0.29	-	-1.65–0.15	-1.26–0.18
Adj. r^2 or ϕ^2 (ODR)	**0.74**	**0.88**	**0.87**	**0.74**	**0.91**
F prob. value	0.00	0.00	-	0.00	0.00
B. Taking the CPIs					
Slope ($\hat{b}_2$)	**1.07**	**0.91**	**1.15**	**1.22**	**1.00**
t-stat.	10.52	13.39	-	10.43	13.32
Prob. value	0.00	0.00	-	0.00	0.00
95% conf. int.	0.85–1.29	0.77–1.06	-	0.97–1.47	0.84–1.15
99% conf. int.	0.77–1.37	0.71–1.11	-	0.88–1.57	0.78–1.22
Constant ($\hat{a}_2$)	**-0.33**	**-0.26**	**-0.37**	**-0.40**	**-0.39**
t-stat.	-1.71	-1.67	-	-2.00	-2.87
Prob. value	0.11	0.11	-	0.06	0.01
95% conf. int.	-0.75–0.09	-0.59–0.07	-	-0.83–0.03	-0.68–-0.10
99% conf. int.	-0.91–0.25	-0.72–0.20	-	-0.99–0.19	-0.79–0.01
Adj. r^2 or ϕ^2 (ODR)	**0.87**	**0.94**	**0.94**	**0.87**	**0.95**
F prob. value	0.00	0.00	-	0.00	0.00

a/ **All values are given for the Y = a + bX relationship**. The t-statistics and confidence intervals from the X-on-Y regressions were transformed so as to be comparable to the Y-on-X results by applying the delta method.

b/ Unweighted.

c/ WLS uses as weights the 1990 values of the various countries' GDP converted into a common currency via the 1990 PPP exchange rates as calculated by the OECD.

Table 4

Third Parity: The Real Interest Rate Condition

$$\underbrace{(I_i - I_{USA})}_{\mathbf{Y}} = a_3 + b_3\underbrace{(\Pi_i - \Pi_{USA})}_{\mathbf{X}} + e_{3,i}$$

	Regressing **Y** on **X**[a]		Orthogonal Regression[a,b]	Regressing **X** on **Y**[a]	
	OLS	WLS[c]		OLS	WLS[c]
A. Taking the GDP Deflators					
Slope ($\hat{b}_3$)Error! Bookmark not defined.	**0.97**	**0.76**	**1.16**	**1.33**	**1.37**
t-stat.	6.47	4.46	-	6.32	4.25
Prob. value	0.00	0.00	-	0.00	0.00
95% conf. int.	0.65–1.28	0.40–1.12	-	0.88–1.78	0.68–2.05
99% conf. int.	0.53–1.40	0.26–1.26	-	0.71–1.96	0.42–2.31
Constant ($\hat{a}_3$)	**-0.07**	**0.33**	**-0.21**	**-0.34**	**-0.69**
t-stat.	-0.24	0.87	-	-0.98	-1.56
Prob. value	0.81	0.40	-	0.34	0.14
95% conf. int.	-0.73–0.59	-0.49–1.15	-	-1.06–0.39	-1.64–0.25
99% conf. int.	-0.98–0.84	-0.80–1.47	-	-1.34–0.67	-2.14–0.75
Adj. r^2 or ϕ^2 (ODR)	**0.71**	**0.74**	**0.85**	**0.71**	**0.76**
F prob. value	0.00	0.00	-	0.00	0.00
B. Taking the CPIs					
Slope ($\hat{b}_3$)	**1.03**	**0.70**	**1.19**	**1.31**	**1.27**
t-stat.	7.70	4.44	-	7.58	4.23
Prob. value	0.00	0.00	-	0.00	0.00
95% conf. int.	0.75–1.32	0.36–1.04	-	0.94–1.68	0.63–1.90
99% conf. int.	0.64–1.43	0.24–1.16	-	0.80–1.82	0.39–2.15
Constant ($\hat{a}_3$)	**0.14**	**0.51**	**0.07**	**0.01**	**-0.38**
t-stat.	0.56	1.42	-	0.04	-0.81
Prob. value	0.58	0.18	-	0.97	0.43
95% conf. int.	-0.41–0.70	-0.26–1.27	-	-0.60–0.63	-1.39–0.62
99% conf. int.	-0.63–0.91	-0.55–1.56	-	-0.84–0.86	-1.78–1.01
Adj. r^2 or ϕ^2 (ODR)	**0.77**	**0.74**	**0.89**	**0.77**	**0.73**
F prob. value	0.00	0.00	-	0.00	0.00

a/ **All values are given for the Y = a + bX relationship**. The t-statistics and confidence intervals from the X-on-Y regressions were transformed so as to be comparable to the Y-on-X results by applying the delta method.

b/ Unweighted.

c/ WLS uses as weights the 1990 values of the various countries' GDP converted into a common currency via the 1990 PPP exchange rates as calculated by the OECD.

Table 5

Nominal UIP: The "F"-Differentials or Financial/Institutional Disadvantages

Reference Country[a]	Estimated Constants or Own-Country Residuals, in Percentage Points[b]							
	Y-on-X Regression[c]				X-on-Y Regression[c]			
	OLS		WLS		OLS		WLS	
	$\hat{a}_1$	p-value	**$\hat{a}_1$**	p-value	**$\hat{a}_1$**	p-value	**$\hat{a}_1$**	p-value
Switzerland	**-1.60**	0.01	**-1.02**	0.28	**-2.44**	0.00	**-3.55**	0.00
Sweden	**-1.09**	0.00	**-0.98**	0.00	**-0.89**	0.04	**-0.90**	0.05
Australia	**-0.95**	0.00	**-0.88**	0.00	**-0.66**	0.12	**-0.59**	0.21
New Zealand	**-0.72**	0.02	**-0.66**	0.02	**-0.39**	0.31	**-0.28**	0.50
Canada	**-0.40**	0.08	**-0.24**	0.33	**-0.31**	0.25	**-0.45**	0.15
USA	**-0.39**	0.09	**-0.14**	0.68	**-0.50**	0.05	**-0.84**	0.01
Norway	**-0.37**	0.11	**-0.23**	0.35	**-0.25**	0.36	**-0.35**	0.25
UK	**-0.19**	0.40	**-0.03**	0.91	**-0.10**	0.69	**-0.23**	0.45
France	**-0.01**	0.98	**0.18**	0.50	**0.04**	0.87	**-0.14**	0.66
Spain	**0.01**	0.97	**-0.01**	0.98	**0.50**	0.13	**0.79**	0.00
Germany	**0.07**	0.83	**0.46**	0.42	**-0.35**	0.27	**-1.01**	0.07
Italy	**0.09**	0.81	**0.05**	0.90	**0.62**	0.07	**0.95**	0.00
Netherlands	**0.40**	0.16	**0.74**	0.13	**0.09**	0.78	**-0.46**	0.43
Belgium	**0.50**	0.03	**0.73**	0.03	**0.44**	0.12	**0.15**	0.73
Austria	**0.62**	0.04	**0.97**	0.06	**0.30**	0.43	**-0.27**	0.68
Finland	**0.63**	0.02	**0.72**	0.01	**0.87**	0.00	**0.91**	0.00
Japan	**1.47**	0.00	**1.97**	0.02	**0.80**	0.24	**-0.12**	0.92
Denmark	**1.93**	0.00	**2.00**	0.00	**2.23**	0.00	**2.32**	0.00

a/ In ascending order for the OLS direct regression.

b/ Shaded values are significant at the 5% level.

c/ **All values are given for the Y = a + bX relationship**.

Table 6

Relative PPP: The "R"-Differentials or Real/Structural Advantages

Reference Country[a]	Estimated Constants or Own-Country Residuals, in Percentage Points[b]							
	Y-on-X Regression[c]				X-on-Y Regression[c]			
	OLS		WLS		OLS		WLS	
	$\hat{a}_2$	p-value	**$\hat{a}_2$**	p-value	**$\hat{a}_2$**	p-value	**$\hat{a}_2$**	p-value
Australia	**-2.55**	0.00	**-2.45**	0.00	**-2.58**	0.00	**-2.64**	0.00
Canada	**-1.51**	0.00	**-1.39**	0.00	**-1.71**	0.00	**-1.69**	0.00
Norway	**-1.08**	0.00	**-0.98**	0.00	**-1.11**	0.02	**-1.17**	0.00
Sweden	**-0.75**	0.03	**-0.69**	0.00	**-0.31**	0.51	**-0.60**	0.03
USA	**-0.53**	0.07	**-0.41**	0.10	**0.75**	0.02	**-0.72**	0.00
Netherlands	**-0.42**	0.34	**-0.25**	0.53	**-1.18**	0.00	**-0.88**	0.01
Belgium	**-0.24**	0.42	**-0.12**	0.67	**-0.57**	0.07	**-0.49**	0.05
Finland	**-0.11**	0.68	**-0.03**	0.87	**0.05**	0.88	**-0.11**	0.56
France	**0.02**	0.93	**0.12**	0.51	**0.10**	0.73	**-0.01**	0.95
New Zealand	**0.12**	0.79	**0.15**	0.53	**0.91**	0.02	**0.44**	0.02
Denmark	**0.39**	0.15	**0.49**	0.02	**0.40**	0.22	**0.32**	0.20
Germany	**0.44**	0.23	**0.59**	0.09	**-0.12**	0.78	**0.08**	0.84
UK	**0.46**	0.14	**0.53**	0.00	**0.82**	0.01	**0.57**	0.00
Spain	**0.75**	0.18	**0.75**	0.02	**1.78**	0.00	**1.20**	0.00
Austria	**0.90**	0.01	**1.04**	0.00	**0.45**	0.36	**0.60**	0.19
Switzerland	**1.09**	0.01	**1.14**	0.00	**0.56**	0.32	**0.74**	0.15
Italy	**1.15**	0.07	**1.23**	0.00	**2.35**	0.00	**1.69**	0.00
Japan	**1.86**	0.00	**2.04**	0.00	**0.90**	0.31	**1.29**	0.11

a/ In ascending order for the OLS direct regression.

b/ Shaded values are significant at the 5% level.

c/ **All values are given for the Y = a + bX relationship**.

Table 7

RIP: The Combined "F" and "R" Differentials, Overall Disadvantages or RIR Premia

Reference Country[a]	Estimated Constants or Own-Country Residuals, in Percentage Points[b]							
	Y-on-X Regression[c]				X-on-Y Regression[c]			
	OLS		WLS		OLS		WLS	
	$\hat{a}_3$	p-value	**$\hat{a}_3$**	p-value	**$\hat{a}_3$**	p-value	**$\hat{a}_3$**	p-value
Switzerland	**-3.19**	0.00	**-3.80**	0.00	**-2.56**	0.05	**-2.17**	0.18
Japan	**-0.80**	0.16	**-1.70**	0.04	**0.36**	0.58	**0.79**	0.45
Italy	**-0.73**	0.28	**-0.17**	0.77	**-2.18**	0.00	**-1.97**	0.00
Germany	**-0.63**	0.13	**-1.27**	0.03	**0.05**	0.93	**0.44**	0.58
New Zealand	**-0.63**	0.20	**-0.35**	0.36	**-1.58**	0.00	**-1.33**	0.00
UK	**-0.59**	0.10	**-0.61**	0.03	**-1.02**	0.01	**-0.72**	0.04
Austria	**-0.48**	0.20	**-1.03**	0.05	**0.06**	0.88	**0.44**	0.52
Spain	**-0.43**	0.47	**0.02**	0.97	**-1.67**	0.00	**-1.45**	0.02
Sweden	**-0.22**	0.54	**-0.18**	0.52	**-0.75**	0.04	**0.46**	0.19
France	**0.00**	0.99	**-0.20**	0.48	**-0.10**	0.77	**0.23**	0.56
USA	**0.07**	0.81	**-0.33**	0.40	**0.34**	0.33	**0.69**	0.14
Netherlands	**0.62**	0.20	**-0.15**	0.82	**1.54**	0.00	**1.95**	0.00
Belgium	**0.70**	0.04	**0.23**	0.61	**1.10**	0.01	**1.46**	0.00
Norway	**0.78**	0.01	**0.51**	0.13	**0.81**	0.06	**1.15**	0.01
Finland	**0.89**	0.01	**0.75**	0.01	**0.69**	0.15	**1.01**	0.03
Canada	**1.15**	0.00	**0.76**	0.06	**1.40**	0.01	**1.75**	0.00
Denmark	**1.72**	0.00	**1.47**	0.00	**1.72**	0.01	**2.05**	0.00
Australia	**1.77**	0.00	**1.49**	0.00	**1.81**	0.01	**2.15**	0.00

a/ In ascending order for the OLS direct regression.

b/ Shaded values are significant at the 5% level.

c/ **All values are given for the Y = a + bX relationship**.

Table 8

Triple Parity: SUR/FIML Pairwise Results[a]

Estimated equations taken two by two[b]	**Slope ($\hat{b}$)**Error! Bookmark not defined.	**Constant ($\hat{a}$)**	**Adj. r^2**
(10) UIP	**1.00** (10.3; 3.0) (0.098; 0.332)	**–0.46** (–2.1; –1.1) (0.21; 0.43)	0.82
(11) PPP	**1.20** (8.0; 4.6) (0.149; 0.261)	**–0.69** (–2.4; –1.4) (0.282; 0.490)	0.74
(10) UIP	**1.00** (10.9; 3.0) (0.092; 0.332)	**–0.46** (–2.1; –1.1) (0.213; 0.434)	0.82
(12) RIP	**1.19** (7.7; 3.6) (0.155;0.331)	**–0.23** (–0.8; –0.5) (0.304; 0.445	0.70
(11) PPP	**1.20** (9.9; 4.6) (0.121; 0.261)	**–0.69** (–2.5; –1.4) (0.276; 0.490)	0.70
(12) RIP	**1.19** (8.9; 3.6) (0.134; 0.331)	**–0.23** (–0.8; –0.5) (0.299; 0.445)	0.66

a/ t-statistics (second line) and standard errors (third line) in parentheses below coefficients; first figure: SUR; second: FIML.

b/ All results are given for the Y-on-X specification. Inflation differentials from GDP deflators. Country of reference: USA. SUR: simultaneous weighting matrix and coefficient iteration.

Table 9

Triple Parity: Full System SUR vs OLS Results[a]

Equation	SUR Results			OLS Results[b]		
	Slope ($\hat{b}$)	**Constant ($\hat{a}$)**	**Adj. r^2**	**Slope ($\hat{b}$)**	**Constant ($\hat{a}$)**	**Adj. r^2**
(10) UIP	**0.97** (11.2) (0.086)	**–0.43** (–2.1) (0.206)	0.82	**0.90** (9.1) (0.099)	**–0.39** (–1.8) (0.219)	0.83
(11) PPP[c]	**1.14** (9.5) (0.121)	**–0.65** (–2.5) (0.263)	0.73	**1.28** (7.1) (0.180)	**–0.75** (–2.3) (0.321)	0.75
(12) RIP	**1.13** (9.4) (0.119)	**–0.19** (–0.7) (0.278)	0.69	**0.97** (6.5) (0.149)	**–0.07** (–0.2) (0.300)	0.71

a/ t-statistics (second line) and standard errors (third line) in parentheses below coefficients. Inflation differentials from GDP deflators. Country of reference: USA. SUR: one-step weighting matrix and coefficient iteration.

b/ See tables 2 and 4 (direct OLS regressions for UIP and RIP) and Table 3 (reverse OLS regression for PPP).

c/ Renormalized – see text. Equation estimated as: ΠUSA = – (c(3)/c(4)) + (1/c(4))*DUSA.

Figure 1

Uncovered Nominal Interest Parity

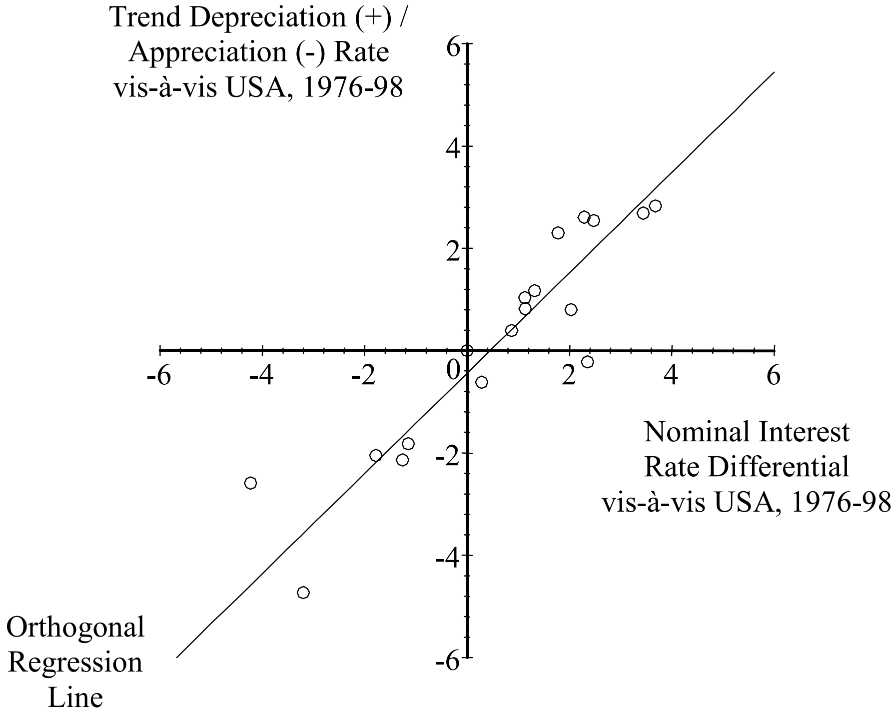


Figure 2

GDP Deflator-Based Purchasing Power Parity

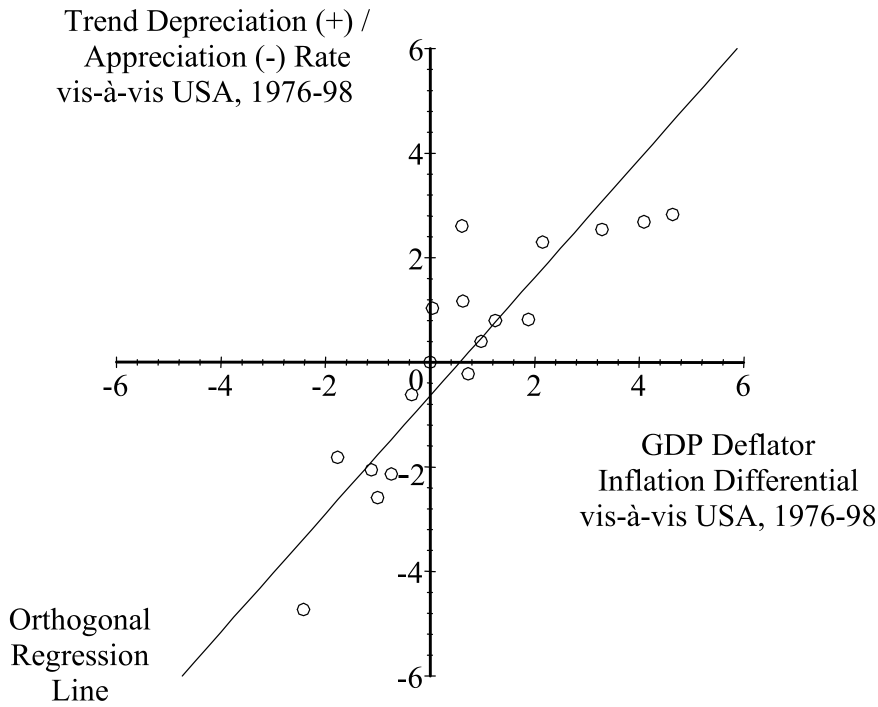


Figure 3

CPI-Based Purchasing Power Parity

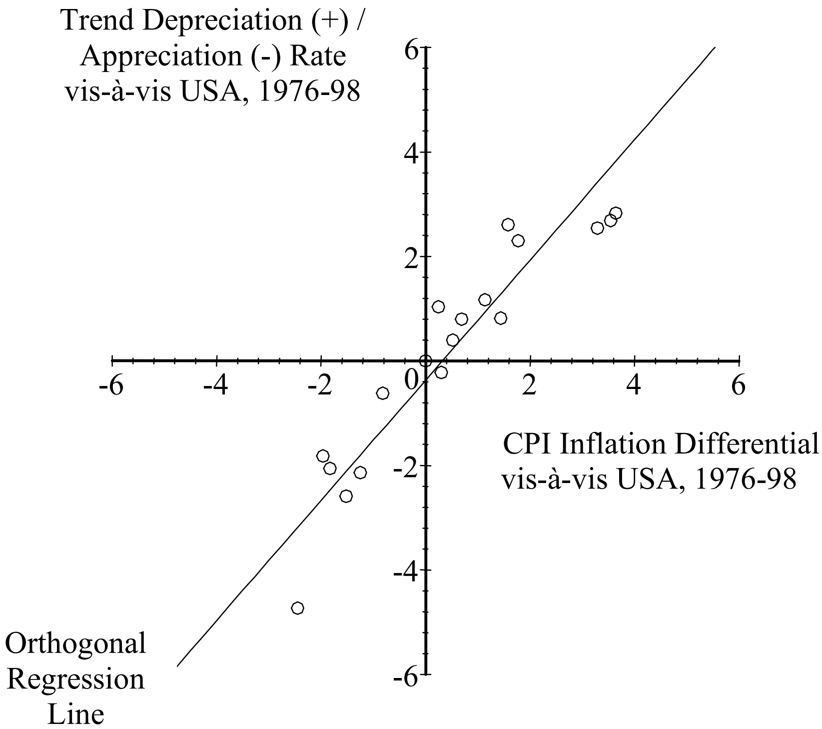


Figure 4

GDP Deflator-Based Real Interest Parity

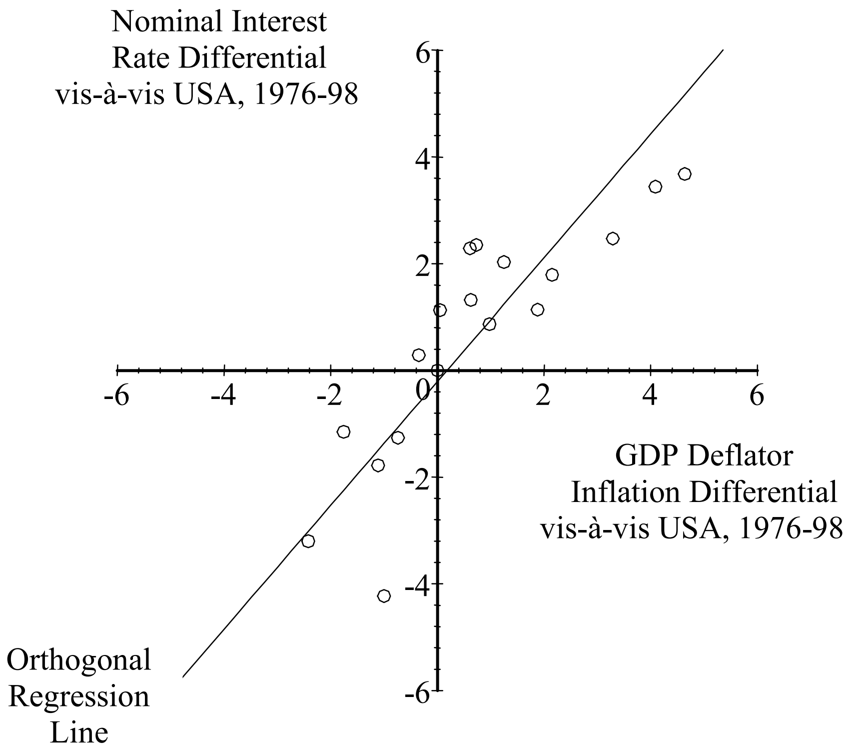


Figure 5

CPI-Based Real Interest Parity

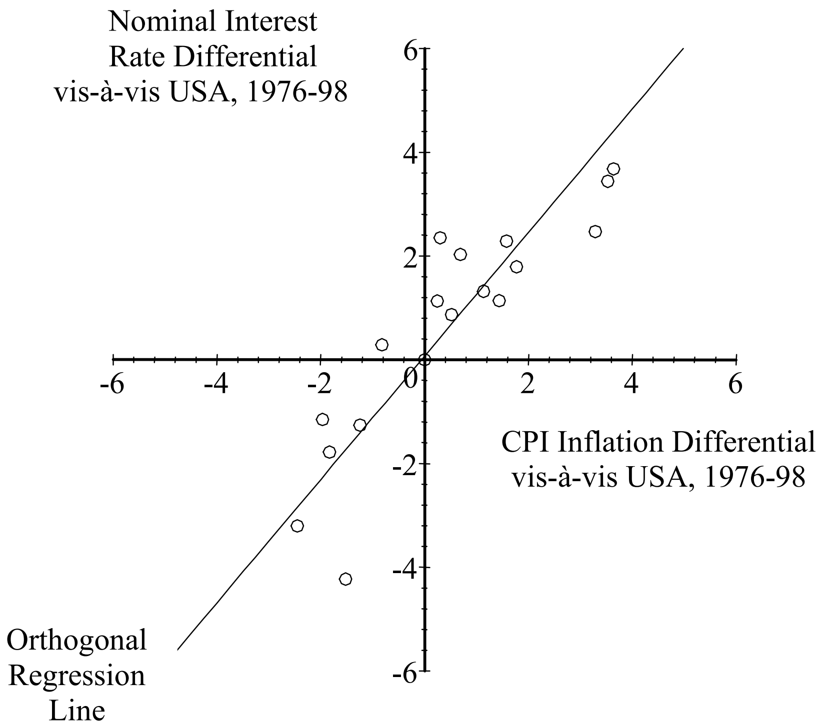


Figure 6

1976-1998 Average Real Long-Term Interest Rates Using GDP Deflators

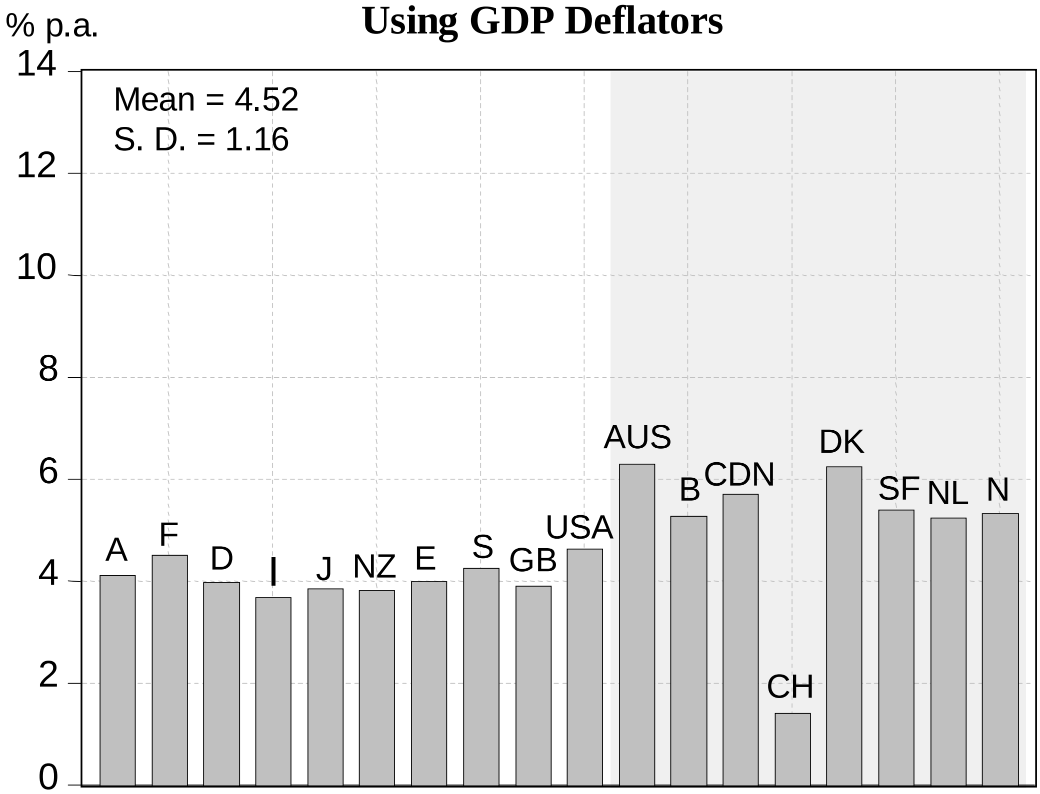
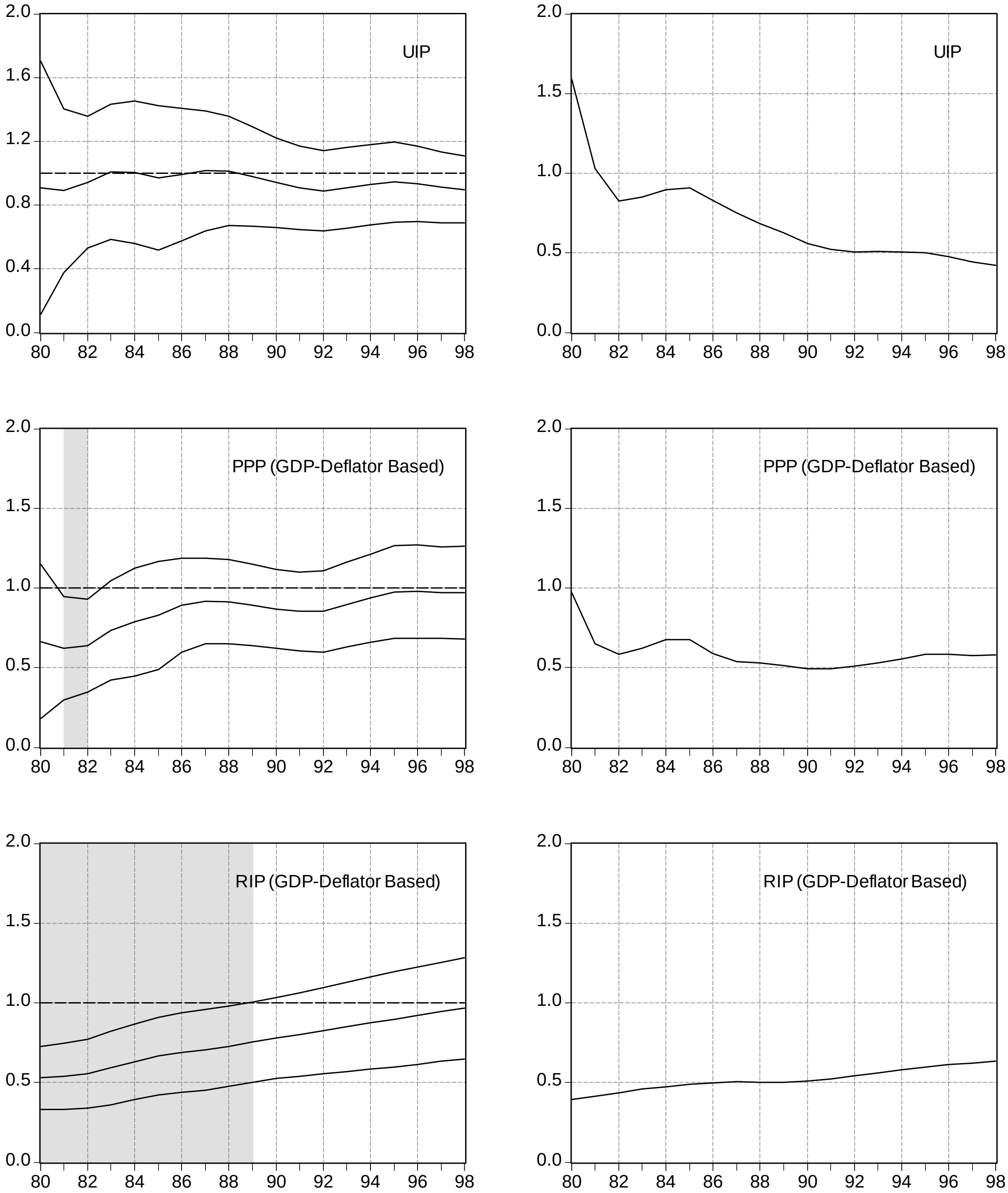


Figure 7

"Long-Run" Cumulative 95% Confidence Intervals for the Triple-Parity Law



LHS column: point estimates and 95% confidence interval band; RHS column: width of the 95% confidence interval band. Gray zones indicate "failure" of the respective equilibrium condition to prevail in the data, as measured by the 95% confidence interval.