

The world real interest rate: stochastic index number perspectives

Li Lian Ong, Kenneth W. Clements, H.Y. Izan*

Department of Accounting and Finance, University of Western Australia, Nedlands 6907, Australia

Abstract

In a recent paper, Gagnon and Unferth (1995). [Is there a real world interest rate? *J. Int. Money Finance* 14, 845–855] test the hypothesis of real interest rate equality by regressing real interest rates for all major countries on dummy variables for each country and each period. We show that this model can be interpreted as an application of the stochastic approach to index numbers. This approach leads to estimates of, and standard errors for, the real rate for the world as a whole, which have appealing index-number interpretations and foundations. We also extend Gagnon and Unferth's approach by weighting countries to recognise their differing importance in the world economy, in terms of relative GDP and trade. Weighting does not lead to appreciable changes in the results. However, when the WPI, which gives more prominence to goods entering into international trade, is used as the deflator, our results indicate that real rates are equalized internationally. © 1999 Elsevier Science Ltd. All rights reserved.

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1. Introduction

How do we determine if two regions share the same market? What about two countries which may or may not be neighbours? A useful way of thinking about this problem is to invoke some form of the 'law of one price' whereby regions/countries are part of one market if the prices of relevant commodities/assets are more

*Corresponding author. Tel.: +61 8 93802971, fax: +61 8 93801047, e-mail: hizan@ecel.uwa.edu.au

or less the same once appropriate adjustments have been made. Such adjustments would take into account differing currencies and, possibly, impediments to price equalization which are approximately constant over time. Under this approach, the basic mechanism which equalizes prices is one of arbitrage — agents buy where the price is low and sell where high. A related mechanism which prevents prices within the same market from diverging is expectations. The expectation that pricing anomalies will soon be eliminated can itself rapidly bring about price equalization without any physical movement of goods or assets. Application of these ideas in international finance include the celebrated (but controversial) conditions of purchasing power parity, the link between exchange rates and prices, and the various forms of interest rate parity, the relationship between interest rates and exchange rates.

This paper deals with a particular version of interest rate parity, namely the equalization of real interest rates in different countries. Such equalization has profound effects. If interest rates in real terms are the same internationally, then the cost of borrowing is independent of where that borrowing occurs and in what currency the loan is denominated. Moreover, the equalization of real rates means that the role of monetary policy within any one country is very much circumscribed; in the limit, monetary policy can only affect domestic interest rates via its influences on the world interest rate. A world with highly mobile capital and integrated financial markets would seem to be not inconsistent with these implications of the equalization of real interest rates.

We begin with a brief review of the basic interest rate parity conditions.

1.1. Covered interest rate parity

The theory of covered interest rate parity (CIRP) is ‘a relationship between the premium (or discount) on a forward contract for foreign exchange and the differential in interest rates on securities that are identical in all respects except for the currency of denomination’ (Frenkel and Levich, 1975, p. 337). This theory thus provides a link between domestic and foreign interest rates, and the spot and forward exchange rates. It is synonymous with covered arbitrage of nominal interest rates. When interest rates on comparable securities are used, tests give strong support to covered interest rate parity.¹

1.2. Uncovered interest rate parity

The uncovered interest rate parity (UIRP) condition entails uncovered arbitrage of nominal interest rates so that the interest differential equals the expected change in the spot exchange rate.² If the foreign exchange market is efficient, then

¹See also Frenkel and Levich (1977); Dooley and Isard (1980); Popper (1990) and Mussa and Goldstein (1993).

²A related concept is real uncovered interest rate parity which links real interest differentials and real exchange rates. See Edison and Pauls (1993) for some recent evidence.

the forward rate is an unbiased predictor of the future spot rate ('speculative efficiency', Bilson, 1981). Note that efficiency here means that participants in the foreign exchange market exhibit rational expectations and risk neutrality. CIRP plus speculative efficiency jointly imply UIRP. Uncovered interest parity is contentious and a large body of research has investigated whether or not investors demand a risk premium to compensate for an unhedged exposure to exchange-rate uncertainty. In general, as this research has yielded contradictory results, this issue is still unsettled.³

1.3. Real interest rate equality

The final parity condition relates to the equality of inflation-adjusted interest rates in different countries. Like UIRP, this is also an uncovered relationship between interest rates. As shown in Section 2 below, ex ante relative purchasing power parity and UIRP imply real interest rate equality (RIRE). In what follows, we provide a brief overview of previous research on RIRE.

The bilateral tests of Hodrick (1979) on ex ante real interest rates in the US and other OECD countries indicate that the evidence is not inconsistent with the hypothesis of RIRE. Mishkin (1984), in turn, explores the multi-lateral performance of short-term real interest rates in the Euromarket for seven OECD countries. Unlike Hodrick, he is able to decisively reject the hypothesis of equality of real interest rates in these countries. The joint hypothesis of UIRP and ex ante relative PPP is also strongly rejected by Mishkin. Cumby and Obstfeld (1984) carry out tests of UIRP, ex ante relative PPP, as well as bilateral tests of ex ante RIRE between the US and five other OECD countries. With the exception of the US\$–Deutsche mark exchange rate, they also reject the hypothesis of UIRP. Furthermore, the expected exchange rate changes appear to be poor and biased predictors of relative inflation rates, thereby calling into question relative PPP. Cumby and Obstfeld also reject real interest rate equality in almost all cases.

Mark (1985) incorporates the effects of taxes into his analysis of RIRE but finds little support for the hypothesis of net-of-tax real rate equality. Similarly, Cumby and Mishkin (1986) find that although a significant positive correlation exists between real interest rate movements in the US and seven major industrialized countries, international linkages are less than complete. They also find no evidence that real rates are more closely linked within Europe than with the US.⁴

Dutton (1993) argues that for the purposes of international comparisons, real

³Hansen and Hodrick (1983); Fama (1984); Hodrick and Srivastava (1984); Hsieh (1984) and Cumby (1988) find support for the existence of such premia; however, Frenkel (1982) and Domowitz and Hakkio (1985) find otherwise. Pope and Peel (1991) suggest that the foreign exchange market is possibly inefficient, while Evans and Lewis (1995) show that systematic forecast errors can explain some of the deviations from uncovered interest rate parity, which also points in the direction of inefficiency. For surveys, see Frenkel (1988) and Lewis (1995).

⁴See Mussa and Goldstein (1993) and Piggott (1993) for further discussions on the convergence of real interest rates.

interest rates should be defined in terms of traded goods prices alone; as non-traded goods are poorly arbitrated internationally, the prices of these goods should be excluded from the computation of real returns. Consequently, Dutton uses price indices constructed from the traded goods component of the CPI for the major OECD countries to deflate nominal interest rates; her results provide broad support for the equality hypothesis.

In a pioneering study, Gagnon and Unferth (1995) estimate the common component of real interest rates for all major countries, which they interpret as the world real interest rate. Their data consist of real interest rates of nine OECD countries over the 1977–1993 period. The authors argue that their study differs from previous ones in that they: (1) focus on countries and time period with the most liberal domestic and international capital markets; (2) allow for a constant risk premium across countries; and (3) treat all countries symmetrically, whereas previous studies focus on bilateral parity relationships between the US and selected partner countries.

Our objective in this paper is to illustrate how stochastic index-number methodology (see, e.g. Selvanathan and Prasada Rao, 1994) can be used to measure the world real interest rate and test the hypothesis of RIRE. According to the stochastic approach to measuring inflation, each commodity price change is viewed as containing some information about the true (or underlying) rate of inflation, as well as some commodity-specific noise. This leads naturally to a signal-extraction problem of how to combine the noisy prices in such a way so as to yield an estimate of the rate of inflation which possesses some desirable statistical properties. In addition to a point estimate, this approach also leads to a standard error of inflation which, under certain conditions, increases when there is more variability in relative prices. This attractive result agrees with intuition that in times of large changes in relative prices, the meaning of the overall rate of inflation becomes more ambiguous. The availability of this standard error means that confidence intervals can be constructed for the underlying rate of inflation, and other inference procedures can be employed. It seems to us that this type of information would be highly valued by financial markets. It is for these reasons that advocates of the stochastic approach would like to see government agencies publish standard errors for inflation on a routine basis.

This paper is organized as follows. In Section 2, we demonstrate the conditions under which real interest rates are equalized internationally. Section 3 sets out stochastic index-number methodology. In Section 4, we show how Gagnon and Unferth's (1995) model can be reinterpreted in terms of the stochastic approach. Extensions of the basic idea to deal with differing country sizes, the distinction between traded and non-traded goods, the 'true' real interest rate etc. are contained in Sections 5 and 6. The conclusion follows in Section 7.

2. When are real interest rates equalized internationally?

In this section, we set out sufficient conditions for the equalization of real

interest rates internationally. We begin by defining the relative version of PPP, whereby the change in the exchange rate is equal to the inflation differential:

$$\log(S_t/S_{t-1}) = \log(P_t/P_{t-1}) - \log(P_t^*/P_{t-1}^*),$$

where S is the spot exchange rate, and P and P^* are the price levels at home and abroad. Let this PPP relationship hold ex ante or in an expected sense.⁵ Thus if E_t represents expectations conditional on information available at time t , then for any horizon of k periods,

$$E_t[\log(S_{t+k}/S_t)] = E_t[\log(P_{t+k}/P_t) - \log(P_{t+k}^*/P_t^*)]. \quad (1)$$

In words, the expected depreciation of the exchange rate reflects the expected inflation differential.

Under UIRP, the nominal interest differential between similar k -period bonds denominated in different currencies ($R_{kt} - R_{kt}^*$) must equal the expected change in the exchange rate over the holding period k :

$$R_{kt} - R_{kt}^* = E_t[\log(S_{t+k}/S_t)], \quad (2)$$

where $R_{kt} = \log(1 + n_{kt})$ and $R_{kt}^* = \log(1 + n_{kt}^*)$. The variables n_{kt} and n_{kt}^* represent the k -period nominal interest rates for the domestic country and the foreign country at time t , respectively. Substituting Eq. (2) into Eq. (1), we obtain $R_{kt} - R_{kt}^* = E_t[\log(P_{t+k}/P_t) - \log(P_{t+k}^*/P_t^*)]$, which could also be rearranged as $R_{kt} - E_t[\log(P_{t+k}/P_t)] = R_{kt}^* - E_t[\log(P_{t+k}^*/P_t^*)]$. Accordingly, if we define $r_{kt} \equiv R_{kt} - E_t[\log(P_{t+k}/P_t)]$ and $r_{kt}^* \equiv R_{kt}^* - E_t[\log(P_{t+k}^*/P_t^*)]$ as the domestic and foreign ex ante real rate, we have the result that

$$r_{kt} = r_{kt}^*.$$

Thus, under ex ante relative PPP and UIRP, ex ante real interest rates are equalized internationally.

3. Index numbers as means

Government agencies typically measure inflation by collecting data on individual prices and then combining those in the form of some index-number formula. This approach is deterministic as it involves no uncertainty, the result being that the inflation rate ‘is what it is’ — no more, no less.

An alternative approach is to view inflation as a signal-extraction problem, so that the individual prices contain information on the underlying rate of inflation, but this information is contaminated with ‘noise’. The problem then becomes how

⁵In contrast to older results, recent evidence on PPP is fairly supportive, at least for longer-term horizons (Rogoff, 1996) and especially for hamburgers (Click, 1996; Cumby, 1996; Pakko and Pollard, 1996; Ong, 1997).

to combine the individual prices so that the effects of noise are minimized. The solution involves some form of averaging of the prices. Under certain conditions, these averages coincide with popular index-number formulae used in the conventional approach. Although they are familiar, these measures of inflation have very different foundations from those employed by government agencies. The foundations are statistical, so that uncertainty plays a prominent and explicit role. For example, not only do we obtain a point estimate of inflation, but also a standard error, so that confidence intervals for the underlying rate of inflation can be constructed. This is known as the stochastic approach to index numbers.

The stochastic approach has recently been rehabilitated by Clements and Izan (1987) and Selvanathan and Prasada Rao (1994).⁶ As this will be used later in the paper to measure the world interest rate, we now set out the elements of this approach to illustrate the principles involved. For convenience, the measurement of inflation will be used here.

3.1. The basic idea

Let there be n consumer goods with prices in period t , P_{1t}, \dots, P_{nt} , and $p_{it} = (P_{it}/P_{i,t-1})$ let be the i^{th} price for this period in terms of last period's value. Note that $p_{it} = 1 + g_{it}$, with $g_{it} = (P_{it} - P_{i,t-1})/P_{i,t-1}$ the growth rate in the i^{th} price. Consider the following statistical model for the evolution of p_{it} :

$$p_{it} = \gamma_t + \varepsilon_{it}, \quad (3)$$

where γ_t is an unknown parameter and ε_{it} is an error term, such that

$$E[\varepsilon_{it}] = 0, \quad \text{cov}[\varepsilon_{it}, \varepsilon_{jt}] = \sigma_t^2 \delta_{ij}, \quad (4)$$

where δ_{ij} is the Kronecker delta which takes the value 0 if $i \neq j$ and 1 if $i = j$. Eq. (3) holds for $i = 1, \dots, n$ prices, so the parameter γ_t is interpreted as (one plus) the expected value of the growth rate common to all prices, i.e. $E(p_{it}) = 1 + E(g_{it}) = \gamma_t$. In other words, each price increases at the same rate except for a zero-mean random component. Accordingly, γ_t could also be interpreted as the underlying rate of inflation.

The second member of Eq. (4) for $i \neq j$ states that the random components are uncorrelated over commodities. If we take γ_t to the left-hand side of Eq. (3), we obtain $p_{it} - \gamma_t = \varepsilon_{it}$. The left-hand side of this equation is the growth in the i^{th} price relative to the common growth rate, or the growth in the relative price of good i . Accordingly, the error term ε_{it} is the i^{th} relative price change.

Next, consider the n price changes p_{1t}, \dots, p_{nt} . Under the assumptions in Eq. (4),

⁶These authors provide references to previous literature on the subject. See Diewert (1995) for a critical review of the methodology. Other recent contributions to this area include Crompton (1996) and Selvanathan et al. (1997).

the average of these n price changes provide the best linear unbiased estimator (BLUE) of the rate of inflation, γ_t , in Eq. (3):

$$\hat{\gamma}_t = \frac{1}{n} \sum_{i=1}^n p_{it}.$$

The sampling variance of $\hat{\gamma}_t$ is

$$\text{var}\hat{\gamma}_t = \frac{1}{n} \sigma_t^2, \quad (5)$$

where σ_t^2 can be estimated unbiasedly by

$$\hat{\sigma}_t^2 = \frac{1}{n-1} \sum_{i=1}^n (p_{it} - \hat{\gamma}_t)^2.$$

This expression is a measure of the variability in relative prices in that, when relative prices are unchanged, $p_{it} = \hat{\gamma}_t (i = 1, \dots, n)$ and $\hat{\sigma}_t^2 = 0$; and when there is greater change in relative prices, $\hat{\sigma}_t^2$ increases.

The above perspective on the measurement of inflation illustrates the intuitive result that when there is more dispersion in relative prices, it is more difficult to obtain a precise estimate of the underlying rate of inflation. The sampling variance expression in Eq. (5) increases with relative price variability, and thus provides a natural formalization of intuition.

3.2. The weighted case

The preceding approach can be improved upon by attributing more weight to those commodities which are more important in the consumer's budget. Let $q_{i,t-1}$ be the quantity demanded of good i in period $t-1$, so that $P_{i,t-1}q_{i,t-1}$ is the expenditure on i and $M_{t-1} = \sum_{i=1}^n P_{i,t-1}q_{i,t-1}$ is the total expenditure on all goods. Consequently, $w_{i,t-1} = (P_{i,t-1}q_{i,t-1})/M_{t-1}$ is the share of total expenditure devoted to i , or the budget share of good i .

We replace Eq. (4) with

$$E[\varepsilon_{it}] = 0, \quad \text{cov}[\varepsilon_{it}, \varepsilon_{jt}] = \frac{\lambda_t^2}{w_{i,t-1}} \delta_{ij}, \quad (6)$$

where λ_t^2 is a constant with respect to commodities. Thus, the variance of the relative price of i is $\lambda_t^2/w_{i,t-1}$ and is inversely proportional to $w_{i,t-1}$. This means that the variability of a relative price falls as the commodity becomes more important in the consumer's budget. Clements and Izan (1987) justify this assumption on the basis that as a commodity becomes more important in the overall economy, there is less scope for its relative price to change.

Eq. (6) means that generalized least squares (LS) is appropriate, which we shall

implement in the form of weighted LS. We multiply both sides of Eq. (3) by $\sqrt{w_{i,t-1}}$ to give

$$y_{it} = \gamma_t x_{i,t-1} + u_{it}, \quad (7)$$

where $y_{it} = p_{it}\sqrt{w_{i,t-1}}$, $x_{i,t-1} = \sqrt{w_{i,t-1}}$ and $u_{it} = \varepsilon_{it}\sqrt{w_{i,t-1}}$. It follows from Eq. (6) that

$$\text{cov}[u_{it}, u_{jt}] = w_{i,t-1} \text{cov}[\varepsilon_{it}, \varepsilon_{jt}] = \lambda_t^2 \delta_{ij},$$

so that $\text{var}[u_{it}] = \lambda_t^2$, which is common for all commodities. Under the stated assumptions, we can thus apply LS to Eq. (7) to obtain the BLUE of γ_t :

$$\hat{\gamma}_t = \frac{\sum_{i=1}^n y_{it} x_{i,t-1}}{\sum_{i=1}^n x_{i,t-1}^2} = \sum_{i=1}^n w_{i,t-1} p_{it}, \quad (8)$$

where the second step follows from $\sum_{i=1}^n x_{i,t-1}^2 = \sum_{i=1}^n w_{i,t-1} = 1$. Eq. (8) is a Laspeyres' price index. The variance of $\hat{\gamma}_t$ is given by

$$\text{var}\hat{\gamma}_t = \frac{\lambda_t^2}{\sum_{i=1}^n x_{i,t-1}^2} = \frac{\lambda_t^2}{\sum_{i=1}^n w_{i,t-1}} = \lambda_t^2.$$

An unbiased estimate of λ_t^2 is given by

$$\hat{\lambda}_t^2 = \frac{1}{n-1} \sum_{i=1}^n (y_{it} - \hat{\gamma}_t x_{i,t-1})^2 = \frac{1}{n-1} \sum_{i=1}^n w_{i,t-1} (p_{it} - \hat{\gamma}_t)^2,$$

which we write as

$$\text{var}\hat{\gamma}_t = \frac{\Pi_t}{n-1},$$

where $\Pi_t = \sum_{i=1}^n w_{i,t-1} (p_{it} - \hat{\gamma}_t)^2$ is a weighted variance of relative price changes.⁷ This demonstrates that the variance of the price index $\hat{\gamma}_t$ is proportional to the degree of relative price variability across commodities. Consequently, the same basic result emerges, but in a weighted form.

3.3. Systematic changes in relative prices

In the previous sub-sections, relative price changes are all assumed to be zero on average, i.e. the error term in Eq. (3), which represents the i^{th} relative price

⁷This variance is closely related to the Divisia variance of prices (Theil, 1975/76).

change, has zero expectation. Following Clements and Izan (1987), we now extend our analysis to allow for systematic changes in relative prices.

We redefine p_{it} as the sum of the common trend in all prices, γ_t , a commodity-specific component, β_i , and a zero-mean random component, ζ_{it} :

$$p_{it} = \gamma_t + \beta_i + \zeta_{it}, \quad i = 1, \dots, n; \quad t = 1, \dots, T, \quad (9)$$

where T is the number of time periods. As before, γ_t is the underlying rate of inflation. The parameter $\beta_i = E(p_{it} - \gamma_t)$ is the expected change in the i^{th} relative price. We assume that ζ_{it} is independent over commodities and time and that its variance is inversely proportional to the corresponding budget-share:

$$\text{cov}[\zeta_{it}, \zeta_{jt}] = \frac{\eta_t^2}{w_{i,t-1}} \delta_{ij},$$

where η_t^2 is a constant with respect to commodities.

Since any increase in γ_t for each t by any number k and a corresponding decrease in β_i for each i by the same k would not affect the right-hand side of Eq. (9), we need to impose a normalization in order to identify the model. Clements and Izan (1987) set a budget-share-weighted average of the β_i 's to zero:

$$\sum_{i=1}^n w_{i,t-1} \beta_i = 0. \quad (10)$$

Although there are an infinite number of normalizations Eq. (10) has the attraction of permitting a simple, intuitive interpretation, namely, a budget-share-weighted-average of the relative price changes is zero.

We proceed as before and multiply both sides of Eq. (9) by $\sqrt{w_{i,t-1}}$ to obtain

$$y_{it} = \gamma_t x_{it} + \beta_i x_{it} + \zeta'_{it}, \quad (11)$$

where $\zeta'_{it} = \zeta_{it} \sqrt{w_{i,t-1}}$. The LS estimators of Eq. (11), constrained by Eq. (10), are

$$\hat{\gamma}_t = \sum_{i=1}^n w_{i,t-1} p_{it}, \quad \hat{\beta}_i = \sum_{t=1}^T \phi_t (p_{it} - \hat{\gamma}_t), \quad (12)$$

where $\phi_t = (1/\eta_t^2) / \sum_{s=1}^T (1/\eta_s^2)$. It is apparent that $\hat{\gamma}_t$ is the Laspeyres' index, as before, while the estimator of the systematic component of the change in the relative price, $\hat{\beta}_i$, is a weighted average of the relative price changes ($p_{it} - \hat{\gamma}_t$) over all T periods. The weights ϕ_t are inversely proportional to η_t^2 , which, in turn, is proportional to the error variance in period t . This means that less weight is accorded to those observations with a higher variance.⁸

⁸See Clements and Izan (1987) for details of the derivation of Eq. (12).

The variances of the estimators in Eq. (12) are

$$\text{var}\hat{\gamma}_t = \eta_t^2, \quad \text{var}\hat{\beta}_i = \frac{1}{T} \frac{1}{\sum_{t=1}^T (1/\eta_t^2)} \left(\frac{1}{w_{i,t-1}} - 1 \right). \quad (13)$$

Here, we see that the variance of $\hat{\gamma}_t$ increases with η_t^2 , which, in turn, rises with relative price variability. Therefore, we once again find that the variance of the estimator of overall price index is higher the larger the relative price movements. The variance of $\hat{\beta}_i$ is proportional to the difference between $1/w_{i,t-1}$ and a constant term, which means that this variance increases as $w_{i,t-1}$ falls.⁹

4. Application to interest rates

Recall that in Eq. (9) we defined the price relative p_{it} as the sum of the common trend in all prices, γ_t , a commodity-specific component, β_i , and a zero-mean random component, ζ_{it} . Consider a similar model for real interest rates whereby each country's rate is the sum of the common component in all countries' real interest rates, a country-specific component and a zero-mean random component. Accordingly, we replace the commodity subscript i in Eq. (9) with a country subscript c ($c = 1, \dots, C$ countries) and the price relative p with the real interest rate r :

$$r_{ct} = \gamma_t + \beta_c + \zeta_{ct}. \quad (14)$$

If we reparameterize β_c as $\mu + \alpha_c$, Eq. (14) then becomes

$$r_{ct} = \gamma_t + \mu + \alpha_c + \zeta_{ct}. \quad (15)$$

As before, the parameters of Eq. (15) are not identified. Consider the following normalizations:

$$\sum_{c=1}^C \alpha_c = 0, \quad \sum_{t=1}^T \gamma_t = 0. \quad (16)$$

This is an alternative to the normalization in Eq. (10). The interpretation of Eq. (15) is as follows. Since r represents the change in a bond price, it is completely analogous to the price relative p , which is defined as $1 + g$, with g the rate of growth of the nominal price in question. The parameters $\mu + \gamma_t$ are the common components in all interest rates in the world, which is analogous to the overall rate of inflation in the case of prices. The parameter α_c is the country effect, which, as countries now play the role of commodities, is analogous to the relative price

⁹See Clements and Izan (1987) for details of the derivation of Eq. (13).

change before. This country effect tells us whether the interest rate in a particular country deviates from the world rate.

To proceed further, consider a variable x_{ct} ($c = 1, \dots, C$; $t = 1, \dots, T$) and let $x_{.t} = (1/C)\sum_{c=1}^C x_{ct}$ be the average over the c subscript and $x_{..} = (1/T)\sum_{t=1}^T x_{.t}$ be the average of x_{ct} over both subscripts. Thus, if we average Eq. (15) over $c = 1, \dots, C$, we obtain $r_{.t} = \gamma_t + \mu + \zeta_{.t}$, which follows from the first member of normalization Eq. (16). Next, we average the above over $t = 1, \dots, T$, so that $r_{..} = \mu$, which follows from the second member of Eq. (16) and the assumption $E(\zeta_{it}) = 0$. This means that μ is interpreted as the average real interest rate across all countries and time periods. In this sense, μ is the ‘world’ interest rate.

Gagnon and Unferth (1995) implement Eq. (15) as a dummy variable regression. The parameter γ_t is estimated as the coefficient of a dummy variable that takes the value ‘1’ in period t and ‘0’ in other periods, the parameter μ is the regression constant and α_c is estimated as the coefficient of a dummy variable that takes the value ‘1’ for country c and ‘0’ for other countries.¹⁰ Clearly, Gagnon and Unferth’s estimates can be re-interpreted in terms of stochastic index numbers. Gagnon and Unferth’s data consist of annual real interest rates for Belgium, Britain, Canada, Denmark, Germany, Holland, Japan, Switzerland and the US, for the period September 1977–December 1993.¹¹ Thus, $C = 9$ countries and $T = 16$ periods when annual data are used. The authors define the real interest rate r as the nominal interest rate observed in the last month of period $t - 1$ minus the percentage change in the consumer price index from $t - 1$ to t ; accordingly, r is an ex post real interest rate. Their results for annual data are given in the left-hand panel of Table 1, while Fig. 1 shows the estimated country effects, α_c , and the world real interest rate in year t , $\mu + \gamma_t$, together with its 95% confidence interval.¹² The right-hand panel of Table 1 will be discussed later. We see from the left-hand panel that the country effects for all countries except Britain and Canada are significant. This suggests that there are significant departures of domestic interest rates from the world rate. Thus, these results imply a rejection of the hypothesis of strict RIRE. Panel (II) of Fig. 1 indicates that the real rate for the world varies substantially over time and the width of the 95% confidence interval averages 1.3%.

Gagnon and Unferth (1995) estimate their model by multivariate LS, with the covariance matrix of the residuals restricted to be scalar. This procedure yields the same coefficient estimates as ordinary least squares (OLS); however, the standard

¹⁰Another perspective of Gagnon and Unferth’s approach is that of a panel regression, whereby data are pooled over countries and time. The panel regression approach has two advantages over pure time-series regressions. It reduces the amount of time-series data required, but still provides sufficient data for powerful tests; moreover, it increases the volatility in the data by exploiting cross-sectional variation (see Hakkio, 1984 and Frenkel and Rose, 1995).

¹¹These countries are identified by the OECD as having relatively open and unrestricted financial markets (see OECD, 1990).

¹²The parameter μ is the overall average real rate and γ_t is the deviation at time t from this average. The standard errors underlying the confidence interval are calculated as the square roots of $(\text{var } \mu + \text{var } \gamma_t)$ as $\text{cov}(\mu, \gamma_t) = 0$.

errors are computed differently.¹³ As there is no compelling reason to employ multivariate LS, we re-estimate by OLS using the same data for comparison purposes. We write Eq. (15) as a dummy variable regression:

$$r_{ct} = \mu + \sum_{s=1}^T \gamma_s d_{cts} + \sum_{d=1}^C \alpha_d z_{ctd} + \zeta_{ct},$$

where $d_{cts} = 1$ if $t = s$, 0 otherwise; and $z_{ctd} = 1$ if $c = d$, 0 otherwise. Using constraint Eq. (16) in the form $\alpha_c = -\sum_{c=1}^{C-1} \alpha_c$, $\gamma_T = -\sum_{t=1}^{T-1} \gamma_t$ in the above, we obtain

$$r_{ct} = \mu + \sum_{s=1}^{T-1} \gamma_s d_{cts} - \left(\sum_{s=1}^{T-1} \gamma_s \right) d_{ctT} + \sum_{d=1}^{C-1} \alpha_d z_{ctd} - \left(\sum_{d=1}^{C-1} \alpha_d \right) z_{ctC} + \zeta_{ct},$$

or

$$r_{ct} = \mu + \sum_{s=1}^{T-1} \gamma_s d'_{cts} + \sum_{d=1}^{C-1} \alpha_d z'_{ctd} + \zeta_{ct}, \quad (17)$$

where $d'_{cts} = d_{cts} - d_{ctT}$; and $z'_{ctd} = z_{ctd} - z_{ctC}$.

We use Gagnon and Unferth's data to estimate Eq. (17) by OLS and the results are contained in the right-hand panel of Table 1. Note that the estimates of the coefficients are virtually identical (as expected), while the standard errors increase so that the country and time effects are less significant. Now we find that the country effects for Holland and the US, in addition to Britain and Canada, are insignificant. However, Gagnon and Unferth's rejection of the strict hypothesis of RIRE still holds, as the α_c 's for Belgium, Denmark, Japan and Switzerland are significant.

5. A weighted extension

The significance of the country effects in Gagnon and Unferth's results imply a rejection of the hypothesis of a common real interest rate for the world. It could, however, be that their result is an artefact of their Eq. (17), whereby equal importance is assigned to each country's influence on the world interest rate. In applying unweighted LS to determine the world real interest rate, the same weight is given to interest rates in the US as, say, those of Denmark. Intuitively, we would expect the US, which on average makes up 50% of total GDP and contributes 30% of the total trade of our sample countries, to exert more influence on world interest rates than Denmark, which makes up 1% of GDP and 2% of total trade. Therefore, in this section we explore an alternative approach whereby countries

¹³We are grateful to Joseph Gagnon for clarifying these issues, as well as supplying the interest rate and inflation data used in this paper.

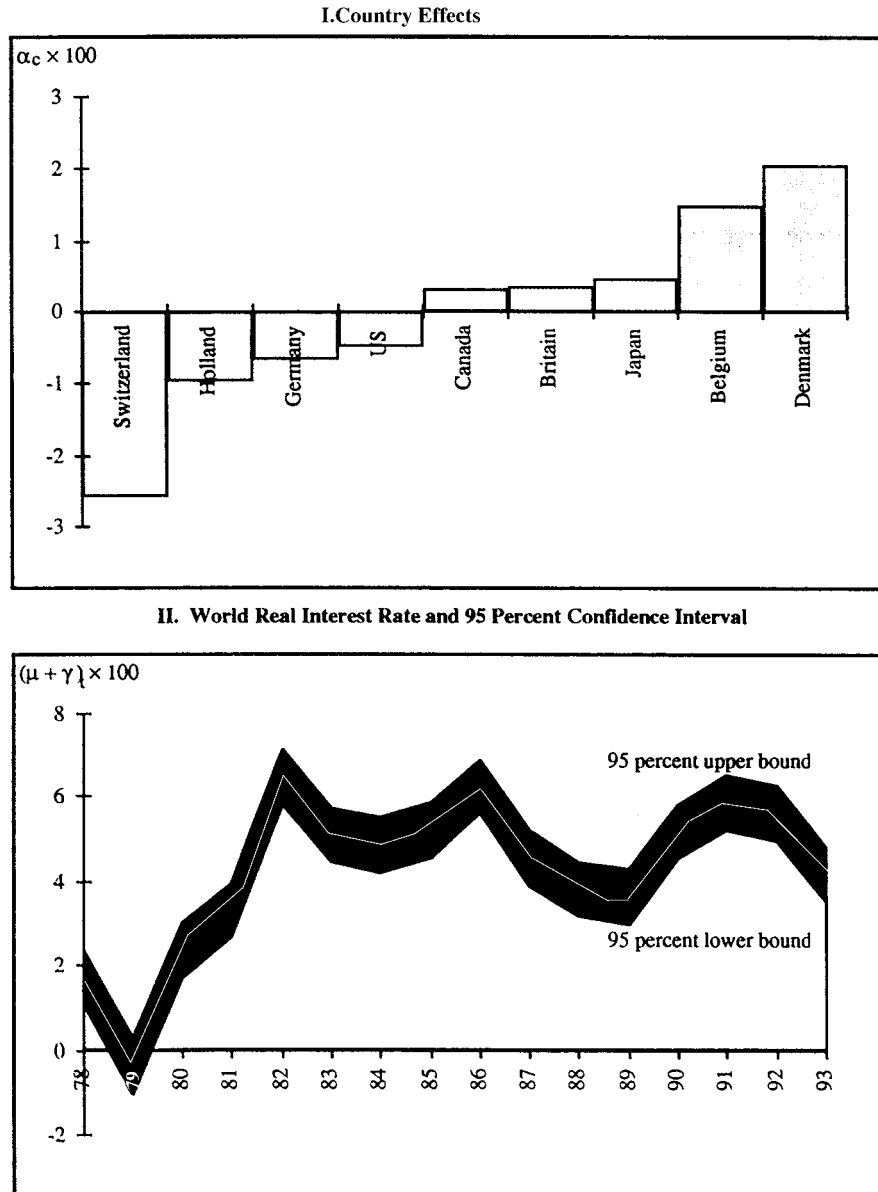


Fig. 1. Gagon and Unferth's estimates of the world interest rate model: annual data, 1978–1993 (I) Country Effects (II) World real interest rate and the 95% confidence interval.

exert differing influences on the world real interest rate, depending on their overall economic influence and integration with the rest of the world. This leads to a weighted world interest rate model which can be regarded as an application of (weighted) stochastic index-number methodology, as discussed above.

Table 1
Estimates of the world interest rate model: annual data, 1978–1993 (S.E. within parentheses)

Coefficient	Gagnon and Unferth's results			OLS results		
	Estimate	<i>t</i> -statistic	% level of significance	Estimate	<i>t</i> -statistic	% level of significance
Intercept, μ	4.24 (0.08)	53.00	0.1	4.23 (0.14)	30.97	0.1
Country effects, α_c						
Belgium	1.50 (0.24)	6.25	0.1	1.50 (0.39)	3.90	0.1
Britain	0.37 (0.24)	1.54	–	0.38 (0.39)	0.98	–
Canada	0.32 (0.24)	1.33	–	0.33 (0.39)	0.85	–
Denmark	2.07 (0.24)	8.63	0.1	2.08 (0.39)	5.38	0.1
Germany	–0.67 (0.24)	–2.79	1	–0.67 (0.39)	–1.73	10
Holland	0.47 (0.24)	1.96	10	0.48 (0.39)	1.24	–
Japan	–0.98 (0.24)	–4.08	0.1	–0.98 (0.39)	–2.54	5
Switzerland	–2.58 (0.24)	–10.75	0.1	–2.64 (0.39)	–6.83	0.1
US	–0.49 (0.24)	–2.04	5	–0.49 (0.39)	–1.26	–
Time effects, γ_t						
1978	–2.50 (0.32)	–7.81	0.1	–2.49 (0.53)	–4.71	0.1
1979	–4.60 (0.32)	–14.38	0.1	–4.60 (0.53)	–8.69	0.1
1980	–1.86 (0.32)	–5.81	0.1	–1.85 (0.53)	–3.50	0.1
1981	–0.93 (0.32)	–2.91	1	–0.93 (0.53)	–1.75	10
1982	2.22 (0.32)	6.94	0.1	2.23 (0.53)	4.21	0.1
1983	0.84 (0.32)	2.63	1	0.85 (0.53)	1.60	–
1984	0.60 (0.32)	1.88	10	0.61 (0.53)	1.16	–
1985	0.96 (0.32)	3.00	1	0.97 (0.53)	1.83	10
1986	1.99 (0.32)	6.22	0.1	1.88 (0.53)	3.56	0.1
1987	0.31 (0.32)	0.97	–	0.32 (0.53)	0.61	–
1988	–0.42 (0.32)	–1.31	–	–0.41 (0.53)	–0.78	–
1989	–0.62 (0.32)	–1.94	10	–0.62 (0.53)	–1.17	–
1990	0.94 (0.32)	2.94	–	0.95 (0.53)	1.79	10
1991	1.63 (0.32)	5.09	0.1	1.64 (0.53)	3.10	1
1992	1.37 (0.32)	4.28	0.1	1.37 (0.53)	2.60	1
1993	–0.08 (0.32)	–0.25	–	0.08 (0.53)	0.15	–

Notes: 1. The coefficient estimates and standard errors are to be divided by 100. 2. The level of significance in the last column shows the probability of incorrectly rejecting the null hypothesis that the individual country or time effects are not significantly different from zero. Where '–' is indicated, the null hypothesis cannot be rejected at any level less than 10%.

To estimate Eq. (17) by weighted LS, it is appropriate to weight the observations for country c by the square root of some measure of the relative size of that country. The idea is that there is likely to be more variability in the interest rates in smaller countries, just as there is likely to be more variability in the relative price of less important commodities, as discussed in Clements and Izan (1987). Of course, giving less weight to smaller countries in this context can be justified directly without any appeal to the variability of interest rates. Let $w_{c,t-1}$ be a

weight for country c (the definition of this weight will be specified subsequently). We then multiply both sides of Eq. (17) by $\sqrt{w_{c,t-1}}$ to yield

$$y_{ct} = \gamma_t x_{ct} + \mu x_{ct} + \alpha_c x_{ct} + \zeta'_{ct}, \quad (18)$$

where $y_{ct} = r_{ct} \sqrt{w_{c,t-1}}$, $x_{ct} = \sqrt{w_{c,t-1}}$ and $\zeta'_{ct} = \zeta_{ct} \sqrt{w_{c,t-1}}$. The corresponding normalizations are

$$\sum_{c=1}^C \alpha_c x_{ct} = 0, \quad \sum_{t=1}^T \gamma_t x_{ct} = 0. \quad (19)$$

The LS estimators of Eq. (18), constrained by Eq. (19), are the analog of Eq. (12). Similarly, their sampling variances are the analog of Eq. (13) so that the variance of the estimated time component $\hat{\gamma}_t$ increases with real interest rate volatility, while the variance of the country component α_c increases as the country 'size' falls.

We employ two different sets of weights: (1) the share of each country's total trade (exports plus imports) in total world trade; and (2) the share in world GDP.¹⁴ The trade weighting scheme is used as it measures the degree of integration of the country in question with the world economy, while the GDP weights represent economic influence in terms of the contribution to world income.¹⁵ Our sample countries are Britain, Canada, Denmark, Germany, Japan, Holland and the US.¹⁶ As Belgium and Switzerland are now omitted, we re-estimate the unweighted Eq. (15) for the seven OECD countries. A comparison of Table 2 with the right-hand panel of Table 1 reveals that the country effects are again significant for Denmark, Germany and Japan, with Britain, Canada, Holland and the US remaining insignificant.

We now apply the weighted Eq. (18) to the data for the seven countries. The results for both GDP and trade weights are presented in Table 3. The country effects are now significant for Germany, Japan and the US, where GDP weights are used, and significant for Denmark, Germany, Japan and the US when trade weights are applied. It is also interesting to note that the time effects are mostly insignificant for the latter half of the sample for both sets of weights; this same pattern occurred previously. The estimates of the average real rate for the world

¹⁴GDP and trade data are obtained from the *International Financial Statistics* of the International Monetary Fund.

¹⁵The two sets of weights reveal some interesting patterns and contrasts. Not surprisingly, we find that the three major economies of US, Germany and Japan have the highest GDP and trade weights. However, while Japan is second richest in terms of GDP, its trade is less than that of Germany. Also of note is that there is more cross-country dispersion in the GDP shares than in the trade shares. The GDP share for the US increased during the first half of the 1980s, but has since fallen, while the opposite is true for Germany. Japan's share of world GDP has increased since the mid-1980s. Similarly, the US share of total world trade increased during the 1980s, while German trade fell. Japan's trade share increased in the early 1980s, but has remained constant since.

¹⁶We are unable to include the full set of countries used by Gagnon and Unferth (1995) as the data for Belgium and Switzerland are incomplete.

Table 2
 Estimates of the world interest rate model for seven OECD countries: annual data, 1978–1993
 (standard errors in parentheses)

Coefficient	Estimate	<i>t</i> -statistic	Percentage level of significance
Intercept, μ	4.39 (0.17)	26.44	0.1
Country effects, α_c			
Britain	0.22 (0.41)	0.53	–
Canada	0.17 (0.41)	0.41	–
Denmark	1.92 (0.41)	4.71	0.1
Germany	–0.83 (0.41)	–2.04	5
Holland	0.32 (0.41)	0.77	–
Japan	–1.14 (0.41)	–2.80	1
US	–0.65 (0.41)	–1.59	–
Time effects, γ_t			
1978	–3.06 (0.64)	–4.76	0.1
1979	–4.78 (0.64)	–7.43	0.1
1980	–2.27 (0.64)	–3.53	0.1
1981	–0.78 (0.64)	–1.21	–
1982	2.50 (0.64)	3.89	0.1
1983	1.14 (0.64)	1.78	10
1984	0.82 (0.64)	1.28	–
1985	1.11 (0.64)	1.72	10
1986	1.87 (0.64)	2.91	1
1987	0.31 (0.64)	0.49	–
1988	–0.45 (0.64)	–0.70	–
1989	–0.24 (0.64)	–0.37	–
1990	0.96 (0.64)	1.49	–
1991	1.69 (0.64)	2.63	1
1992	1.18 (0.64)	1.83	10
1993	0.00 (0.64)	0.00	–

Note: 1. The coefficient estimates and standard errors are to be divided by 100. 2. The level of significance in the last column shows the probability of incorrectly rejecting the null hypothesis that the individual country or time effects are not significantly different from zero. Where ‘–’ is indicated, the null hypothesis cannot be rejected at any level less than 10%.

are quite stable at a little over 4% per annum, whether or not weighting is used. A comparison of the Table 3 results with those in Table 2 reveals that there is only one major effect of employing country weights — the country effect for the US becomes significant. However, the conclusion from this section must be that weighting countries does not seem to appreciably affect the tests of RIRE.

6. Further extensions

So far, we have tested for RIRE using annual data. We now re-estimate Eq. (18)

using quarterly data to test for the short-term validity of RIRE. In doing so, we find that country effects are similar to those using annual data, i.e. α_c , for Germany, Japan and the US are significant when GDP weights are used, while those for Denmark, Germany, Japan and the US are significant when trade

Table 3
Estimates of the weighted world interest rate model: annual data, 1978–1993 (standard errors in parentheses)

Coefficient	Estimate	<i>t</i> -statistic	Percentage level of significance
Weighted by GDP			
Intercept, μ	0.99 (0.07)	15.26	0.1
Country effects, α_c			
Britain	0.93 (0.52)	1.79	10
Canada	-0.06 (0.64)	-0.09	-
Denmark	-3.50 (1.08)	-3.23	1
Germany	0.75 (0.30)	2.45	5
Holland	-1.65 (0.84)	-1.96	10
Japan	1.15 (0.32)	3.57	0.1
US	2.37 (0.22)	10.81	0.1
Time effects, γ_t			
1978	-4.29 (0.57)	-7.51	0.1
1979	-5.51 (0.57)	-9.65	0.1
1980	-2.81 (0.57)	-4.92	0.1
1981	1.40 (0.57)	2.46	5
1982	4.34 (0.57)	7.60	0.1
1983	1.77 (0.57)	3.09	1
1984	1.90 (0.57)	3.32	0.1
1985	1.79 (0.57)	3.14	1
1986	2.98 (0.57)	5.21	0.1
1987	-0.85 (0.57)	-1.49	-
1988	-0.26 (0.57)	-0.46	-
1989	0.12 (0.57)	0.22	-
1990	-0.59 (0.57)	-1.03	-
1991	1.37 (0.57)	2.39	5
1992	-0.11 (0.57)	-0.19	-
1993	-1.25 (0.57)	-2.18	5
Weighted by trade			
Intercept, μ	1.34 (0.07)	20.08	0.1
Country effects, α_c			
Britain	0.76 (0.44)	1.73	10
Canada	-0.08 (0.51)	-0.16	-
Denmark	-2.88 (0.83)	-3.47	0.1
Germany	0.74 (0.31)	2.37	5
Holland	0.22 (0.50)	0.43	-
Japan	-0.05 (0.38)	-0.12	-
US	1.30 (0.29)	4.41	0.1

Table 3 (Continued)

Coefficient	Estimate	<i>t</i> -statistic	Percentage level of significance
Time effects, γ_t			
1978	–3.75 (0.62)	–6.07	0.1
1979	–5.14 (0.62)	–8.34	0.1
1980	–2.40 (0.62)	–3.89	0.1
1981	0.27 (0.62)	0.43	–
1982	3.46 (0.62)	5.61	0.1
1983	1.19 (0.62)	1.93	10
1984	1.36 (0.62)	2.21	5
1985	1.34 (0.62)	2.17	5
1986	2.65 (0.62)	4.29	0.1
1987	–0.16 (0.62)	–0.26	–
1988	–0.49 (0.62)	–0.80	–
1989	0.06 (0.62)	0.10	–
1990	0.18 (0.62)	0.29	–
1991	1.50 (0.62)	2.43	5
1992	0.64 (0.62)	1.03	–
1993	–0.70 (0.62)	–1.14	–

Note: 1. The coefficient estimates and standard errors are to be divided by 100. 2. The level of significance in the last column shows the probability of incorrectly rejecting the null hypothesis that the individual country or time effects are not significantly different from zero. Where ‘–’ is indicated, the null hypothesis cannot be rejected at any level less than 10%.

weighting is applied. The significance of the time effects are similar to the annual estimates.¹⁷

To determine whether capital market intervention affects interest rate equalization, we extend the sample to include countries that have more restricted capital markets, namely, Austria, France, Italy, Norway and Sweden (OECD, 1990). Surprisingly, we find that all five ‘restricted’ countries generally do not have significant country effects, except for France when trade weights are used.¹⁸ The country effects for the ‘unrestricted’ economies are similar to earlier estimates, but are now more significant.

We also analyze the effects of using the WPI as an alternative to the CPI to derive real interest rates. Dutton (1993) argues that because the prices of non-traded goods are poorly arbitrated, real interest rates should be defined in terms of the prices of traded goods alone for the purposes of international comparisons. Since the WPI is made up of mainly traded goods, we re-estimate using annual real interest rates defined in terms of the WPI for the seven OECD countries with free

¹⁷ These estimates are available on request. Results for the unweighted model with quarterly data show significant country effects for Denmark, Japan and the US, compared to Denmark, Germany and Japan for annual data.

¹⁸ Detailed results are available on request.

capital markets.¹⁹ The results are presented in Table 4. In this table, there are three sets of results: (1) unweighted; (2) weighted by GDP; and (3) weighted by trade. Comparing the unweighted results with those obtained with the CPI (Table 2), we see that using the WPI has the effect of making all the country coefficients insignificant, whereas in Table 2, three of the seven country coefficients are significant. This means that RIRE now holds, contradicting Gagnon and Unferth's (Gagnon and Unferth, 1995) findings with the CPI. Similar results of insignificant country effects emerge in the unweighted WPI case when we use quarterly data for (1) the seven countries; and (2) the 12 countries (the additional five have restricted capital markets).²⁰

Next, we compare the GDP-weighted results, given in the middle panel of Table 4, with those in the left-hand panel of Table 3. Since GDP weights are used in both cases, this comparison involves only the WPI (Table 4) against the CPI (Table 3). The country effects are now insignificant in all cases, while previously, three were significant. We thus conclude that here also the WPI results are more supportive of RIRE.²¹ Finally, regarding the trade weights, we compare the results in the far right-hand panel of Table 4 with those in the right panel of Table 3. As can be seen, the effect of moving from the CPI to the WPI is to reduce the number of significant country coefficients from four to none.²²

The main finding thus far is as follows. When real rates are defined in terms of the WPI, which excludes many goods that do not enter into international trade, none of the country effects are significant. In other words, there is now more support for the hypothesis that real rates are equalized internationally. This result seems to hold on both a quarterly and annual basis.

Consider now a more formal comparison between the CPI and the WPI. Let the 'true' rate of inflation be a weighted average of the growth rates of the CPI and WPI, π^{CPI} and π^{WPI} :

$$\pi = \lambda(\pi^{\text{CPI}}) + (1 - \lambda)(\pi^{\text{WPI}}),$$

where λ is the weight given to the CPI. The real interest rate in terms of the true deflator (or the 'true' real rate) is then $r = i - \pi$, or $r = i - [\lambda(\pi^{\text{CPI}}) + (1 - \lambda)(\pi^{\text{WPI}})]$. Alternatively, we could write it as $r = r^{\text{WPI}} - \lambda(\pi^{\text{CPI}} - \pi^{\text{WPI}})$, where $r^{\text{WPI}} = i - \pi^{\text{WPI}}$ is the real rate defined in terms of the WPI. It then follows that if we use the true real rate, Eq. (15) becomes

$$r_{ct}^{\text{WPI}} = \gamma_t + \mu + \alpha_c + \lambda(\pi_{ct}^{\text{CPI}} - \pi_{ct}^{\text{WPI}}) + \zeta_{ct}. \quad (20)$$

¹⁹The WPI data are also provided by Joseph Gagnon.

²⁰Detailed results are available on request.

²¹We also use GDP weights and the WPI with (1) quarterly data; and (2) the 12 countries. For (1), the findings are similar with those using annual data, with all country effects being insignificant. For (2), again all country effects are insignificant.

²²We also use trade weights and the WPI with (1) quarterly data; and (2) the 12 countries. As with the GDP weights, we find that all country effects are insignificant in both cases.

Table 4
 Estimates of the weighted world interest rate model for seven OECD countries: annual data deflated by WPI, 1978 to 1993 (standard errors in parentheses)

Coefficient	Unweighted			Weighted by GDP			Weighted by trade		
	Estimate	t-statistic	% level of significance	Estimate	t-statistic	% level of significance	Estimate	t-statistic	% level of significance
Intercept, μ	6.13 (0.38)	16.07	0.1	6.06 (0.62)	9.73	0.1	6.12 (0.49)	12.36	0.1
Country effects α_c									
Britain	1.47 (0.93)	1.58	-	1.47 (1.16)	1.28	-	1.50 (1.00)	1.50	-
Canada	-0.45 (0.93)	-0.48	-	-0.43 (1.38)	-0.31	-	-0.49 (1.16)	-0.42	-
Denmark	1.33 (0.93)	1.43	-	1.42 (2.72)	0.52	-	1.42 (2.15)	0.66	-
Germany	-1.10 (0.93)	-1.18	-	-1.04 (0.66)	-1.57	-	-1.05 (0.70)	-1.49	-
Holland	0.43 (0.93)	0.46	-	0.15 (1.85)	0.08	-	0.27 (1.14)	0.23	-
Japan	-0.77 (0.93)	-0.82	-	-0.75 (0.85)	-0.89	-	-0.73 (0.90)	-0.81	-
US	-0.91 (0.93)	-0.98	-	-0.83 (0.73)	-1.13	-	-0.93 (0.75)	-1.24	-
Time effects, γ									
1978	-2.71 (1.48)	-1.84	10	-4.77 (1.23)	-3.86	0.1	-3.89 (1.40)	-2.78	1
1979	-12.43 (1.48)	-8.42	0.1	-11.08 (1.23)	-8.98	0.1	-11.95 (1.40)	-8.56	0.1
1980	-3.80 (1.48)	-2.57	5	-3.76 (1.23)	-3.04	1	-3.33 (1.40)	-2.38	5
1981	-2.93 (1.48)	-1.99	5	0.76 (1.23)	0.62	-	-1.38 (1.40)	-0.99	-
1982	2.59 (1.48)	1.75	10	3.74 (1.23)	3.03	1	3.30 (1.40)	2.37	5
1983	0.23 (1.48)	0.15	-	2.49 (1.23)	2.01	5	0.94 (1.40)	0.68	-
1984	-0.65 (1.48)	-0.44	-	1.85 (1.23)	1.50	-	0.49 (1.40)	0.35	-
1985	4.54 (1.48)	3.08	1	3.78 (1.24)	3.06	1	4.07 (1.40)	2.92	1
1986	8.82 (1.48)	5.98	0.1	6.73 (1.24)	5.45	0.1	8.00 (1.40)	5.73	0.1
1987	-0.28 (1.48)	-0.19	-	-0.64 (1.23)	-0.52	-	-0.44 (1.40)	-0.32	-
1988	-2.46 (1.48)	-1.67	10	-1.37 (1.23)	-1.11	-	-2.24 (1.40)	-1.60	-
1989	-1.25 (1.48)	-0.85	-	-1.70 (1.23)	-1.38	-	-1.37 (1.40)	-0.98	-
1990	2.44 (1.48)	1.65	-	-0.16 (1.23)	-0.13	-	1.36 (1.40)	0.97	-
1991	5.15 (1.48)	3.49	0.1	3.78 (1.23)	3.07	1	4.35 (1.40)	3.12	1
1992	1.69 (1.48)	1.14	-	0.19 (1.23)	0.15	-	1.38 (1.40)	0.99	-
1993	1.06 (1.48)	0.72	-	0.17 (1.23)	0.14	-	0.70 (1.40)	0.50	-

Notes: 1. The coefficient estimates and standard errors are to be divided by 100. 2. The level of significance in the last column shows the probability of incorrectly rejecting the null hypothesis that the individual country or time effects are not significantly different from zero. Where '-' is indicated, the null hypothesis cannot be rejected at any level less than 10%.

This equation shows that the unknown weight λ can be estimated as a coefficient. Eq. (20) also reveals that λ has the additional interpretation as the response of the WPI real rate to relative prices, $\partial r^{\text{WPI}} / \partial(\pi^{\text{CPI}} - \pi^{\text{WPI}})$, which is reminiscent of the Tobin–Mundell effect.²³

Estimates of Eq. (20) using annual data are presented in Table 5. Interestingly, we find that the results are very similar to those in Table 2, where CPI real interest rates are used. The WPI real rate appears highly responsive to relative price changes, with λ insignificantly different from 1. This indicates that the true real rate is very similar to the CPI real rate. Similar results to those using quarterly CPI-deflated interest rates are obtained when we re-estimate equation Eq. (20) with data for: (1) the seven countries; and (2) the 12 countries including those with restricted capital markets. For these cases, the estimated CPI weight λ remains high at 0.92 and 0.93, respectively.²⁴

We thus conclude that the true rate of inflation is largely based on CPI prices, where ‘true’ is interpreted in terms of the appropriate deflator for the real interest rate. It appears that while the WPI real rate is equalized internationally, neither the CPI real rate nor the true real rate is equalized. Another interpretation of our results is that the goods and services that enter into the calculation of the CPI are very different from the mainly tradeable goods that are used in deriving the WPI.

7. Summary and conclusion

The integration of world capital markets would tend to equalize the cost of borrowing funds, in real terms, across countries. Recent research has investigated the extent to which this in fact occurs, by examining departures of domestic interest rates from those prevailing in the world. While earlier research tests whether pairs of countries have the same interest rates, in a pioneering study, Gagnon and Unferth (1995) introduce a methodology which applies to all major countries simultaneously. Their approach is to regress interest rates across countries and over time on dummy variables for each period and each country. Interest rates in a given country are then said to exhibit departures from the world rate if the coefficient of the dummy variable for that country is significant.

Another recent area of research is what is known as ‘stochastic index numbers’ (Clements and Izan, 1987, Selvanathan and Prasada Rao, 1994). When applied to prices, the underlying rate of inflation is treated as an unknown parameter to be estimated from the individual price data. Each individual price contains noisy information on the underlying rate of inflation and the effect of the noise is minimized by employing some type of averaging of the individual prices. The attractiveness of this approach is that it leads to not only a point estimate of

²³See Mundell (1963) and Tobin (1965).

²⁴Detailed results are available on request.

Table 5
 Estimates of the true real interest rate with relative price changes for seven OECD countries: annual data, 1978–1993 (standard errors in parentheses)

Coefficient	Estimate	<i>t</i> -statistic	Percentage level of significance
Intercept, μ	4.33 (0.19)	22.79	0.1
Country effects, α_c			
Britain	0.18 (0.41)	0.43	–
Canada	0.19 (0.41)	0.46	–
Denmark	1.94 (0.41)	4.76	0.1
Germany	–0.82 (0.41)	–2.01	5
Holland	0.31 (0.41)	0.77	–
Japan	–1.15 (0.41)	–2.83	1
US	–0.64 (0.41)	–1.57	–
Time effects, γ_t			
1978	–3.07 (0.65)	–4.76	0.1
1979	–4.52 (0.76)	–5.94	0.1
1980	–2.22 (0.65)	–3.42	0.1
1981	–0.71 (0.66)	–1.08	–
1982	2.50 (0.65)	3.87	0.1
1983	1.18 (0.65)	1.82	10
1984	0.87 (0.65)	1.34	–
1985	0.99 (0.67)	1.48	–
1986	1.64 (0.74)	2.21	5
1987	0.33 (0.65)	0.52	–
1988	–0.39 (0.65)	–0.60	–
1989	–0.20 (0.65)	–0.31	–
1990	0.91 (0.65)	1.39	–
1991	1.57 (0.67)	2.35	5
1992	1.16 (0.65)	1.79	10
1993	–0.02 (0.65)	–0.04	–
Relative Price Changes, λ	1.03 (0.05)	19.56	0.1

Notes: 1. The coefficient estimates and standard errors are to be divided by 100. 2. The level of significance in the last column shows the probability of incorrectly rejecting the null hypothesis that the individual country or time effects are not significantly different from zero. Where ‘–’ is indicated, the null hypothesis cannot be rejected at any level less than 10%.

inflation, but also a standard error. This standard error increases with the degree of relative price variability, which agrees with the intuitive notion that, in some sense, inflation is less well-defined when there is substantial variation in relative prices.

In this paper, we combine these two areas of research. We show that Gagnon and Unferth’s model of real interest rates for the world can be interpreted as an application of the stochastic approach to index numbers. Additionally, we extend their approach by weighting countries to recognize their differing importance in

the world economy, in terms of their shares in income and trade. It turns out, however, that weighting per se does not substantially affect the results.

As non-traded goods are poorly arbitrated, it can be argued that for international comparisons, real interest rates should be expressed in terms of traded goods only. We investigated whether our tests are sensitive to the choice of the price index used to deflate nominal interest rates. We used the Wholesale Price Index, which comprises a high proportion of traded goods, as the deflator. In contrast to earlier results which use the CPI, we are unable to reject the hypothesis that real rates are equalized internationally.

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