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THE TRIPLE-PARITY LAW

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Abstract

Scientists and epistemologists generally agree that a scientific law must be (a) relatively simple, and (b) fully corroborated by the available evidence. In this paper, we propose one such law pertaining to international economics. It integrates three well-known equilibrium conditions which are shown to be fully verified in the long run, on average and ex post: (i) the uncovered, but risk-inclusive nominal interest rate parity condition; (ii) the purchasing power parity condition, in its relative form and allowing for inter-country productivity growth differentials; (iii) the real interest rate parity condition, allowing for both country-specific risks and inter-country productivity growth differentials.

Taken individually, none of our empirical results is really new, with the possible exception of the various countries' estimated "R" (risk-premium) and "T" (technology-related productivity-enhancing growth) factors; what may also be novel is that the analyses and the empirical results make up an integrated whole. The test period (1976-1998) will perhaps turn out to have been unique: if the euro system does not disintegrate in the future, and all the more so if Britain and the other EU outsiders or candidates should join it, the industrialized world might long remain divided into, and dominated by, three major currency areas only: the dollar, the euro and the yen zones. This would mean that data such as those used here will henceforth make up a much less richer sample.

The law holds in the long run, on average and ex post, and consequently it says nothing about short-term dynamics although it can provide useful benchmarks in this context too. Yet, it confirms that "reversal to the mean" must eventually occur and it may also be useful as a long-term equilibrium condition in (e.g.) error-correction models. Furthermore, it highlights the divergence in some important exchange rate reactions depending on the time horizon adopted: in the short-term, if interest rates increase in a given country, one normally expects its currency to appreciate; according to the triple-parity law, the opposite obtains in the long run. Finally, it confirms a basic choice for open economies under floating exchange rates: in the longer run, a given country can follow a strong currency, low inflation and low nominal interest rate policy; or it can pursue a weak currency, high inflation and high nominal interest rate policy; or anything else along this axis. But it cannot, for example, have both permanently low interest rates to stimulate investment and a permanently weak currency to foster exports. Country-specific factors, i.e. comparative advantages or disadvantages such as political stability, an especially efficient financial sector, a banking secrecy law or being a tax haven, can however mitigate this choice to some extent, at least for a few small countries. With the same proviso, real interest rates tend to be equalized the world over. The triple-parity law may also illustrate a more general 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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> 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[2]

Scientists and epistemologists generally agree that, to be worthy of the name, a scientific law must be (a) relatively simple, and (b) fully corroborated by the available evidence.[3] Here is one such law pertaining to international economics. It integrates three well-known equilibrium conditions which are shown to be fully verified in the long run, on average and ex post: the uncovered, but risk-inclusive nominal interest rate parity condition; the purchasing power parity condition, in its relative form and allowing for inter-country productivity growth differentials; the real interest rate parity condition, allowing for both country-specific risks and inter-country productivity growth differentials.

The underlying basic theory is summarized in a first section with only a few references to the literature; the same goes for the empirical results in section 2. How the theory and the econometric findings relate to the existing literature is discussed in more detail in section 3.[4] Our conclusions are set forth in section 4.

1. Theoretical Framework

Consider a two-country world (A and B) where there exists "sufficient" – i.e. not necessarily perfect – mobility of capital, goods and services across the border.[5] Because of capital mobility, the following (first) arbitrage condition must hold – and will be shown to hold – in the long run, on average and ex post:

(1) $D_{A/B} = (R_A - R_B) + (I_A - I_B) + \varepsilon_1$

All variables are expressed in percent p.a. $D_{A/B}$ is the **d**epreciation rate of currency A with respect to currency B as measured by the spot exchange rate[6]. R_A and R_B are country-specific **r**isk premia, about which more in the next paragraph. I_A and I_B are long-term **i**nterest rates; their exact definition and measurement will be discussed in section 2. Equation (1) is the classical uncovered but risk-inclusive nominal interest rate parity condition in its ex post formulation. It holds whether the currencies are floating or not; a fixed exchange rate system simply means $D_{A/B} = 0$ (~0, because of gold points or allowable fluctuation margins).

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

[3] This article builds on a previous paper by the first author (Lambelet 1996).

[4] Putting most references to – and comparisons with – the existing literature into a special section will make the article easier to read, e.g. for graduate students starting in international economics.

[5] The lesser the mobility, the longer it will generally take for the three parity conditions to be realized.

[6] Being the price of one unit of B's currency in terms of A's currency. Consequently, $D_{A/B} > 0$ means that the currency of country A is **depreciating** with respect to that of country B.

"Country-specific risk premium" must be understood here in the broadest sense as it may include other factors besides the habitually mentioned risks (i.e. the default risk, the political or sovereign 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*, R_A will then be less than R_B. As these examples show, the risk premium could also be a safety premium, so that "comparative disadvantage/advantage" is a more general and hence better description.[7] Section 2 will provide estimates of the average "R" factor differential for each of 18 industrialized countries over the 1976-1998 period.

If, like capital, goods and services are sufficiently mobile, arbitrage also ensures that the following (second) condition will be fulfilled in the long run, on average and ex post:

$$(2)\quad D_{A/B} = (T_B - T_A) + (CP_A - CP_B) + \varepsilon_2$$

which is one version of the familiar purchasing power parity (PPP) condition, in its relative form[8] and with explicit allowance for inter-country productivity growth differentials. CP_A and CP_B are national inflation rates (CP = **c**hange in **p**rices, in percent p.a.). T_B and T_A are **t**echnology-related productivity-enhancing growth factors, also expressed in percent p.a., to be estimated in section 2, and henceforth called the "T" factors. [9]

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

$$(3)\quad (R_A - R_B) + (I_A - I_B) = (T_B - T_A) + (CP_A - CP_B) + (\varepsilon_2 - \varepsilon_1), \text{ or equivalently}$$

$$(4)\quad (I_A - CP_A) = [(T_B - T_A) - (R_A - R_B)] + (I_B - CP_B) + (\varepsilon_2 - \varepsilon_1),$$

which is one version of the – perhaps less well-known – real interest rate parity (and third) condition, with explicit allowance for both risk and productivity differentials. It is not always realized that if (1) and (2) hold, (3)-(4) must too. In other words, the uncovered nominal interest rate parity condition and the relative PPP condition mean, when taken together, that real interest rates must also be equalized internationally in the long run, on average and ex post. Whether (3)-(4) might nevertheless be considered as a separate condition "in its own right" will be discussed later.

Combining (1), (2) and (3), and ignoring the error terms, we get:

$$\boxed{\mathbf{(5)}\ \underbrace{D_{A/B} - (R_A-R_B) - (I_A-I_B)}_{\substack{\text{Risk-inclusive nominal}\\ \text{interest rate parity}}} = \underbrace{D_{A/B} - (T_B-T_A) - (CP_A-CP_B)}_{\substack{\text{Relative PPP allowing for}\\ \text{productivity differentials}}} = \underbrace{(R_A-R_B) + (I_A-I_B) - (T_B-T_A) - (CP_A-CP_B)}_{\substack{\text{Real interest rate parity allowing for}\\ \text{risk and productivity differentials}}} = 0}$$

which is the *triple-parity law*, to be tested below.

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[7] This broader perspective is often ignored – thus Fujii and Chinn (2000, 4): "The existence of a covered interest differential is often taken as a manifestation of 'political risk', caused either by capital controls, or the threat of their imposition. In the absence of such barriers, such differentials should not exist because they imply unlimited arbitrage profit opportunities."

[8] Because some goods (the "non-tradables") cannot be exchanged internationally due to their physical nature (e.g. housing) or for other reasons (e.g. prohibitive tariffs or transportation costs), this parity condition is different from the generalized "law of one price", i.e. the absolute purchasing power parity condition. For the resulting "Scandinavian model" (already found in Ricardo), see e.g. Dornbusch (1987) and Artus (1989, 260-7). Lambelet (1996, 4-5) shows how this model leads directly to relative PPP with productivity differentials.

[9] On these productivity growth factors in the relative PPP condition, see Dornbusch (1987). Notice that the T factors are reversed in (2): T_B comes before T_A. This is because they are defined positively while the R factors in (1) are defined negatively; i.e. a large R is a "bad", but a large T is a "good".

Note that this law is entirely specified in terms of rates of change over time[10] and is thus compatible with any number of different combinations of interest and inflation differentials. This is illustrated in table 1 if, leaving the risk and productivity differentials aside for simplicity's sake, we rewrite (5) in the following manner:[11]

(6) $D_{A/B} = I_A - I_B = CP_A - CP_B$,

Table 1 (% p.a.)

	$D_{A/B}$	=	I_A	–	I_B	=	CP_A	–	CP_B	Real interest rate in both countries
(a)	4%	=	8%	–	4%	=	5%	–	1%	3%
	(4)	=		(4)		=		(4)		
(b)	0	=	5	–	5	=	2	–	2	3
	(0)	=		(0)		=		(0)		
(c)	4	=	12	–	8	=	9	–	5	3
	(4)	=		(4)		=		(4)		
(d)	4	=	8	–	4	=	4	–	0	4
	(4)	=		(4)		=		(4)		

In example (a), B is a strong currency, low nominal interest rate, low inflation country, and conversely for A. In example (b) the two countries are identical on all three counts, as will be the case under a fixed exchange rate system.[12] Also note that according to the triple-parity law the depreciation/appreciation rate must be equal to the inflation and interest rate *differentials*, but it says nothing about the particular values of the inflation and interest rates which make up any given differential – see (c) as compared to (a). In examples (a) through (c) the real interest rate is 3% p.a. in both economies. There is nothing preordained about that particular value: the real interest rate might just as well be 4%, as shown by example (d) compared to (a) and (c). All that the triple-parity law requires is that the real interest be the same in both countries – if, as in (6), we ignore the "R" and "T" factors.

2. Empirical Verification

The following data, all from the IMF's *International Financial Statistics* (IFS) or from the OECD's national accounting publications, were collected for each of 18 industrialized countries[13] over the 1976-1998 period[14]: the average annual values of the nominal spot

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

[11] This is how this law was specified in the first author's 1996 paper, where it was dubbed "the double equality".

[12] Always assuming that capital as well as goods and services are sufficiently mobile, which was hardly the case for most countries during most of the Bretton-Woods era (1944-1973) – with however a few exceptions such as the USA and Switzerland. In the case of those two countries, the first author found in some earlier work (Lambelet 1986) that the respective interest rates on comparable long-term U.S. and Swiss government bonds did tend to be equalized in the Bretton-Woods period, with however a permanent risk differential of about half a percentage point in favor of Switzerland, probably due to that country's banking secrecy law, its international aloofness in tax matters, its political and economic stability, and perhaps also its efficient financial sector.

[13] 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 could be presumed to exist over most of the sample period: Australia, Austria, Belgium, Canada, Denmark, Finland, France, Germany, Italy, Japan, Netherlands, New Zealand, Norway, Spain, Sweden, Switzerland, the UK, and the USA.

[14] 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 (*note continued on next page*)

exchange rate vis-à-vis the US dollar; the average annual levels of both the GDP deflator and the CPI; the average annual interest rate on long-term government bonds. In the latter case, our objective was to select homogeneous bonds with a long and uniform maturity (say, 10 years), but this proved unfeasible.[15] Accordingly, this series could easily be the one most likely to suffer from a serious 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, i.e. those in the IFS, therefore reflect various country-specific definitions. Broadly speaking, there are three groups of countries. First, some such as Austria, Germany and Japan report the average yield on all government bonds.[16] Second, most 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;[17] the same rule applies, but for maturities in excess of 10 years, to Canada, Belgium[18] and Italy;[19] 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[20] and New Zealand, 10 years for Finland,[21] the Netherlands and the USA, and 20 years for the UK.[22]

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,[23] and each national series' mean value was taken.

In a first step, average depreciation/appreciation, inflation and interest rate differentials were calculated with respect to the USA. In a second step, these differentials were then computed taking in turn each of the other economies as the reference country, so as to obtain estimates of the various relative risk premia and productivity differentials – see below.

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the emergence 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 (see the concluding section).

[15] Fujii and Chinn (2000) were able to use two apparently more homogeneous series for the G7 countries. The first is the yields on outstanding government bonds with a 10-year maturity used by Edison and Pauls (1993). The second consists of the synthetic "constant maturity" 5- and 10-year yields interpolated from the yield curve of outstanding government securities, as "obtained from the IMF country desks". We'll try to secure these series for the final version of our paper, although it is very doubtful that they exist for all 18 countries in our sample.

[16] 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.

[17] 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.

[18] 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.

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

[20] Yield to maturity.

[21] Since the Finnish 10-year government bond yield series was not available in the IFS for all years in our sample period, we have used instead a series (for 1973-1998) kindly provided by Erkki Kujala, Bank of Finland, to whom we owe our thanks.

[22] We tried obtaining homogeneous interest rate series by contacting various central banks and statistical offices, but ran into insurmountable "objective" obstacles (e.g. the raw statistical data do not always exist); see fn. 15.

[23] Applying the following formula for country i: $I_{i,t} = \log(1+IR_{i,t})$ where $IR_{i,t}$ is the reported interest rate. In this respect, another data problem with the interest rate series is whether interest is paid once per year or at a higher frequency, about which the data sources say little.

The sample is thus a *cross-section* of *time-series* trends and averages. It consists of 18 (shifting) observations,[24] 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*. In our case, 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 the sample mean or trend values of various annual national time series, means that a lot of information about short-term dynamics is lost. Here, we are however *solely* interested in estimating a set of long-term equilibrium conditions, with a sample including as many 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.[25] Put differently, we prefer to filter out what is sure to include a lot of irrelevant high-frequency noise and we are not interested in estimating a short-term relationship, possibly an artifact, such as (e.g.) $D_{A/B,t} = a + b(I_{A,t} - I_{B,t}) + \varepsilon_{1,t}$, with t = month or quarter – or any dynamic or more sophisticated version thereof. To repeat: we want to estimate directly a set of long-term equilibrium conditions without short-term noise getting into the way. Even so, it will be hinted in the concluding section that our results may have some relevance in a short-term context too. Lastly, we are focusing here on *realized* outcomes and not on any ex ante relationships.[26]

Also consider this: given that were are dealing with long-term equilibrium conditions resting on arbitrage, the *functional form* of the equations is known with certainty, the problem thus being restricted to estimating a set of parameters and testing a number of hypotheses. The traditional, structural-type econometric approach can therefore be applied in a straightforward manner, since the desirable properties of the estimators involved do not, under these circumstances, depend in any way on the variables being stationary or not.[27]

2.1 Taking the USA as the Reference Country: OLS, WLS and ODR Results

All three parity conditions will be tested, even though the real interest one has been derived above from the other two.[28] This testing is however not as straightforward as it might seem, because all conditions rest on an *arbitrage* mechanism. Taking for example

(7) $D_{A/B} = (R_A - R_B) + (I_A - I_B) + \varepsilon_1$,

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

(8) $D_{i/USA} = a + b(I_i - I_{USA}) + \varepsilon_{1,i}$ (i = country) ,

[24] "Shifting" because all 18 countries will be used in turn as the reference-currency one.

[25] Higher frequency – e.g. quarterly or monthly – data are in any case not available for all variables and for all countries in the sample: thus, many countries do not have quarterly national accounts and hence GDP deflators for the full 1976-1998 period. Hence, with a sample of 23 annual observations for each country, no co-integration tests can be performed. In a sense, we trade smaller time-series for a much broader country coverage.

[26] However: "Although one does not observe the expected [i.e. ex ante] real interest rates, they can be approximated in a variety of manners in empirical analyses. The first is to use the unbiasedness hypothesis, and calculate ex post real interest rate differentials." (Fujii-Chinn, 2000, 6; also see Obstfeld-Taylor, 2000, 2).

[27] If one examines the proofs of these properties (unbiasedness, efficiency, etc.), at no point is the hypothesis of stationary variables called upon. – Another way to look at this issue : if the functional form of an equation is known with certainty, the variables involved are *necessarily* co-integrated.

[28] It is not clear from the literature whether real interest rate parity is really a separate condition "in its own right"; we shall return to this question later on and conclude that all three equations must be estimated.

where, as explained further on, â includes – but is not necessarily equal to – the estimated average risk premium of the USA versus the other countries, and where the first null is a=0 and the second b=1. Or whether one should rather estimate the reverse relationship[29]

(9) $(I_i - I_{USA}) = a' + b'(D_{i/USA}) + \varepsilon'_{1,i}$,

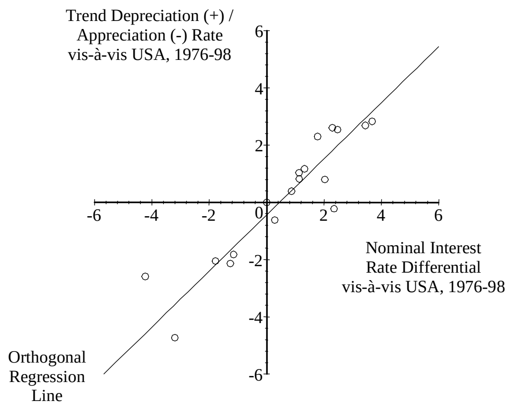
which will yield different numerical estimates[30].

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.[31] Following the discussion in Maddala (1992, 74-76, 447-72), we shall consequently estimate in all cases both an equation like (8) and one like (9), the results to be interpreted, according to the same author, as "bounds" around the true value of the parameters.[32]

Orthogonal distance regression (ODR) rather than OLS and WLS would seem an obvious choice in such circumstances.[33] This method appears to be quite popular in other sciences such as medicine or engineering, where it is sometimes claimed that it allows the error-in-data problem to be sidestepped; this is however only partially true as it turns out that the ratio of the error variances must be supplied extraneously.[34] Moreover, orthogonal estimators have infinite higher moments[35] (at least in the case of linear models[36]), so that no hypothesis testing can be done and no confidence intervals can be constructed. Nevertheless, we shall also supply ODR estimates, which will of course lie between the two bounds mentioned above.[37]

Table 2 lists the OLS, WLS and ODR results for the risk-inclusive nominal interest parity condition, the reference country being the USA.[38] Graph 1 gives an impression of the sample.

Graph 1: Uncovered Nominal Interest Rate Parity (18 OECD Countries)



[29] With analogous nulls, and of course comparing â with (–â'/ $\hat{b}$ '), and $\hat{b}$ with (1/ $\hat{b}$ ').
[30] At least with *single-equation* estimation methods such as OLS and WLS, but not with FIML. This section concentrates on OLS, WLS and also ODR (see below) results. FIML results are given in section 2.3.
[31] 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 – see below.
[32] Maddala cautions that these "bounds" should not be misinterpreted as confidence intervals since the estimated bounds have standard errors.
[33] Methods of estimation other than OLS, WLS and ODR are discussed in section 2.3.
[34] For example, see Ammann-Ness (1988).
[35] See Anderson (1976, 1984) as quoted in P.T. Boggs *et al.* (1988,172).
[36] See Boggs-Rogers (1990).
[37] As the standard orthogonal regression package available on the web ("ODRPACK") could not be rendered operational for some mysterious reason, we have used a simple estimation program of our own, based on an algorithm in Malinvaud (1970, 35-39). This *Gauss* program is available on request.
[38] And taking for the WLS weights the 1990 values of the various countries' GDP converted into a common currency by means of the 1990 PPP exchange rates as calculated by the OECD.

Table 2
First Parity: The Risk-Inclusive Uncovered Nominal Interest Rate Condition

$$\underbrace{D_{i/USA}}_{\mathbf{Y}} = a + b\underbrace{(I_i - I_{USA})}_{\mathbf{X}} + \varepsilon_i$$

	Regressing **Y** on **X**[a]		Orthogonal	Regressing **X** on **Y**[a]	
	OLS	WLS[c]	Regression[a,b]	OLS	WLS[c]
Slope ($\hat{b}$)	**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 ($\hat{a}$)	**-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	**0.83**	**0.75**	**0.83**	**0.83**	**0.80**
F prob. value	0.00	0.00	0.00	0.00	0.00

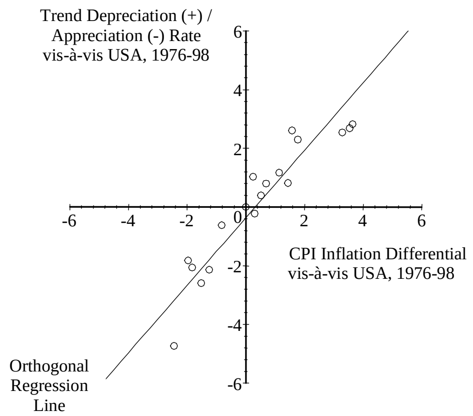
a/ **All values are given for the Y = a + bX relationship**. The t-statistics and confidence intervals from the X-on-Y regressions were transformed so as to be comparable to the Y-on-X results by applying a Taylor expansion and the Delta-method; see the **Appendix**.
b/ Unweighted.
c/ For the weights used, see note 38.

As the table shows, the data do not reject the b=1 null, i.e. the theoretically expected value. In other words, the uncovered interest rate parity condition stands verified on average, in the long run and ex post. In a pure cross-section context and with a sample of 18 observations, adjusted r^2's of around 0.8 would seem rather comforting too.[39] As to the various point estimates of **b**, those resulting from the X-on-Y regressions may be here preferable to those from the Y-on-X regressions because the interest rate differentials are more likely to suffer from a serious error-in-data problem than the depreciation differentials – see above. Be that as it may, it is striking that the central ODR point estimate of **b** is almost exactly unity. It was mentioned above 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 discussing it to section 2.2. 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.[40]

Table 3 lists our results for the relative PPP condition, again taking the USA as the reference country. Graph 2 gives a visual impression of the sample when the CPI's are used to measure inflation differentials, and graph 3 when the GDP deflators are taken instead.

[39] Graph 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, 1998, 299-300, 304-5; Hayashi, 2000, 21-3); the results of this analysis are available on request.
[40] The irrelevance of a Wald test in this context seems to be often overlooked in the empirical literature.

Graph 2: CPI-Based Purchasing Power Parity



Graph 3: GDP Deflator-Based Relative Purchasing Power Parity

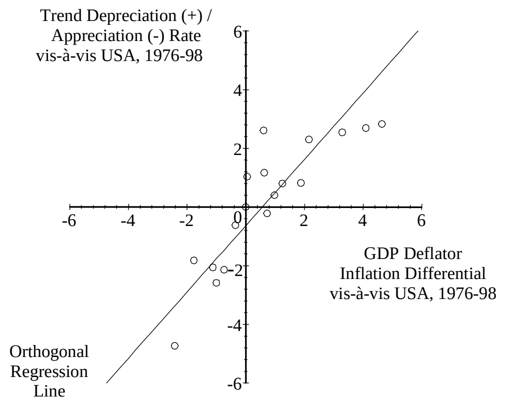


Table 3
Second Parity: The Relative PPP Condition

$$\underbrace{D_{i/USA}}_{\mathbf{Y}} = a + b\underbrace{(CP_i - CP_{USA})}_{\mathbf{X}} + \varepsilon_i$$

	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 ($\hat{b}$)	**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 ($\hat{a}$)	**-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	**0.74**	**0.88**	**0.74**	**0.74**	**0.91**
F prob. value	0.00	0.00	0.00	0.00	0.00
B. Taking the CPIs					
Slope ($\hat{b}$)	**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 ($\hat{a}$)	**-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	**0.87**	**0.94**	**0.87**	**0.87**	**0.95**
F prob. value	0.00	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 as explained in the **Appendix**.
b/ Unweighted.
c/ For the weights used, see footnote 38.

Again, the data do not reject the hypothesis that b=1, i.e. the theoretically expected value. The PPP condition in its relative form is thus also corroborated empirically in the long run, on average and ex post. That the adjusted r^2's are higher when taking the CPIs is surely due to the direct impact of the exchange rate on the CPIs (which includes the prices of imported consumer goods). The GDP deflators are the theoretically more relevant price indices since they are supposed to capture "home-grown" inflation; on the other hand, the CPIs are in general measured more precisely. The estimated constants are discussed in section 2.2.

Table 4 lists our results for the real interest rate parity condition, still taking the USA as the reference country.[41] Graphs 4 and 5 give a visual impression of the sample.

Table 4
Third Parity: The Real Interest Rate Condition

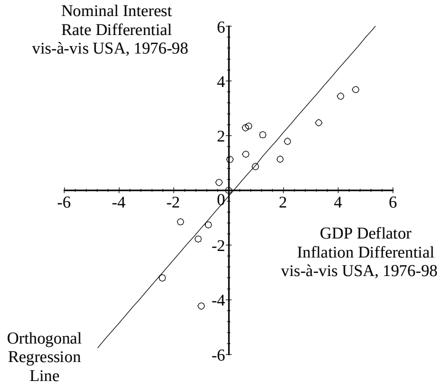
$$\underbrace{(I_i - I_{USA})}_{\mathbf{Y}} = a + b\underbrace{(CP_i - CP_{USA})}_{\mathbf{X}} + \varepsilon_i$$

	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 ($\hat{b}$)	**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 ($\hat{a}$)	**-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	**0.71**	**0.74**	**0.71**	**0.71**	**0.76**
F prob. value	0.00	0.00	0.00	0.00	0.00
B. Taking the CPIs					
Slope ($\hat{b}$)	**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 ($\hat{a}$)	**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	**0.77**	**0.74**	**0.77**	**0.77**	**0.73**
F prob. value	0.00	0.00	0.00	0.00	0.00

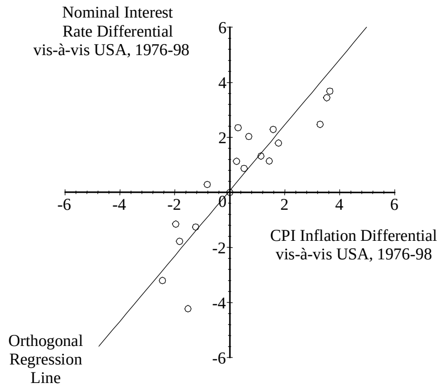
a/ **All values are given for the Y = a + bX relationship**. The t-statistics and confidence intervals from the X-on-Y regressions were transformed as explained in the **Appendix**.
b/ Unweighted
c/ For the weights used, see footnote 38.

[41] Notice that the estimated equation (top of table 4) is different from equation (4). In the latter case, taking the USA as the reference country would mean that the dependent variable ($I_{USA} - CP_{USA}$) is…a constant. Consequently, the equation was re-specified as indicated on top of table 4. In this specification, it is a version of the Fisher equation.

Graph 4: GDP Deflator-Based Real Interest Parity



Graph 5: CPI-Based Real Interest Parity

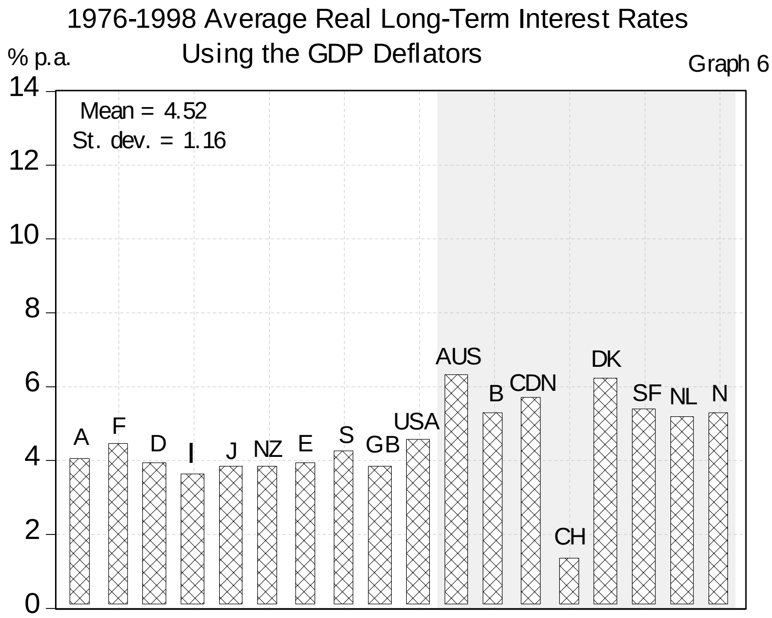


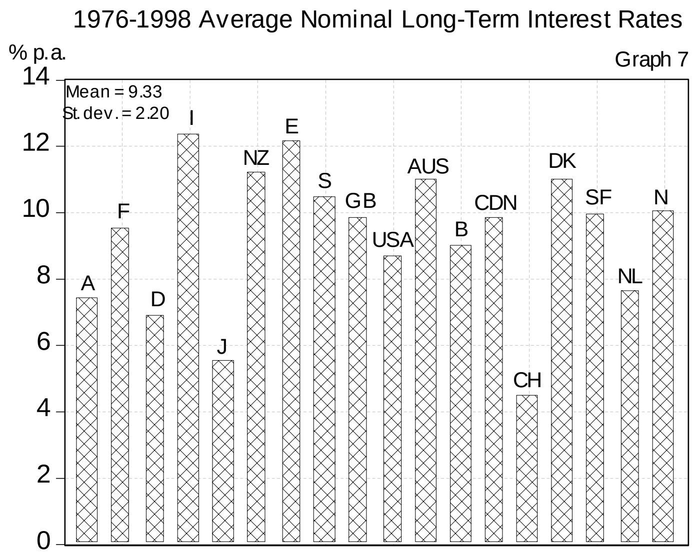
Here too, the theoretically expected value for the slope, i.e. b=1, always lies within the 95 and 99% confidence intervals, so that the real interest rate condition is also corroborated by the data. The estimated intercepts will be discussed in section 2.2. It was mentioned above that it is not clear 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 uncovered nominal interest rate condition is clearly of the essence: she will compare the nominal rate of return on home and foreign assets allowing for the expected evolution 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.[42] 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.[43] We shall again return to this issue in section 2.2.

The triple-parity law says that, except for inter-country risk and productivity differentials, the real interest rate should tend to become equalized in all countries, on average, in the long run

[42] On this, see Marston (1997).
[43] Obstfeld-Taylor (2000, 1) offer another explanation : “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”.

and ex post. This is illustrated in graph 6.[44] 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. Country-specific factors seem important for eight countries (shaded area in graph 6), about which more in the next section where the significance of these deviations will be tested. It is noteworthy that the pattern for nominal interest rates (graph 7) is quite different from that for real interest rates (graph 6), and that the former are more widely spread out than the latter: moving from nominal to real interest rates reduces the standard deviation by almost one half.





2.2 The "R" and "T" Factors

In this section, we shall look at the *significant* deviations from the triple-parity law which we observe in the case of some countries – actually, a small number of small countries, as will be seen. In some sense, these cases should therefore be considered as anomalies which do not

[44] Countries are identified by the country tags one sees on automobiles: A = Austria; F = France; D = Germany; I = Italy; J = Japan; NZ = New Zealand; E = Spain; S = Sweden; GB = United Kingdom; USA = United States of America; AUS = Australia; B = Belgium; CDN = Canada; CH = Switzerland; DK = Denmark; SF = Finland; NL = Netherlands; N = Norway.

gainsay the general validity of the triple-parity law. Prior to that, some further considerations on our estimation methods are in order.

So far, the USA has been taken as the reference country, which is an arbitrary although perhaps natural choice. If another country is selected as the reference one, so that all differentials are computed with respect to that other country, it will change the estimated constants, but *not* the estimated gradients and their associated statistics, which remain exactly the same as in the above tables.[45] 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, all equations are really log-log ones. If the measurement unit of one or several variables is changed in a log-log equation, it will alter the estimated constant only. Thus, changing the reference country is similar to changing the measurement unit.

This point however calls for some additional explaining. When estimating any one among our equations, a natural impulse would be to take as a sample the various differentials for 17 countries out of a total of 18, i.e. omit the reference country altogether. If this is done, the various **b** estimates however turn out to be different depending on which country is taken as the reference one, although the differences are always very small; in particular, the relevant 95% confidence intervals always overlap, and all always include the unit value for the estimated **b**'s. In other words, the triple-parity law remains verified. It is however disturbing from a theoretical standpoint that the – purely "nominal" – choice of the measurement unit should have an effect on the estimated slopes, however small it may be. But if each relevant sample is extended to include the reference country, whose differentials are identically equal to zero, which is what has been done for the estimations in tables 2-4,[46] the theoretically expected result obtains: the estimated slopes and all associated statistics are then exactly the same whatever country is selected as the reference one. The best way we can think of to justify this point is by means of an analogy.

Image a pure barter economy. One good is selected as the numéraire, its price being identically equal to one. We then have a vector of real prices for all goods. Now suppose we want to find out whether there is any significant relationship between the price vector and some other vectors measuring, for example and among other relevant factors, the degree of competition in the various goods markets. There would then be no reason to omit the price of the numéraire good: although it is identically equal to unity, it conveys information that should not be arbitrarily thrown away and it should therefore be included in the sample. By analogy, there is no more reason to exclude the reference country from our various samples.[47] But to repeat: its inclusion or exclusion has no impact on the general empirical validity of the triple-parity law.

Including the zero differentials for the reference country also has an interesting econometric consequence. Take any equation where the USA is the reference country. Given that both the RHS and 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.[48] The same obtains when each one of the other countries is taken in turn as the reference country. Consequently, the various estimated constants and the various reference-country *own* residuals are one and the same

[45] This can be checked easily with our data bank, which is available on request.
[46] Also see graphs 1-5.
[47] Since our equations are really log-log ones, a zero differential is similar to the unit price for the numéraire good in our analogy: the anti-log of zero is one.
[48] Given $y_i = \hat{a} + \hat{b}\, x_i + \hat{e}_i$, $y_i = x_i = 0$ means $\hat{a} = -\hat{e}_i$.

thing, and we do not need to show them separately.[49] Table 5 lists them for the first equation, i.e. the uncovered nominal interest rate parity condition.

Table 5
Uncovered Nominal Interest Rate Parity: The "R" Factors

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	
	â	p-value	â	p-value	â	p-value	â	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**.

It was stated earlier, perhaps a little enigmatically, that each estimated constant includes – but is not necessarily equal to – the estimated average country-specific risk premium. This is so because an estimated constant can be non zero for three non mutually exclusive reasons: (1) $E(\varepsilon_i) =$ some constant $\neq 0$: the measurement errors affecting the dependent variable include a systematic bias which will show up in the estimated intercept; (2) even when $E(\varepsilon_i)=0$, genuinely random shocks will in general not average out to zero in any finite sample; (3) there is a non zero average risk premium differential for the country under consideration. Given that our cross-section sample is made up of trend or mean values 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 examining the results listed in table 5 – especially, it would seem, those for the Y-on-X regressions, since appreciation/depreciation rates are likely to be measured with a high degree of precision, but not so our long-term interest rate series, as argued previously.[50]

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[49] 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. Ideally, all these cross-section residuals should be checked as to their randomness, but – to our knowledge – no explicit tests for auto-correlation are available in a pure cross-section context, and only "eyeballing" can be resorted to. 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. The various residual series and the corresponding graphs are available on request.

[50] On the other hand, the reverse X-on-Y regressions would then suffer from a serious error-in-data problem, meaning that the slope estimates might be biased – and hence also the estimated constants. Consider a two-

Perhaps the most striking thing about the results in table 5 is that they are quite sensitive to the direction of regression and to the estimation method (OLS vs. WLS), as the following correlation matrix confirms.

r	Y on X OLS	Y on X WLS	X on Y OLS	X on Y WLS
Y on X OLS	1.00	0.98	0.91	0.74
Y on X WLS	0.98	1.00	0.82	0.60
X on Y OLS	0.91	0.82	1.00	0.95
X on Y WLS	0.74	0.60	0.95	1.00

This sensitivity to the direction of regression and the estimation method is even more noticeable when considering the *significance* of the estimated constants: had we limited ourselves to the direct OLS regressions, i.e. the first **â** vector in table 5, we would have concluded that the intercept is significant (at the 5% level) for 9 countries out of 18: Switzerland, Sweden, Australia, New Zealand, Belgium, Austria, Finland, Japan and Denmark – all of them small economies, except Australia and Japan.

However, looking across the various **â**'s for each country, and admitting as a rule of thumb that there are "strong indications" of an overall significant intercept when the **â**'s are significant in at least three cases out of four, we find relevant results for four small countries only: Switzerland, Sweden, Finland and Denmark – the latter two being the only ones for which the constants are in all cases significantly different from zero. Moreover, the four point estimates are closely bunched for Sweden, Finland and Denmark, but not for Switzerland. This illustrates the importance of the regression-direction and estimation-method issues when using single-equation estimation methods.

Bearing in mind the preceding caveats about possible error-in-data problems, a small (i.e. negative) intercept is indicative of a small risk premium differential for the country under consideration or, more generally, it is indicative of an international comparative advantage, as argued previously. This is best seen when considering the reverse relationship: a small 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.[51] Seen in this light, the results in table 5 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 real interest rate parity equations. A proviso should however be stated right away: if the risk premium for a given country has remained constant over the sample period, the interpretation (e.g. in terms of structural factors, policies, etc.) is likely to be more straightforward than if it has been changing. Neither will there be a specific econometric problem on this account. If the risk premium has changed over time, but without being correlated with the independent variable, no specific econometric problem arises either, but the estimated constant is then a measure of the *average* risk premium over the sample period

variable, linear and upward-sloping equation: if the estimated slope is biased, say, upwards, the estimated constant will necessarily be biased downwards.

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

and it says nothing about what it might be today or in recent years. In the case of the direct regressions, the national interest rates are widely flucutating but stationary series whereas the risk premium is likely to change – if it changes – slowly and smoothly over time, and the two are therefore unlikely to be significantly correlated. We have a problem if (a) the changing risk premium and the interest rate differentials nevertheless happen to be correlated;[52] and/or (b) if, as stated above, $E(\varepsilon_i)$ = some (sizable) constant $\neq 0$ (i.e. measurement error with bias).

The same general comments apply to the results in table 6, i.e. the constants in the equations for the real interest rate parity conditions.[53] Using the same rule of thumb as above, we now find that six countries appear to be at a significant comparative advantage (Switzerland, the UK) or disadvantage (Belgium, Finland, Denmark and Australia). The point estimates are however closely bunched only in the case of Finland, Denmark and Australia.

Table 6
Real Interest Rate Parity Condition: The Combined "R" and "T" Factors

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	
	$\hat{a}$	p-value	$\hat{a}$	p-value	$\hat{a}$	p-value	$\hat{a}$	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**.

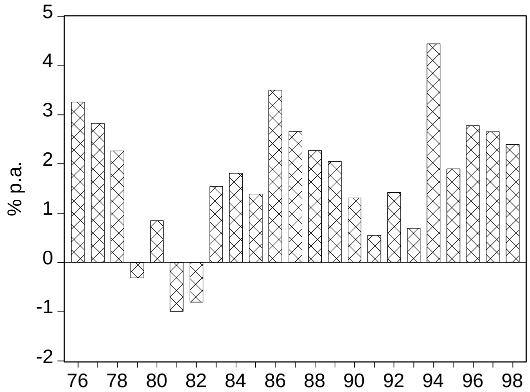
Summing up our results for both the nominal and the real interest rate parity conditions, we therefore conclude that only Denmark, Finland and Switzerland appear to constitute significant anomalies on *both* counts.[54] Why should that be so? E.g. why should Finland and Denmark appear to be at a significant disadvantage on both the nominal and the real interest rate count whereas Switzerland would seem to enjoy a significant comparative advantage? In

[52] For the inverse regressions: if the changing risk premium is correlated with the depreciation rate.
[53] In table 6, the sign of the estimated intercepts have 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.
[54] To be more precise: there is significant econometric evidence that these three countries constitute anomalies. Other countries may be anomalous too, but our data may not enable us to identify them.

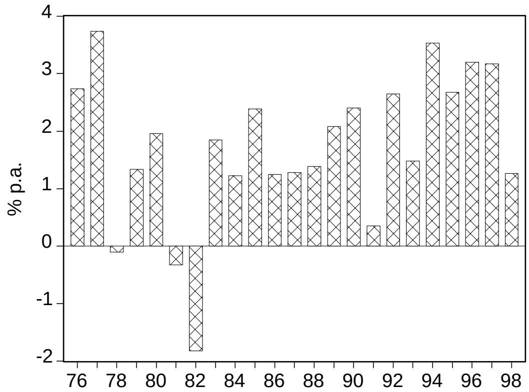
the case of Switzerland, the explanation is most likely to be found in that country's reputation for political and economic stability, its efficient financial sector as well as its banking secrecy law and its status as an international tax heaven. Here is not the place to go into the pros and cons of 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). As to Finland and Denmark, the issue is currently being discussed with people who know more about them than we do, the outcome of these contacts to be reported in the final version of our paper. For all three countries, an important issue is whether or not their relative risk premia have been more or less constant over the sample period. Graphs 8-13 may shed some light on this issue, focusing as they do on ex post real annual interest rates (notice the different scales).

Graphs 8-13

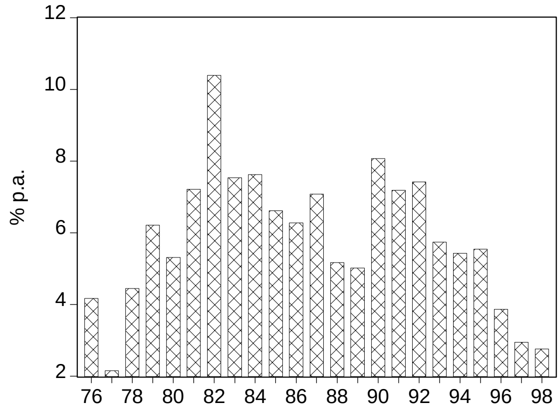
Switzerland: Long-Term CPI-Defined Real Interest Rate



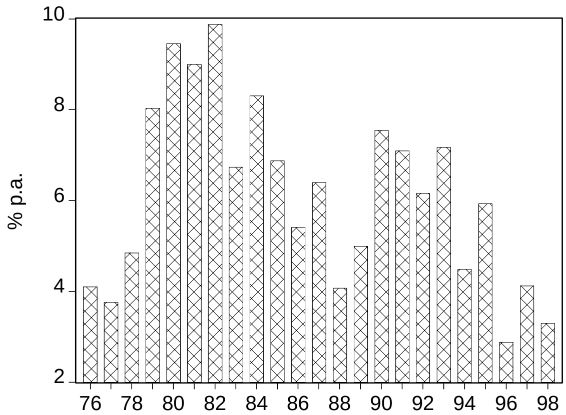
Switzerland: Long-Term GDP Deflator-Defined Real Interest Rate



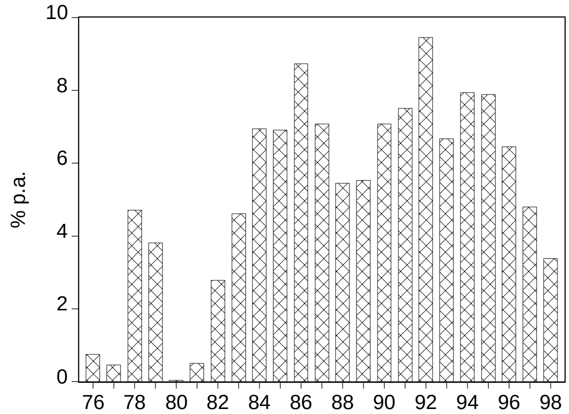
Denmark: Long-Term CPI-Defined Real Interest Rate



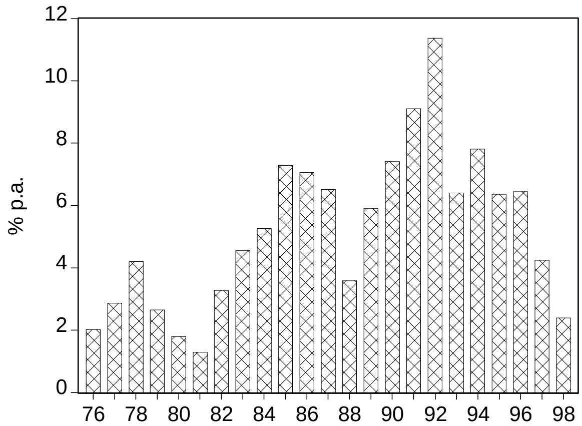
Denmark: Long-Term GDP Deflator-Defined Real Interest Rate



Finland: Long-Term CPI-Defined Real Interest Rate



Finland: Long-Term GDP Deflator-Defined Real Interest Rate



These graphs suggest that Switzerland's low risk premium is a permanent factor whose expected value may have been approximately constant over the sample period. In the case of Denmark, its high average risk premium appears to be linked with high real interest rates in the 1980's and in the first half of the 1990's (the Danish economy was then 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 risk 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.

We now turn to the estimated constants (or residuals) in the PPP equations. 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 7 is indicative of low productivity growth, 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 lower-than-expected productivity growth: Australia, Canada, Norway and Sweden, all of them important producers and exporters of primary commodities and raw materials. The estimated intercept is significant in all four cases and the points estimates lie close to one another. At the other end, productivity seems to have grown at a significantly faster-than-expected pace in the UK, Spain and Italy. The points estimates are closely bunched for the UK, but they are small. They are larger for Spain and Italy, but also rather spread out. The explanation may be that these three countries have undergone especially rapid modernization in the sample period.

Table 7
Relative PPP Condition: The "T" Factors

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	
	â	p-value	**â**	p-value	**â**	p-value	**â**	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
inland	**-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**.

Looking back at equation (4), it is seen that the constant in the real interest rate equation is equal to the productivity growth differential in the PPP equation minus the risk premium differential in the nominal interest rate equation: a country will enjoy a comparative *real* interest rate advantage if its productivity growth differential is larger than its nominal interest rate risk premium (which is an interesting proposition in itself). This also affords us a way to check whether the estimated country-specific constants in the real interest rate equations (table 6), which we shall call the *direct* estimates of the real interest rate risk premia, are consistent with the differences between the estimated productivity growth differentials (table 7) and the estimated nominal interest rate risk premia (table 5), these differences to be dubbed here the *indirect* estimates of the real interest risk premia – see table 8.

Table 8

OLS Estimates: Y on X

		Indirect	Direct	Difference[a]
1	Australia	-1.60	-1.77	-0.17
2	Austria	0.28	0.48	0.20
3	Belgium	-0.74	-0.70	0.04
4	Canada	-1.11	-1.15	-0.04
5	Denmark	-1.54	-1.72	-0.18
6	Finland	-0.74	-0.89	-0.15
7	France	0.03	0.00	-0.03
8	Germany	0.37	0.63	0.26
9	Italy	1.06	0.73	-0.33
10	Japan	0.39	0.80	0.41
11	Netherlands	-0.82	-0.62	0.20
12	New Zealand	0.84	0.63	-0.21
13	Norway	-0.71	-0.78	-0.07
14	Spain	0.74	0.43	-0.31
15	Sweden	0.34	0.22	-0.12
16	Switzerland	2.69	3.19	0.50
17	UK	0.65	0.59	-0.06
18	USA	-0.14	-0.07	0.07
	Mean	0.00	0.00	0.00
	S.D.	1.06	1.15	0.23
	Correlation coefficient 0.98			

OLS Estimates: X on Y

		Indirect	Direct	Difference[a]
1	Australia	-1.92	-1.81	0.11
2	Austria	0.15	-0.06	-0.21
3	Belgium	-1.01	-1.10	-0.09
4	Canada	-1.40	-1.40	0.00
5	Denmark	-1.83	-1.72	0.11
6	Finland	-0.82	-0.69	0.13
7	France	0.06	0.10	0.04
8	Germany	0.23	-0.05	-0.28
9	Italy	1.73	2.18	0.45
10	Japan	0.10	-0.36	-0.46
11	Netherlands	-1.27	-1.54	-0.27
12	New Zealand	1.30	1.58	0.28
13	Norway	-0.86	-0.81	0.05
14	Spain	1.28	1.67	0.39
15	Sweden	0.58	0.75	0.17
16	Switzerland	3.00	2.56	-0.44
17	UK	0.92	1.02	0.10
18	USA	-0.25	-0.34	-0.09
	Mean	0.00	0.00	0.00
	S.D.	1.32	1.36	0.26
	Correlation coefficient 0.98			

a/ Direct minus indirect estimates.

It is seen that the direct and indirect estimates are actually quite close, being highly correlated. It is also seen that the differences between them average out to zero and that none of these differences is 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 (see above), we should rather expect the direct estimates to be different from the indirect ones because these agents' perceived risk premia may be different from that of other investors or firms. But this ignores that *all* agents operate on the same financial and goods markets where their interactions result in average market-wide premia. No matter how we look at it, the deterministic part of equation (1) and the deterministic part of equation (2) *necessarily* imply the deterministic part of equation (3) : if the uncovered interest rate parity and the PPP equilibrium conditions hold, the real interest rate parity condition *must* hold too. It follows that, seen in this light, the real interest parity condition 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 results in separate *shocks*, so that equation (3) should be rewritten as:

(10) $(R_A - R_B) + (I_A - I_B) = (T_B - T_A) + (CP_A - CP_B) + (\varepsilon_2 - \varepsilon_1) + \varepsilon_3$

Under these conditions, it is *econometrically* justified to test the real interest parity relationship as a separate condition, as we and others have done. Furthermore, it is also possible that the existence of this separate class of arbitrageurs, supposing it really exists, will reinforce and speed up the realization of the uncovered interest rate and PPP conditions.

2.3 Other Estimation Methods

Besides OLS, WLS and ODR, our three basic equations have also been estimated by the following methods: (1) White’s heteroskedasticity-consistent method; (2) Zellner’s method for seemingly unrelated equations (SUR); (3) full information maximum likelihood (FIML), with Marquardt’s algorithm.

Concerning (1), the USA-based equations were checked for heteroskedasticity, with a few testing positive.[55] They were consequently re-estimated by White’s method. The estimated **b** coefficients are not changed hereby, of course, but their standard errors are: the latter turn out to be somewhat larger in some cases and somewhat smaller in others, but all the confidence intervals continue to support the basic E(b)=1 hypothesis – and hence the triple-parity law.[56]

For reasons that will be seen presently, the SUR and FIML methods must be discussed together.

FIML – being a system method of estimation – has the nice property of invariance (see e.g. 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 discuss them, let us simplify our notation and rewrite our three basic relationships as a system of equations:

(11) $DUSA_i = c(1) + c(2)*IUSA_i + \varepsilon_{1,i}$

i.e. the uncovered nominal interest rate parity condition (the *UIP equation*), with $DUSA_i$ = average depreciation rate of country i’s currency w.r.t. the dollar, and $IUSA_i$ = i’s average nominal interest differential vis-à-vis the USA,

(12) $DUSA_i = c(3) + c(4)*CPUSA_i + \varepsilon_{2,i}$

i.e. the relative PPP condition (the *PPP equation*), with $CPUSA_i$ = country’s i’s average inflation rate differential w.r.t. the USA,

(13) $IUSA_i = c(5) + c(6)* CPUSA_i + (\varepsilon_{1,i} - \varepsilon_{2,i}) + \varepsilon_{3,i}$

i.e. the real interest rate parity condition (the *RIP equation*).

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 (it includes $\varepsilon_{3,i}$ 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

[55] Given that the general triple-parity law has to do with the **b** coefficients, and that these do not change when the reference country is changed, only the 10 USA-based OLS equations were tested.

[56] Results available on request.

hence no reduced form.[57] Consequently, the FIML method cannot be applied to our full system of three equations.[58] If we want to use FIML, we must instead take them two by two, for a total of three combinations.

Doing this however re-introduces the question of normalization, and hence that of the direction of causality. If, for example, we try to apply FIML to (11) and (12) as normalized above, i.e. with the same dependent variable, the algorithm breaks down[59]. 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 (11)-(12) must therefore be renormalized, which – can be done in two ways:

(14) $DUSA_i = c(1) + c(2)*IUSA_i + \varepsilon_{1,i}$
(15) $CPUSA_i = -[c(3)/c(4)] + [1/c(4)]*DUSA_i - [1/c(4)]*\varepsilon_{2,i}$, or

(16) $IUSA_i = -[c(1)/c(2)] + [1/c(2)]*DUSA_i - [1/c(2)]*\varepsilon_{1,i}$
(17) $DUSA_i = c(3) + c(4)*CPUSA_i + \varepsilon_{2,i}$

Both systems are recursive; or they are only triangular if the two ε's are not independent,[60] 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 (14), and then $CPUSA_i$ is determined by $DUSA_i$ via (15).[61] Since (14) 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 (15) will however be different from the OLS ones. Given that the deterministic parts of (14)-(15) 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 (16)-(17). 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 (16)-(17) for c(1) and c(2), and those
- given by (14)-(15) for c(3) and c(4),

and we need not reproduce the OLS point estimates again.[62]

The same reasoning holds for the other two combinations of equations (11)-(13) taken two by two.[63] Table 9 gives the results for all three combinations.

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[57] This does not mean, of course, that country *i* has no choice as to its inflation rate, its nominal interest rate level and the depreciation tendency of its currency: all variables in the system are *differentials* w.r.t. the country of reference – see table 1.
[58] The system's parameters might perhaps be estimated by some numerical methods such as are used in (e.g.) physics, but we did not further pursue this matter, in particular because we are not sure to what extent these methods lend themselves not only to parameter estimation, but also to classical hypothesis testing.
[59] "Cannot compute likelihood function derivatives at current parameter values", where it makes no difference what initial guesses are supplied for the parameter values, e.g. c(1)=c(3)=0 and c(2)=c(4)=1, or any other values.
[60] And so is the system (11)-(13) if (11) is renormalized so as to determine $CPUSA_i$ – see below.
[61] Notice that, by the order condition, both equations are exactly identified.
[62] The OLS standard errors are different, as said, but all coefficients remain significant in the triple-parity law sense. Hence we dispense with reproducing them.

Table 9
SUR/FIML Results[a]

Estimated equations taken two by two[b]	**Slope ($\hat{b}$)**	**Constant ($\hat{a}$)**	**Adj. r^2**
(11) UIP	**1.00** (10.3;3.0)	**–0.46** (–2.1;–1.1)	0.82
(12) PPP	**1.20** (8.0;4.6)	**–0.69** (–2.4;–1.4)	0.74
(11) UIP	**1.00** (10.9;3.0)	**–0.46** (–2.1;–1.1)	0.82
(13) RIP	**1.19** (7.7;3.6)	**–0.23** (–0.8;–0.5)	0.70
(12) PPP	**1.20** (9.9;4.6)	**–0.69** (–2.5;–1.4)	0.70
(13) RIP	**1.19** (8.9;3.6)	**–0.23** (–0.8;–0.5)	0.66

a/ t-statistics in parentheses: 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.

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. Comparing the results in table (9) with the OLS ones in tables 2, 3 and 4 also shows that using system estimation methods leads to a significant improvement:[64] 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 an even stronger verification of the triple-parity law.

As stated above, the three-equation system (11)-(13) cannot be estimated as such by FIML, because it is a closed one, but it can be estimated by SUR if (12) is renormalized so as to determine $CPUSA_i$, thus making the deterministic parts of (11)-(13) a fully triangular system – see table 10, which also gives the OLS results for comparison purposes.

[63] Combination (10) and (12) can be estimated as it is because the system is triangular. For the reasons indicated in the text, the estimated coefficients (but not the standard errors) for (12) will the same as the OLS direct-regression ones in table 4.
[64] 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.

Table 10
System (11)-(13): SUR vs. OLS Results[a]

	SUR Results			OLS Results[b]		
Equation	**Slope ($\hat{b}$)**	**Constant ($\hat{a}$)**	**Adj. r^2**	**Slope ($\hat{b}$)**	**Constant ($\hat{a}$)**	**Adj. r^2**
(11) UIP	**0.97** (11.2) (0.086)	**–0.43** (–2.1) (0.206)	.82	**0.90** (9.1) (0.099)	**–0.39** (–1.8) (0.219)	.83
(12) PPP[c]	**1.14** (9.5) (0.121)	**–0.65** (–2.5) (0.263)	.73	**1.28** (7.1) (0.180)	**–0.75** (–2.3) (0.321)	.75
(13) RIP	**1.13** (9.4) (0.119)	**–0.19** (–0.7) (0.278)	.69	**0.97** (6.5) (0.149)	**–0.07** (–0.2) (0.300)	.71

a/ t-statistics and standard errors 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 regression for UIP and RIP)
and table 3 (reverse OLS regression for PPP)
c/ Renormalized – see text. Equation estimated as: cpusa = -(c(3)/c(4)) + (1/c(4))*dusa

It is seen that for the UIP and PPP equations, but regrettably 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 fully stand up under SUR too. Finally, the SUR full-system results in table 10 are also distinctly better than the SUR/FIML two-equations-at-a-time results in table 9, in the sense that the standard errors for the slope coefficients are smaller.

3. The Triple-Parity Law and the Existing Literature

Let us first focus on three recent and particularly relevant studies.

The closest conceptual proximity to our integrated three-relationship analysis we are aware of is to be found in Marston (1997[65]). But this author does not quite go as far as to characterize the triple parity as a general economic law.[66] On the empirical side, he 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 research where their interdependency is stressed. However, Marston also performs ex post tests of the uncovered nominal interest rate condition, of the relative PPP condition, and hence of the real interest rate parity condition too, using the same monthly and quarterly time series for the G-5 countries[67] over the June 1973–December 1992 period as in his ex ante tests. These ex post tests confirm that all three equilibrium conditions ultimately hold, on average, for the G-5 countries.

[65] Also see 1995, Ch. 6.
[66] Moreover, he apparently does not all allow for the "T" factors in the PPP relationship and performs a number of joint Wald tests (which, we have argued, are irrelevant).
[67] France, Germany, Japan and the UK, with the USA used as the reference country.

Using very long historical series for eight countries,[68] Obstfeld and Taylor (2000) also find that the long-run real interest rate parity condition did hold ex post, most strikingly in the 1890-1914 gold standard period and somewhat less so in the post-1974 float. They argue that previous studies[69] had found negative results, or at least had not been very supportive, due in part to overdifferencing and filtering problems in the approach to the data. They also explicitly allow for the normalization and direction-of-causality issue.

Fujii and Chinn (2000) find that real interest rate parity holds better at longer than at short horizons, the hypothesis being decisively rejected in the latter case. Using our notation, their results show that the E(b)=1 hypothesis in the RIP relationship is not rejected for most countries in their sample when using 5- or 10-year interest rates, and either consumer or wholesale price indices.[70]

All three studies use time series data at relatively high frequencies (monthly, quarterly) and test the RIP relationship for each country in their sample taken separately (i.e. 5, 8 and 7 large countries, respectively[71]). Obstfeld-Taylor and Fjuii-Chinn, but not Marston, also take the co-integration route. We did not because, among other things, we wanted to trade a small number of annual observations in exchange for as large a sample of industrialized countries as possible. Hence, we have tested all three relationships for and *across* 18 countries.

In what follows, we summarize the theory and evidence in the rest of the existing literature that underlie each of the three equilibrium propositions making up the triple-parity law. In doing so, we put our own econometric tests and results in the context of the related empirical work.

3.1 The Law of One Price and Purchasing Power Parity

According to the "Law of One Price" (LOP), the domestic-currency prices – in any two countries and at any given date – of any pair of individual comparable goods must be the same once converted into a common currency via the spot exchange rate. The major reason for such price equalization is the free arbitrage of analogous (or homogeneous) pairs of goods that leads to a squeeze-out of any price-differential profit unrelated to quality or any other goods characteristics that would make the analogues compared different.

Purchasing Power Parity (PPP), one of the oldest concepts in economics,[72] is a generalization of LOP. There have been various ways to express PPP when testing it, but the two most common forms are absolute PPP, which is defined over some (comparable) national price index, i.e. the general price level, and relative PPP, defined instead over national price-level rates of change, i.e. inflation rates. Absolute PPP essentially states that national price levels should equalize once all goods prices in two or more economies are duly aggregated and converted into a common currency. Relative PPP, in turn, requires that nominal exchange-rate adjustment, i.e. currency appreciation or depreciation, ultimately equalizes the inflation differential between the countries compared, ex post and over a past (long) period. This is

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

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

[70] 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, note 2).

[71] Since all three studies take the USA as the country of reference, and since none seems to discuss the issue of whether that country should not also be included, the number of countries is actually reduced to 4, 7 and 6.

[72] An extensive discussion of the origins of PPP can be found in Officer (1982, Ch. 3). For briefer and more recent interpretations, one may also wish to look at the summary article of Dornbusch (1987) or the frequently cited survey by Rogoff (1996).

equivalent to a hypothesis according to which the real exchange rate is stationary.[73] Relative PPP can also be written in its ex ante form[74] if one wants to test whether (expected) depreciation and the (expected) inflation differential are equalized over some future period(s).

Early studies of LOP, such as Isard (1977) and Richardson (1978), have documented large and persistent deviations from LOP, over a broad range of analogous goods and even simple good components, that are highly correlated with exchange-rate movements. Subsequent empirical work has proceeded in three major directions: (i) inter- vs. intra-national comparisons of price volatility have found a large "border effect", e.g. Engel (1993), Engel-Rogers (1995) and Rogers-Jenkins (1995); (ii) 20th century vs. earlier centuries comparisons of price volatility have concluded that the volatility of deviations from LOP, even among highly traded goods, is remarkably stable over a period spanning the 14th through the 20th century, as documented in Froot-Kim-Rogoff (1995) for grains and diary products in Holland and England; (iii) trading frictions and nationally-segmented markets have been identified and modeled as the principal cause of deviations from LOP, and hence absolute PPP, in many more or less recent papers such as Krugman (1987), Knetter (1989, 1993, 1994), Kasa (1992), Feenstra (1995), Froot-Rogoff (1995), Betts-Devereux (1996, 2000), Devereux-Engel (1998, 2000) and Bacchetta-van Wincoop (1998, 2000).

The basic theoretical justifications for the widely documented empirical failure of LOP are well known. First, transportation costs and tariff and non-tariff barriers to international trade drive a wedge between the prices of similar goods in different countries, with the size of this wedge depending on the tradability of the goods concerned. Second, there are a number of other reasons, among which Rogoff (1996) enumerates country differences in: (i) inclusion or non inclusion of value-added taxes; (ii) profit margins (across market locations depending on competition); and (iii) product bundling. Nevertheless, for some highly traded and homogeneous commodities where free arbitrage is not hindered, such as gold, LOP does perform very well in the data.[75]

It is in general confirmed empirically that over the very long run, real exchange rates tend to conform to PPP, but with a very slow speed of convergence.[76] A second finding supported by the data is that short-run deviations from PPP are large and volatile, even for relatively homogenous classes of highly traded goods.[77] To reconcile the above two empirical facts has turned out to be quite problematic, hence the term "PPP puzzle" in the international economics literature. Most proposed explanations for short-run exchange-rate volatility have stressed the importance of financial factors. Obstfeld-Rogoff (1996) point in particular to changes in portfolio preferences, short-term asset price bubbles and monetary shocks, the latter being especially relevant in the presence of nominal rigidities that allow for (short-term) real effects of monetary policy.

In general, it has thus been hard to verify empirically, in a convincing manner, short-run LOP and PPP[78] outside a fairly small range of homogenous goods. Nevertheless, long-run macroeconomic data essentially support convergence to PPP, by roughly 15% a year,[79] a

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[73] Unless, that is, there are significant productivity-growth differentials (Dornbusch 1987).
[74] As first done in the study by Roll (1979).
[75] Table 2 in Rogoff (1996) illustrates this fact quite convincingly.
[76] Rogoff (1996) considers this to be the first basic fact about which a "surprising" degree of consensus has emerged in the empirical PPP literature.
[77] Also emphasized by Rogoff (1996).
[78] It should be noted in passing that Frenkel (1978) does find some evidence on PPP in hyperinflationary contexts, which is not surprising given the predominance of monetary shocks in such economic environments.
[79] Rogoff (1996), pp. 657-658.

speed that is much slower than average estimates of aggregate price (or other nominal) stickiness would imply.

The first econometric tests of PPP focused on rejecting or not rejecting a random walk hypothesis for the real exchange rate. Most research along this line found it difficult to reject the random walk null, hence PPP (that is, mean reversion) could not be justified econometrically. The earliest simpler studies include Roll (1979), Darby (1983), Adler and Lehmann (1983) and Edison (1983); later unit root approaches are represented by Huizinga (1987) and Meese-Rogoff (1988); and still more recent co-integration techniques follow the lead of Breuer (1994). The next wave of empirical research looks at much longer data sets, most of them employing samples of 100 years or more, thus testing in “the very long run”. Frankel (1986) succeeds in rejecting the random walk hypothesis, using Dickey-Fuller (1979) tests. Edison (1987), using an error-correction approach, obtains slightly weaker rejections. A consensus estimate of the half-life of PPP deviations, i.e. 3 to 5 years, has emerged out of the long-horizon empirical literature of the early 1990s.[80] A major problem was, however, that the long-horizon data mixed peg and float periods whereas Mussa (1986) had demonstrated that real exchange rates tend to be more volatile under flexible than under fixed regimes. A further advance was thus to distinguish between periods of peg and float, while at the same time increasing the number of countries examined. Evidence on mean reversion has in consequence increased, as confirmed by De Gregorio, Giovannini-Wolf (1994) and Wei-Parsley (1995) for tradable-sector cross-country comparisons, as well as by Frankel-Rose (1995) for floating-rate panel data.

Two remaining problems are that (i) most cross-sections include 10 to 15 OECD countries, i.e. highly industrialized ones, whereas Froot-Rogoff (1995) present evidence that the Argentinean peso has fallen sharply in real terms against both the US dollar and the British pound since the beginning of the 20th century; and (ii) that the standard practice is to calculate all real rates relative to the US dollar, which can lead – according to O’Connel (1996) – to cross-sectional dependence in time series panel data. In our triple-parity law tests we control for this (USA-centered) bias by taking each of our 18 OECD economies in turn as "the country of reference", as explained in section 1. As to the former (industrialized-sample) bias, data homogeneity limitations unfortunately do not allow us to duly account for it as well, although the crucial assumption of “sufficient” goods mobility may justify this omission.

3.2 Nominal Interest Parity: Covered and Uncovered Versions

Nominal interest parity (NIP) is the proposition that, in an environment of free capital mobility, international arbitrage by agents optimizing their asset holdings should result in equal(ized) nominal interest rates (i.e. nominal returns) on comparable financial instruments across countries. Unlike PPP, this relation is typically defined on an ex ante basis, due to the fact that it is the expectation of a potential return on an asset that matters for an investor/borrower at the moment she makes the decision to undertake or not a given financial investment/borrowing. Nevertheless, an ex post (actual) version of the same relation should ultimately materialize in real-world open economies.[81] NIP is analyzed in the literature under the form of either covered interest parity (CIP) or uncovered interest parity (UIP), depending on the consideration (availability) or not of forward exchange-rate markets (i.e. of flexible exchange-rate arrangements).

[80] This is one issue on which our approach cannot say anything.

[81] “Open” in the sense of having completed (i) deregulation of national financial markets (for more on this point, see Marston 1995, Ch. 2), and (ii) liberalization of national capital controls (for further detail, see *ibid.*, Ch. 3).

Even CIP, i.e. the concept which is easiest to test, may fail empirically for four principal reasons, well-known in the literature and originating essentially in different national financial transactions regulations:[82] (i) differences in default risk; (ii) regulated (hence, non-competitive) financial markets; (iii) capital controls;[83] and (iv) differences in political (or sovereign) risk. Reasons (iii) and (iv) are often unified in a joint concept which gives rise to what is known as "country premium/discount", a notion which we have discussed and estimated in section 2 under the name of country-specific average risk factors, however given them a broader interpretation (comparative advantage/disadvantage).

In many tests of CIP, data from the eurocurrency markets, which operate free of any capital controls since as early as their emergence in the late 1960s, have therefore been preferred to national data by researchers. In this type of tests, which eliminate by construction the impact of national regulations, CIP has generally been found to hold between any pair of eurocurrency deposit rates. Such tests also constitute the "purest" version of NIP, in the sense that the interest rates compared relate to (almost) identical financial instruments in all characteristics except their currency of denomination, and also in the sense that the forward rate provides an anchor for expectations.

As to UIP, it is closely related to the so-called expectations (or speculative) theory of the forward rate. According to that theory, the forward premium/discount should be equal to the market's expectation of future depreciation/appreciation, which is exactly the case if CIP holds. So, as Marston (1995)[84] has pointed out, testing UIP is nearly identical to testing the expectations theory of the forward rate, the only difference being due to deviations from CIP attributable to such factors as transactions costs.

As mentioned, NIP is by definition an ex ante concept. Yet, actual returns are not necessarily equal to the expected ones because of forecast errors in predicting future spot exchange rates, quite sizeable in the data. Theoretically, if the foreign-exchange market is efficient, then these forecast errors should be a serially uncorrelated random variable with zero expected mean. As Levich (1985) has emphasized, it is not possible to test market efficiency by itself since the latter notion is always defined relative to some underlying asset-pricing model. In the present context, this pricing model is UIP, so tests of the joint hypothesis of UIP and (foreign-exchange) market efficiency have been attempted in the related literature, instead of UIP tests in isolation, as stressed by Marston (1995).

Simple tests of UIP, such as those in Bilson (1981), Cumby-Obstfeld (1984) and Fama (1984),[85] involve the following regression (very similar to our UIP test equation reported in Table 2):

(18) $D^{A/B}_{t+1} = a + b(I^A_t - I^B_t) + \varepsilon_t$

where the left-hand side is the actual, and not the expected, rate of country A's currency depreciation. Under the joint hypothesis of UIP and market efficiency, which ensures that the disturbance term is uncorrelated with the explanatory variable, the slope (=b) should be equal to one and the intercept (=a) equal to zero.[86] Regressions of the kind have largely provided

[82] E.g. Marston (1995), Ch. 3, p. 43.
[83] Outward and/or inward.
[84] Ch. 4, p. 76.
[85] An extensive survey is Hodrick (1987).

[86] As Marston (1995) points out (Ch. 4, 78-79), perhaps because forward exchange rates were more readily available than Eurocurrency interest rates, many past studies regressed the actual depreciation on the forward premium instead of on the interest differential in (18). Tests of such a form are equivalent to regression (18)

evidence that UIP does not hold. This failure of UIP in the data has mostly been attributed in the literature to two factors: (i) (systematic) forecast errors;[87] and (ii) (time-varying) risk premia, often derived from asset-pricing models with risk-averse investors.[88]

In the preceding sections, however, we did not need pursue such explanations for the failure of UIP, because in our long-run cross-country regressions using average annual data this failure does not occur… Otherwise, the nature of the simple tests we apply is in the tradition of the methodology sketched above, with only a time-indexing difference (averages over a 23-year period are employed instead in our version of (18), corresponding to the main thrust of the present study). Another remark we wish to make in the context of the standard UIP regression (18) is that the intercept (=a) need not be zero since it reflects, besides other things, country-specific factors related to risks arising from national regulations and practices, as discussed in section 2.

In a summary of his own empirical work on UIP, Marston (1995) reports findings that are basically the same as ours. Calculating unconditional and conditional estimates of UIP leads him to conclude that "uncovered interest differentials between the eurodollar and other G-5 eurocurrencies have been quite small *on average*"[89] throughout June 1973-December 1992.

3.3 Real Interest Parity

Real interest parity (RIP) is the proposition that expected real rates of return on investment projects, i.e. real interest rates, are equal(ized) across countries. Another, ex-post (actual or realized) version of RIP is simply to write it as an equality without any expectation operators assuming, for instance, that perfect foresight (as an extreme form of rational expectations) ultimately holds or, as in our case, that RIP is valid on average and in the long run.

It is the expected real rate of return, i.e. the nominal one adjusted for the expected rate of inflation, that matters for both investors and borrowers, the more so in settings of high or variable inflation. Although this seems beyond dispute, it is more difficult to translate it into a straightforward equation for empirical RIP testing than was the case for UIP. As emphasized by Marston (1995, ch.6, 1997), the general reason for these difficulties in applied work is that the real costs of borrowing at home and abroad, from the viewpoint of a firm operating in a given country, are expected costs that should be both "deflated" by the same (the home country's) expected inflation rate and not by the two respective expected inflation rates in the relevant countries.

While Marston (1995) maintains[90] that, unlike for UIP, there is no sound theoretical rationale for RIP, Adler-Lehman (1983) argue that RIP is ensured by financial arbitrage in bonds (at

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itself as long as CIP holds and are known as speculative efficiency tests, because they test the joint hypothesis of (foreign-exchange) market efficiency and the expectations (or speculative) theory of the forward rate, as already noted. Since CIP generally holds for eurocurrency data, UIP and speculative efficiency tests are usually treated as interchangeable in tests based on such series. – As to a=0, see above and below.

[87] For two reasons at least: (i) discrete shifts in regimes that are expected but not realized in the particular sample period, i.e. a situation termed the "peso problem" (in reference to the behavior of the Mexican peso prior to its devaluation in 1976), as in Rogoff (1979), Froot-Thaler (1990), and Engel-Hamilton (1990), among others; and (ii) because market participants may be learning about changes in regimes that have occurred, as in the model of Lewis (1989) where expectations are updated using Bayesian methods.

[88] Such models belong to a line of research starting with Solnik (1974), and going on with Grauer, Litzenberger-Stehle (1976), Sercu (1980) and Stultz (1980), as well as with other, more recent papers, summarized in the surveys by Adler-Dumas (1983), Dumas (1992) and Lewis (1995).

[89] Ch. 4, p. 103. – Emphasis in original.

[90] Ch. 6, p. 152.

least indirectly, e.g. by trade flows). A number of authors nevertheless agree[91] that RIP will hold under two conditions, made evident by decomposing the differential between real interest rates as:

$$(19)\quad E_t(RR^A_t) - E_t(RR^B_t) = \{I^A_t - [I^B_t + E_t(D^{A/B}_{t+1})]\} - \{E_t(CP^A_{t+1}) - [E_t(CP^B_t) + E_t(D^{A/B}_{t+1})]\}$$

with RR denoting the real rate of return (or the real interest rate). An insightful interpretation of the above (expectational) formula is found in Marston (1995), i.e. in his own words:[92]

> 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.[93] The second condition involving ex ante PPP, however, is a condition involving goods markets, not financial markets.

RIP is thus analytically the outcome of the deterministic parts of the two previously discussed parity conditions, in fact, of its key components: UIP and PPP. If these both hold, then the deterministic parts of RIP necessarily hold as well, as they can be thought of as a condition requiring the joint validity of the two preceding ones. This theoretical interdependence, written in the slightly different form of our triple-parity law (8), is one of the central points of our study.

Regressions of a similar kind to that used for UIP tests – see equation (18) above – have generally provided evidence in older studies that RIP does not hold. But as in the case of UIP, Marston (1995) presents his own unconditional and conditional estimates of RIP,[94] which allow him to conclude that "real interest differentials are also quite small on average",[95] for the period June 1973-December 1992 and for the G-5 countries making up his sample.

Apart from the principal reasons for PPP and UIP failures summarized before, complications now also arise, in the case of RIP tests, due to the homogeneity problems when selecting the bond series needed for the regression analysis. It is difficult to collect homogeneous interest rates even for short-term financial instruments: Marston (1995), for example, confines his short-run empirical tests of RIP to money market rates, due to asymmetries related to default risks, maturity, and country premia. This problem is further exacerbated when an attempt is made to select internationally comparable time series of bond analogues. In addition to the stated asymmetries, measuring bond yields creates another problem: Marston (1995) compares yields to maturity and holding period yields for government (and not corporate) bonds, to conclude that, no matter the similarity in results using the two yield definitions, the bond version of RIP[96] is really "difficult to implement empirically".[97] The same problem has arisen in our own empirical tests of ex post long-run UIP and RIP, due to the difficulties in collecting homogeneous government bond time series for our sample of 18 OECD countries.

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[91] Cumby-Obstfeld (1984), Mishkin (1984), Frankel (1986) and Marston (1995, 1997), to mention the most prominent and relevant examples of studies emphasizing the interdependency we call here the triple-parity law.
[92] Ch. 6, pp. 153-154.
[93] "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.
[94] Ch. 6, pp. 163-166.
[95] Ch. 6, p. 177.
[96] Many such studies – e.g. Friend-Tokutsu (1987), Hatsopoulos-Brooks (1986), and McCauley-Zimmer (1989) – calculate some weighted average of the real rate on bank loans and the real rate on bonds.
[97] Ch. 6, p. 167.

4. **Conclusions**

The triple-parity law is relatively simple, integrating as it does three (more or less) familiar parity conditions, and it has been shown to be fully corroborated by the available evidence. Taken individually, none of our empirical results is really new, with the possible exception of the various countries' estimated "R" and "T" factors ; what may also be novel is that the analyses and the empirical results make up an integrated whole. The test period (1976-1998) will perhaps turn out to have been unique: if the euro system does not disintegrate in the future, and all the more so if Britain and the other EU outsiders or candidates should join it, the industrialized world might long remain divided into, and dominated by, three major currency areas only: the dollar, the euro and the yen zones.[98] This would mean that data such as those used here will henceforth make up a much less richer sample.

The law holds in the long run, on average and ex post, and consequently it says nothing about short-term dynamics although it can provide useful benchmarks in this context too.[99] Yet, it confirms that "reversal to the mean" must eventually occur and it may also be useful as a long-term equilibrium condition in (e.g.) error-correction models. Furthermore, it highlights the divergence in some important exchange rate reactions depending on the time horizon adopted: in the short-term, if interest rates increase in a given country, one normally expects its currency to appreciate; according to the triple-parity law, the opposite obtains in the long run. Finally, it confirms a basic choice for open economies under floating exchange rates: in the longer run, a given country can follow a strong currency, low inflation and low nominal interest rate policy; or it can pursue a weak currency, high inflation and high nominal interest rate policy; or anything else along this axis. But it cannot, for example, have both permanently low interest rates to stimulate investment and a permanently weak currency to foster exports. Country-specific factors, i.e. comparative advantages or disadvantages such as political stability, an especially efficient financial sector, a banking secrecy law or being a tax haven, can however mitigate this choice to some extent, at least for a few small countries. With the same proviso, real interest rates tend to be equalized the world over.

Because of its simplicity and its empirically-verified validity, the triple-parity law might – in our opinion should – serve as a natural starting point in lectures and textbooks on international economics (financial and otherwise) when exchange rates, the PPP proposition and the nominal as well as the real interest rate parity conditions are discussed.[100] It may also illustrate a more general 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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[98] With perhaps a few (more or less) independent and (more or less) "fringe" currencies such as the Canadian dollar and the Swiss franc.

[99] A forerunner to this article (Lambelet 1996), circulated as a research paper in January-February 1996, used the estimated PPP condition, with some additional technical adjustments, to calculate a set of "equilibrium" exchange rates. Focusing at one point (pp. 30-2, 36-8) on the US dollar-Swiss franc relationship, it argued – mostly for the benefit of Swiss investors – that the then prevailing current spot exchange rate was so much out of line (i.e. the dollar was so extremely low vis-à-vis the franc) that the U.S. currency was very likely to appreciate considerably before long. The advice offered – of a type generally not found in research papers – thus was: the time is right to buy dollar-denominated financial assets. Those investors who took heed realized an exchange rate profit of nearly 25% within a year (Jan. 22, 1996 - Jan. 21, 1997). This suggests that the triple-parity law can also be used to calculate some benchmarks which may be useful in a short- and medium term context inasmuch as they can help identify extreme disequilibria.

[100] As mentioned in the introduction, this is why we prefer to have a special section on how the theory and the empirical findings relate to the existing literature.

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Appendix

To clarify our exposition, we first propose a notation that is specific to our particular problem here, i.e. transforming into comparable terms the respective parameters from simple OLS regressions run in both directions of causality.

We have sample data (Y_i, X_i), $i = 1,\ldots,18$, which in our case are defined as differentials. Let us denote by *Latin* letters the coefficients of the "direct" regression,[101] i.e. the regression of Y_i on X_i (and, hence, call it the "Latin" regression[102]):

$$Y_i = \hat{a} + \hat{b}X_i + e_i$$

Similarly, let us denote by *Greek* letters the parameters of the "reverse" regression,[103] i.e. the regression of X_i on Y_i (and, hence, call it the "Greek" regression[104]):

$$X_i = \hat{\alpha} + \hat{\beta}Y_i + \varepsilon_i$$

We run both regressions and obtain the OLS estimates for a, b, α and β. It is clear that we cannot compare directly $\hat{b}$ with $\hat{\beta}$ and $\hat{a}$ with $\hat{\alpha}$ because the causality of the two regressions is reversed. Solving the Greek *population* regression for Y_i (as $i \to \infty$, $\varepsilon_i \to 0$ and $e_i \to 0$) results in:

$$Y_i = -\frac{\alpha^0}{\beta^0} + \frac{1}{\beta^0} X_i$$

Hence we have the following one-to-one mapping:

$$a^0 = -\frac{\alpha^0}{\beta^0} \quad \text{and} \quad b^0 = \frac{1}{\beta^0}$$

Since the OLS estimators are unbiased in terms of our *sample* regressions (with $i = 1,\ldots,18$), we have:

$$E(\hat{a}) = a^0 = -\frac{\alpha^0}{\beta^0} = E\left(-\frac{\hat{\alpha}}{\hat{\beta}}\right) \quad \text{and} \quad E(\hat{b}) = b^0 = \frac{1}{\beta^0} = E\left(\frac{1}{\hat{\beta}}\right)$$

In the tables of the main text, we have used the above formulae to make the estimated coefficients from the reverse regressions comparable to those in the direct regressions. As to the somewhat more complicated calculation of their standard errors, t-statistics, probability values and confidence intervals, also reported in our tables, we have applied a Taylor series expansion and the Delta method to approximate the nonlinear relationships between the "Latin" and "Greek" OLS estimators specified above.[105,106]

[101] Terminology in Maddala (1992).
[102] Our terminology.
[103] Terminology in Maddala (1992).
[104] Our terminology.
[105] These are more or less standard techniques in econometrics when dealing with non linearities.
[106] We are indebted to Alberto Holly and Jacques Huguenin for suggesting the general approach and helping with its particular application to our case.

Econometrically, the essence of the problem is to establish the asymptotic distribution of the *transformed* coefficients, which are themselves non linear functions of the OLS estimators of the reverse regression. For our purposes, it is sufficient to derive the relevant expressions only for the first two moments of the distribution of the vector of both estimators, which will enable us to construct confidence intervals for these transformed coefficients.

Since our $\hat{\alpha}$ and $\hat{\beta}$ are the OLS coefficient estimators of the Greek regression, they are both asymptotically normal, so their vector is too, i.e.:

$$\begin{bmatrix}\hat{\alpha}\\ \hat{\beta}\end{bmatrix} \sim N\left(\begin{bmatrix}\alpha^0\\ \beta^0\end{bmatrix}, \Omega^0\right), \text{ with } \Omega^0 = \begin{bmatrix}\omega_{11}^{0\,2} & \omega_{12}^0\\ \omega_{12}^0 & \omega_{22}^{0\,2}\end{bmatrix} \cong \begin{bmatrix}\hat{\omega}_{11}^2 & \hat{\omega}_{12}\\ \hat{\omega}_{12} & \hat{\omega}_{22}^2\end{bmatrix} = \begin{bmatrix}\hat{Var}(\hat{\alpha}) & \hat{Cov}(\hat{\alpha},\hat{\beta})\\ \hat{Cov}(\hat{\alpha},\hat{\beta}) & \hat{Var}(\hat{\beta})\end{bmatrix}$$

In the general case of *any* function of the vector of the OLS sample estimators $\hat{\alpha}$ and $\hat{\beta}$, applying successively the first-order Taylor approximation[107] around the true population parameters α^0 and β^0, and using the Delta method, we obtain:

$$f\begin{pmatrix}\hat{\alpha}\\ \hat{\beta}\end{pmatrix} \cong f\begin{pmatrix}\alpha^0\\ \beta^0\end{pmatrix} + \frac{\partial f\begin{pmatrix}\alpha^0\\ \beta^0\end{pmatrix}}{\partial(\alpha \quad \beta)}\begin{pmatrix}\hat{\alpha}-\alpha^0\\ \hat{\beta}-\beta^0\end{pmatrix} \Leftrightarrow f\begin{pmatrix}\hat{\alpha}\\ \hat{\beta}\end{pmatrix} - f\begin{pmatrix}\alpha^0\\ \beta^0\end{pmatrix} \cong \frac{\partial f\begin{pmatrix}\alpha^0\\ \beta^0\end{pmatrix}}{\partial(\alpha \quad \beta)}\begin{pmatrix}\hat{\alpha}-\alpha^0\\ \hat{\beta}-\beta^0\end{pmatrix}$$

$$\Rightarrow f\begin{pmatrix}\hat{\alpha}\\ \hat{\beta}\end{pmatrix} \sim N\left(f\begin{pmatrix}\alpha^0\\ \beta^0\end{pmatrix}, \frac{\partial f\begin{pmatrix}\alpha^0\\ \beta^0\end{pmatrix}}{\partial(\alpha \quad \beta)}\Omega^0 \frac{\partial f\begin{pmatrix}\alpha^0\\ \beta^0\end{pmatrix}}{\partial\begin{pmatrix}\alpha\\ \beta\end{pmatrix}}\right)$$

In the special case of our functional relationships (defined on the preceding page) we have:

$$f\begin{pmatrix}\hat{\alpha}\\ \hat{\beta}\end{pmatrix} = f\begin{pmatrix}-\frac{\hat{\alpha}}{\hat{\beta}}\\ \frac{1}{\hat{\beta}}\end{pmatrix} \Rightarrow \frac{\partial f\begin{pmatrix}\hat{\alpha}\\ \hat{\beta}\end{pmatrix}}{\partial\begin{pmatrix}\alpha\\ \beta\end{pmatrix}} = \begin{bmatrix}-\frac{1}{\hat{\beta}} & 0\\ \frac{\hat{\alpha}}{\hat{\beta}^2} & -\frac{1}{\hat{\beta}^2}\end{bmatrix} = \frac{1}{\hat{\beta}^2}\begin{bmatrix}-\hat{\beta} & 0\\ \hat{\alpha} & -1\end{bmatrix}$$

And so:

$$\hat{Var}\begin{pmatrix}-\frac{\hat{\alpha}}{\hat{\beta}}\\ \frac{1}{\hat{\beta}}\end{pmatrix} = \frac{1}{\hat{\beta}^4}\begin{bmatrix}-\hat{\beta} & \hat{\alpha}\\ 0 & -1\end{bmatrix}\begin{bmatrix}\hat{\omega}_{11}^2 & \hat{\omega}_{12}\\ \hat{\omega}_{12} & \hat{\omega}_{22}^2\end{bmatrix}\begin{bmatrix}-\hat{\beta} & 0\\ \hat{\alpha} & -1\end{bmatrix} = \frac{1}{\hat{\beta}^4}\begin{bmatrix}\hat{\beta}^2\hat{\omega}_{11}^2 - 2\hat{\alpha}\hat{\beta}\hat{\omega}_{12} + \hat{\alpha}^2\omega_{22}^2 & \hat{\beta}\hat{\omega}_{12} - \hat{\alpha}\hat{\omega}_{22}^2\\ \hat{\beta}\hat{\omega}_{12} - \hat{\alpha}\hat{\omega}_{22}^2 & \hat{\omega}_{22}^2\end{bmatrix}$$

Our *EViews* OLS package gives us directly the values of $\hat{\alpha}$, $\hat{\beta}$, $\hat{\omega}_{11}^2$ and $\hat{\omega}_{22}^2$, so the calculation of the variance of the estimator $\hat{b} = \frac{1}{\hat{\beta}}$, given by $\frac{\hat{\omega}_{22}^2}{\hat{\beta}^4}$ in the above formula, is straightforward. In order to find the variance of the estimator $\hat{a} = -\frac{\hat{\alpha}}{\hat{\beta}}$, given by

[107] We have experimented with second-order Taylor expansion too, but it turns out that this more precise approximation is in our case practically identical to the simpler first-order expansion.

$\frac{\hat{\beta}^2\hat{\omega}_{11}^2 - 2\hat{\alpha}\hat{\beta}\hat{\omega}_{12} + \hat{\alpha}^2\hat{\omega}_2^{22}}{\hat{\beta}^4}$ in the formula above, we also need to compute the sample covariance of the two reverse OLS estimators, i.e. $\hat{\omega}_{12} = \hat{C}ov\left(\hat{\alpha}, \hat{\beta}\right)$. The value of the latter has been obtained using the relationship between the definitions of $\hat{\alpha}$ and $\hat{\beta}$, namely:

$$\hat{C}ov\left(\hat{\alpha}, \hat{\beta}\right) = \hat{C}ov(\overline{X} + \hat{\beta}\overline{Y}, \hat{\beta}) \equiv E\left[\left(\overline{X} + \hat{\beta}\overline{Y}\right) - (\overline{X} + \beta^0\overline{Y})\right]\left(\hat{\beta} - \beta^0\right) = \overline{Y}E\left(\hat{\beta} - \beta^0\right)^2 = \overline{Y}Var\left(\hat{\beta}\right) = \overline{Y}\hat{\omega}_{22}^2$$

The last two values, the sample average for Y_i and the variance of the OLS estimator $\hat{\beta}$, are also directly supplied by *EViews*.

Once the numbers needed for the transformation of the estimated reverse OLS coefficients have been obtained by the algorithm summarized above, calculating the transformed t-statistics and the related probability values and confidence intervals is a standard routine. Our numerical results are reported in the regression tables in the main text.