Journal of International Money and Finance (1985), **4**, 189–208

Some Evidence on the International Inequality of Real Interest Rates

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This paper undertakes econometric tests of the hypothesis that ex-ante real interest rates are equal across countries with highly integrated capital markets. The issue is of practical importance because the violation of real rate equality is a necessary condition for monetary policy to influence the open economy through the real interest rate channel. Although an empirical literature concerning real rate equality already exists, previous investigators have focused on pre-tax real rates. This paper contributes to the literature by attempting to incorporate the effects of taxation into the analysis.

The purpose of this paper is to examine the hypothesis that in a world characterized by a high degree of capital mobility, ex-ante real interest rates are equal across countries which operate under flexible exchange rates. Aside from being of interest in its own right, the present investigation is motivated by three considerations. First, whether or not real interest rates are equal across countries is an issue closely related to the operation of activist stabilization policy in the open economy since one channel through which monetary policy is thought to influence real economic activity is through the real interest rate. This channel would not be available if real rates are equal across countries since the ability of the authorities to influence their own real rate would be limited to the extent to which they could influence the world rate. The inequality of real rates across countries, however, does not provide sufficient evidence to conclude that policy makers have the power to control their own real rate so this study addresses only one facet of the broader question concerning the scope for activist stabilization policy to operate in the open economy. Second, we wish to examine the appropriateness of certain assumptions frequently employed in theoretical models of open economies when applied to the current international environment. We focus on those parity conditions in goods and asset markets which imply real interest-rate equalization across countries. For example, the cross-country equality of real rates is assumed in the early monetary approach to exchange-rate determination papers of Frenkel (1976) and Bilson

* This paper is adapted from Chapter 6 of my dissertation at the University of Chicago. I would like to thank Michael L. Mussa, Jacob A. Frenkel, John Huizinga, Robert J. Barro, Robert E. Cumby, and an anonymous referee for helpful comments and suggestions.

0261-5606/85/02/0189-19 $03.00

(1978), among others.[1] In other models, such as Dornbusch (1976) and Mussa (1982), sticky goods prices permit real interest rates to diverge across countries. Because liquidity effects are possible in these models, real interest rates are influenced in a systematic way by unanticipated monetary disturbances. Third, although an empirical literature concering real rate equality already exists, previous investigators have focused exclusively on the equality of pre-tax real rates. When economic agents pay taxes on interest income and treat interest payments as deductable expenses, the real cost of credit in each country is the net of tax rate and it becomes relevant to ask whether or not net of tax real rates are equal across countries. This paper makes a contribution to the existing literature by incorporating the effects of taxation in the analysis.

The remainder of the paper is organized as follows. In the next section, a hypothesis of the cross-country equality of real interest rates is stated and some recent evidence is reviewed. A description of the data employed here is given in section II, and a regression based strategy for testing real rate equality is implemented in section III. We extend the procedure to test for net of tax real rates in section IV. The analysis to this point is concerned with monthly real rates. In section V, tests of the equality hypothesis are conducted using quarterly real rates. Finally, some concluding remarks are contained in section VI.

I. A Hypothesis of Real Interest-Rate Equality and the Recent Evidence

At time $t+k$, the k-period ex-post real interest rate is defined to be,

$$\langle 1 \rangle \qquad r_{t+k} \equiv i_{t+k} - DP_{t+k}$$

where r_{t+k} is the ex-post real rate at $t+k$, i_{t+k} is the nominal yield on a k-period bond issued at t and matures at $t+k$, and DP_{t+k} is the rate of change in the general price level from t to $t+k$.

At t, people make predictions as to what the ex-post real rate will be at $t+k$, based on the information set, I_t, available at t. Taking expectations on both sides of $\langle 1 \rangle$ conditioned on I_t, we obtain,

$$\langle 2 \rangle \qquad E(r_{t+k}:I_t) = i_{t+k} - E(DP_{t+k}:I_t)$$

where $E(y:X)$ is the mathematical expection of y conditioned on X. This prediction of the ex-post real rate in $\langle 2 \rangle$ then defines the ex-ante real interest rate at t. Similarly, the foreign ex-ante real rate is,

$$\langle 3 \rangle \qquad E(r^*_{t+k}:I_t) = i^*_{t+k} - E(DP^*_{t+k}:I_t)$$

where stars denote variables of the foreign country.

Domestic and foreign ex-ante real interest rates can be hypothesized to be equal through the use of two parity conditions representing asset and commodity-market equilibrium. These are uncovered interest parity (UIP), sometimes referred to as the Fisher Open Hypothesis, and an ex-ante version of purchasing-power-parity (EAPPP), given in equations $\langle 4 \rangle$ and $\langle 5 \rangle$.

$$\langle 4 \rangle \qquad i_{t+k} - i^*_{t+k} = E(s_{t+k} - s_t:I_t) \quad \text{(UIP)}$$

$$\langle 5 \rangle \qquad E(DP_{t+k} - DP^*_{t+k}:I_t) = E(s_{t+k} - s_t:I_t) \quad \text{(EAPPP)}$$

Here, s_t is the (log) spot-exchange rate at t, expressed as the domestic currency price of a unit of foreign exchange. UIP is the equilibrium condition that expected net returns on domestic and foreign assets of similar risk be equal. EAPPP is a fairly weak equilibrium condition imposed on commodity markets which states that the anticipated differential in domestic and foreign inflation equals the anticipated rate of change of the exchange rate. It is implied by the absolute version of PPP ($P_t = s_t + P_t^*$), and the relative version ($DP_{t+k} = DS_{t+k} + DP_{t+k}^*$). However, these implications do not run in reverse. For example, EAPPP would hold if the real exchange rate followed a random walk, but the other two versions would not.[2] Hence, EAPPP is a weaker condition than either the absolute or relative versions of PPP.

To see how these conditions imply the equality of real rates across countries, we substract ⟨5⟩ from ⟨4⟩ to obtain,[3]

$$\begin{aligned} 0 &= (i_{t+k} - E(DP_{t+k}:I_t)) - (i_{t+k}^* - E(DP_{t+k}^*:I_t)) \\ &= E(r_{t+k} - r_{t+k}^*:I_t) \end{aligned}$$

Ex-ante PPP was derived by Roll (1979) as an equilibrium condition arising from intertemporal and international commodity arbitrage, and subjected to econometric testing. Roll notes that his tests of EAPPP can be interpreted as tests of the cross-country equality of real rates if UIP is taken as a maintained hypothesis, and concludes that his evidence weighs in favor of the equality hypothesis. Darby (1983) provides counterevidence in the form of significant moving average terms for the log change in several real exchange rates.

Notice that equations ⟨2⟩ and ⟨3⟩ imply,

$$\langle 6 \rangle \qquad DP_{t+k} - DP_{t+k}^* = E(r_{t+k}^* - r_{t+k}:I_t) + (i_{t+k} - i_{t+k}^*) + u_{t+k}$$

where $u_{t+k} = (DP_{t+k} - DP_{t+k}^*) - E(DP_{t+k} - DP_{t+k}^*:I_t)$. Tests of the equality hypothesis can be conducted by performing a joint test of a zero intercept and unit slope coefficient in regressions of the inflation differential on the nominal interest rate differential. Hodrick (1980) notes that covered interest arbitrage holds well in the Eurocurrency market and replaces the nominal interest rate differential with the forward premium on foreign exchange.[4] Conducting (0,1) tests in regressions of the inflation differential on the forward premium, he concludes that his results support the equality hypothesis. Cumby and Obstfeld (1981) perform (0,1) tests directly on regressions of ⟨6⟩. Using Eurorates and the same countries as Hodrick, they are able to reject the equality hypothesis. Since these regressions are the same in essence as Hodrick's, these differing results must derive from differing sample periods, or estimation procedures.

Mishkin (1982) employs a latent variable statistical model to test for real rate equality. He uses equations ⟨1⟩ and ⟨2⟩ to obtain the decomposition,

$$\langle 7 \rangle \qquad r_{t+k} = E(r_{t+k}:I_t) - e_{t+k}$$

where $e_{t+k} = DP_{t+k} - E(DP_{t+k}:I_t)$. Since the ex-ante real rate is unobservable, it is replaced with its best linear predictor based on X_t, an observable subset of I_t. Denoting this predictor by $X_t b$, the ex-ante real rate is represented as $E(r_{t+k}:I_t) = X_t b + v_{t+k}$, where v_{t+k} is the linear least squares prediction error. Substitution into ⟨7⟩ yields,

$$\langle 8 \rangle \qquad r_{t+k} = X_t b + (v_{t+k} - e_{t+k})$$

Mishkin employs a constant and a fourth order polynominal in time as X_t. He cannot reject the equality hypothesis in bilateral tests of the US *vis-à-vis* six industrialized countries, but is able to reject at very small marginal significance levels in tests of the joint equality of all seven real rates.

Taken as a whole, the evidence on real interest-rate equalization across countries is mixed, and there appears to be room for further research, to which we now turn.

II. The Data

The countries considered here are the US, Canada, Germany, Italy, the Netherlands, and the UK. As emphasized by Mishkin (1982), Eurocurrency rates (obtained from the *Harris Bank Weekly Review*) were used as nominal interest rates for three reasons. First, they are market clearing. Second, because a given offshore bank will issue deposit liabilities in a number of different currencies, Euroloans share similar risk characteristics. Third, being offshore deposits, they are virtually free from regulation and capital control restrictions. The narrow definition of money is used. For the US, this is M1B from the Board of Governors of the Federal Reserve. Money for the other countries and the CPI data were taken from the IFS tapes. Special aggregates of the US and Canadian CPIs were obtained from the *Survey of Current Business* and Statistics Canada's Consumer Price Index publications. All data are monthly and are seasonally unadjusted, and the sample runs from May 1973 to February 1982.

III. Methodology and Empirical Results for Pre-Tax Real Rates

To motivate the approach undertaken here, we restrict economic agent's forecasts to optimal linear prediction rules, so that conditional expectations coincide with linear least squares projections. Taking the difference between equations ⟨2⟩ and ⟨3⟩, we rearrange to obtain,

$$\langle 9\rangle \qquad (i_{t+k}-DP_{t+k})-(i^*_{t+k}-DP^*_{t+k}) = P(r_{t+k}-r^*_{t+k}:I_t)+v_{t+k}$$

where $v_{t+k}=(DP^*_{t+k}-DP_{t+k})-P(DP^*_{t+k}-DP_{t+k}:I_t)$ and $P(y{:}X)$ is the linear least squares projection of y on X. Projecting both sides of ⟨9⟩ on Z_t, an observable subset of I_t, we have by an iterated projections argument,

$$\langle 10\rangle \qquad (i_{t+k}-DP_{t+k})-(i^*_{t+k}-DP^*_{t+k}) = P(r_{t+k}-r^*_{t+k}:Z_t)+w_{t+k}$$

where $w_{t+k}=(DP^*_{t+k}-DP_{t+k})-P(DP^*_{t+k}-DP_{t+k}:Z_t)$. Note that $E(w_{t+j}\, w_t)=0$ for $j\geq k$, provided that $\{DP_t,DP_{t-1},\ldots,DP^*_t,DP^*_{t-1},\ldots\}\subseteq Z_t$. Current and past values of DP and DP^* are included in all of the regressions reported below. In this section, we shall be concerned with monthly forecast horizons and take $k=1$; hence, the w_t process will be serially uncorrelated. Furthermore, if we assume that (i) $P(DP^*_{t+k}-DP_{t+k}:Z_{t-i})=0$ for some $i\geq 0$, and (ii) $\{r_t-r^*_t\}$ and $\{Z_t\}$ are jointly stationary and ergodic, equation ⟨10⟩ can be consistently estimated by least squares.

By running regressions of ⟨10⟩, we estimate economic agents' forecast of the real rate differential $P(r_{t+1}-r^*_{t+1}:I_t)$ by $P(P(r_{t+1}-r^*_{t+1}:I_t)Z_t)=P(r_{t+1}-r^*_{t+1}{:}Z_t)$. The null hypothesis is that $P(r_{t+1}-r^*_{t+1}:I_t)=0$, which implies that $P(r_{t+1}-r^*_{t+1}{:}Z_t)=0$, so that non-zero coefficients in regressions of ⟨10⟩ constitute evidence against the null.

In forming the observable information set Z_t, we consider current and past real interest differentials and rates of monetary growth and inflation at home and abroad. Current and past differentials are likely to provide information about future differentials, and the inflation rates represent one component of the real differential. Also, inflation rates are required as regressors for the forecast error to be serially uncorrelated. Finally, in the event that liquidity effects are present, the behavior of monetary variables may be important. Two different specifications of Z_t are considered. First, current and past ex-post real differentials are employed; then current and past rates of money growth and inflation are used. All regressions include a linear trend and 12 seasonal dummies denoted by $x'l$, but these factors are *not* used as a basis for rejecting the null hypothesis. Thus a significant time varying real differential is required to reject the null hypothesis rather than a constant differential as in the Hodrick and Cumby and Obstfeld analyses.

In this section and the next, the order of the regressions reported were arrived at by conducting a search procedure which minimizes an estimation criterion function discussed by Geweke and Meese (1981).[5] When the set of regressors include the appropriate lagged inflation rates, we also have an identifying restriction in that the residuals will be serially uncorrelated. In cases where the residuals continued to exhibit significant serial correlation, the order of the distributed lag was increased until the hypothesis of no serial correlation could not be rejected.[6]

We adopt the convention that the US be the domestic economy. All of the tests of real rate equality will be relative to the US rate. While tests based on bilateral comparisons lack the statistical power of a joint test (such as in Mishkin, 1982), they provide more information about which countries in the sample contribute towards rejection. Moreover, if equality can be rejected on the basis of bilateral comparisons, a joint test will almost certainly reject also. Indexing the foreign countries in the sample by $j=1,2,\ldots,5$, estimation results for

$$\langle 11\rangle \qquad r_{t+1}-r^*_{j,t+1} = \sum_{i=0}^{k_j-1} a_{ij}\,(r_{t-i}-r^*_{j,t-1})+x'1+w_{t+1}$$

appear in rows 1–5 of Table 1. In all cases, the equality hypothesis can be rejected at the 5% level, with the exception of Germany (marginal significance level $=0.093$).

Because not all prices in the CPI are sampled at a single point in time, part of the first order autocorrelation of CPI measured inflation rates may be spurious due to a time-averaging problem.[7] If this is considered to be a serious problem, part of the correlation between next period's ex-post real rate differential (or CPI inflation rate) may be spurious. To allow for possible time averaging, exclusion restrictions should not be imposed on the current ex-post differential; *i.e.*, we do not impose $a_{0j}=0$. In this case, equality of real rates can be rejected only for the Canadian and Netherlands differentials (m.s.l.$=0.005$ and 0.005 respectively).

In rows 1–5 of Table 2, we report test results from estimation of

$$\langle 12\rangle \qquad r_{t+1}-r^*_{t+1} = \sum_{i=0}^{k_j-1} b^*_{ij}MG^*_{j,t-i}+\sum_{i=0}^{k_j-1} b_{ij}MG_{t-i}+\sum_{i=0}^{k_j-1} d^*_{ij}DP^*_{j,t-i}+\sum_{i=0}^{k_j-1} d_{ij}DP_{t-i}+x'1+w_{t+1}$$

where $MG_t(MG^*_t)$ is the rate of domestic (foreign) money growth from $t-k$ to t. The equality hypothesis can be rejected at the 10% level for all except the Italian case, but the evidence most damaging comes from Canada, Germany, and the

TABLE 1. Tests of monthly pre-tax real interest rate equality based on current and past real interest rate differentials.

Country	$\hat{a}_0$ (s.e.) (confidence)	$\hat{a}_1$ (s.e.) (confidence)	$F(2,89)$ (confidence)	R^2
Canada	0.140 (0.106) (0.810)	0.304 (0.106) (0.995)*	5.821 (0.996)*	0.27
Germany	0.173 (0.102) (0.907)*	—	—	0.32
Italy	0.303 (0.097) (0.998)*	—	—	0.26
Netherlands	0.082 (0.102) (0.578)	0.290 (0.102) (0.995)*	4.690 (0.988)*	0.35
UK	0.287 (0.100) (0.995)*	—	—	0.40
Canada[1]	0.109 (0.108) (0.687)	0.243 (0.108) (0.973)*	3.473 (0.965)*	0.25
Canada[2]	0.086 (0.108) (0.573)	0.181 (0.108) (0.903)*	1.894 (0.844)	0.19

[1] US CPI less shelter.
[2] US and Canadian CPI less shelter.
* Marginal significance <0.1.

Netherlands, where exclusion restriction on economic variables yield F-statistics with marginal significance levels of 0.023, 0.002, and 0.009 respectively. Monetary factors as well as past rates of inflation appear to have been important in predicting the ex-post differentials for these countries, and equality can still be rejected at the 5% level when the coefficients on current inflation are unconstrained to account for possible time averaging.

The last issue to be dealt with in this section concerns the possibility that infrequent sampling of some items in the CPI may have led to spurious rejections of the equality hypothesis. The potential problem exists in the shelter component. In the US, apartment rentals are sampled on a rotating basis. The rent on a given apartment is sampled every six months, and the rents of those not priced in a given month are assumed to remain constant. Because of this, Fama and Schwert (1979) argue that the current period CPI records price changes which occurred in

TABLE 2. Tests of monthly pre-tax real interest rate equality based on current and past money growth and inflation rates.

Country	k_j	m	Test statistic[1] (confidence) MG^*	MG	DP^*	DP	$F(4k_j,m)$ (confidence)	Current lag of inflation rates free $F(4k_j-2,m)$ (confidence)	R^2
Canada	4	73	3.115 (0.980)*	2.470 (0.948)*	3.129 (0.980)*	2.715 (0.964)*	2.018 (0.977)*	2.138 (0.981)*	0.42
Germany	8	53	1.901 (0.921)*	2.433 (0.975)*	2.748 (0.987)*	2.724 (0.986)*	2.433 (0.998)*	2.273 (0.996)*	0.74
Italy	4	73	0.474 (0.245)	0.209 (0.067)	0.644 (0.368)	3.100 (0.979)*	1.128 (0.653)	1.180 (0.691)	0.37
Netherlands	2	83	3.618 (0.969)*	3.394 (0.962)*	3.956 (0.977)*	0.681 (0.491)	2.764 (0.991)*	3.672 (0.997)*	0.43
UK	1	88	−0.035 (0.028)	−1.280 (0.796)	2.575 (0.988)*	−0.243 (0.192)	2.076 (0.909)*	0.824 (0.558)	0.41
Canada[2]	3	78	2.997 (0.964)*	1.536 (0.788)	2.253 (0.911)*	2.148 (0.890)	2.309 (0.986)*	2.507 (0.989)*	0.40
Canada[3]	2	83	2.834 (0.936)*	2.198 (0.883)	0.995 (0.626)	2.974 (0.943)	2.307 (0.972)*	2.240 (0.953)	0.31

[1] t-statistics for UK, $F(k_j,m)$ for other countries. This is the joint test that coefficients on these variables are zero.
[2] US CPI less shelter.
[3] US and Canadian CPI less shelter.
* Marginal significance level <0.1.

preceding periods, and that this creates spurious correlation in CPI measured inflation rates. The Canadian CPI is priced more faithfully, where monthly sampling of major household budget items including housing and rents is the rule. To examine whether or not misdating of price changes in the shelter component imparts a serious bias, we perform the previous calculations for the US–Canadian case, first by using the rate of change in the US CPI less shelter leaving Canadian and inflation unaltered, and also by omitting the shelter component from both Canadian and US CPI inflation rates.[8] These results appear at the bottom of Tables 1 and 2.

In Table 1, equality can be rejected when shelter is omitted only from the US CPI (m.s.l. = 0.035). In Table 2, the null is rejected at the 5% level whether shelter is omitted only from the US CPI, or from both the US and Canadian CPIs. Moreover, if the coefficients on the current inflation rates (or current ex-post differential) are dropped from the set of exclusion restrictions, equality can still be rejected at small marginal significance levels when accounting for possible time averaging.

It appears, at least for the US–Canadian case, that sampling practices in the shelter component of the CPI have not seriously biased the results toward rejection of the equality hypothesis.[9]

IV. On the Equality of Net of Tax Real Rates

Up until now, we have been concerned with the quality of pre-tax real interest rates across countries, which was implied by UIP and EAPPP. However, when economic agents pay taxes on interest income and treat interest payment as deductable expenses, it is the net of tax real interest rate which is economically meaningful, and the equalization or non-equalization of pre-tax real rates tells us nothing about whether net of tax real rates are equal. In this section, the analysis is modified to consider the behavior of net of tax real interest rates across countries.

Let h be the (marginal) tax rate for the domestic investor/borrower, and h^* the tax rate for the foreigner. The k-period ex-post net of tax real interest rate at time $t+k$ at home (r^a_{t+k}) and abroad (r^a_{t+k}) are,

$$\langle 13 \rangle \qquad r^a_{t+k} = (1-h)\, i_{t+k} - DP_{t+k}$$

$$\langle 14 \rangle \qquad r^{*a}_{t+k} = (1-h^*) i^*_{t+k} - DP^*_{t+k}$$

Proceeding as before, we obtain,

$$\langle 15 \rangle \qquad (i_{t+k} - DP_{t+k}) - (i^*_{t+k} - DP^*_{t+k}) = P(r^a_{t+k} - r^{*a}_{t+k} : X_t) + (hi_{t+k} - h^* i^*_{t+k}) + w_{t+k}$$

Taking $k=1$, we can estimate $\langle 15 \rangle$ by OLS and test the null hypothesis that ex-ante net of tax real interest rates are equal across countries; *i.e.*, $P(r^a_{t+k} - r^{*a}_{t+k} : I_t) = 0$. Thus, a time varying wedge $(hi_{t+k} - h^* i^*_{t+k})$ between real rates is introduced in addition to the deterministic displacement incorporated in $x'l$. Note that these net of tax real rates attempt to measure the real borrowing costs or rate of return for an agent who borrows or lends his national currency. It would seem that these are the real rates which are relevant to consumption/investment decisions and to macroeconomic policy. Hence, this section attempts to test for significant divergences in the true borrowing costs across countries.

The scenario for which arbitrage conditions would deliver equality of net of tax real rates requires that UIP be modified such that an agent face $h(h^*)$ when borrowing or lending in the domestic (foreign) currency and that the expected tax liability on foreign exchange capital gains (losses) be small.[10] For an individual this may be a poor representation of his investment opportunities since taxes are not imposed at source in the Eurocurrency market; however, multinational firms are likely to face just such an environment. To maximize the firm's profits, interest will tend to be charged against highly taxed profits and surplus funds will tend to be invested in low tax locations. For example, a foreign based multinational whose subsidiary, in the home country, requires short-term financing may find it advantageous to borrow abroad and make an interest free loan to the subsidiary when the net-of-tax cost of borrowing is lower abroad.

To test the hypothesis of net-of-tax equality of real rates, we shall first perform the joint test that the coefficients on regressors which make up the Z_t set are zero when h and h^* are constrained to equal corporte tax rates. The corporate tax rate is used since, for practical purposes, this rate can be regarded as a flat tax for all but the smallest firms. The tax rates used were those prevailing in 1977, which were 0.48, 0.56, 0.48, 0.48, 0.52, and 0.48 for Canada, Germany, Italy, the Netherlands, the UK, and the US respectively.[11]

Tax laws are highly complicated and vary across countries. We do not presume to

TABLE 3. Tests of monthly net of tax real interest rate equality.

Country	m	n	$\hat{b}$ (s.e.) (confidence)[1]	$\hat{b}^*$ (s.e.) (confidence)[1]	$\hat{a}_0$ (s.e.) (confidence)	$\hat{a}_1$ (s.e.) (confidence)	Tax rates constrained $F(m,n)$ (confidence)	Tax rates unconstrained $F(m-2,n)$ (confidence)	R^2
Canada	4	87	0.373	0.419	0.097	0.282	2.201	4.383	0.30
			(0.207)	(0.253)	(0.107)	(0.106)	(0.925)*	(0.985)*	
			(0.394)	(0.191)	(0.631)	(0.991)*			
Germany	3	89	0.473	0.567	0.149	—	0.772	—	0.38
			(0.185)	(0.201)	(0.099)		(0.488)		
			(0.030)	(0.028)	(0.866)				
Italy	3	89	0.548	0.635	0.186	—	2.718	—	0.47
			(0.201)	(0.112)	(0.087)		(0.951)*		
			(0.265)	(0.834)	(0.965)*				
Netherlands	4	87	0.611	0.817	0.019	0.210	2.426	2.412	0.45
			(0.199)	(0.217)	(0.096)	(0.097)	(0.946)*	(0.904)*	
			(0.104)	(0.880)	(0.156)	(0.967)*			
UK	3	89	0.150	0.069	0.287	—	3.398	—	0.40
			(0.285)	(0.295)	(0.101)		(0.979)*		
			(0.753)	(0.874)	(0.994)*				

TABLE 3 (*cont.*)

B. Tests based on current and past money growth and inflation rates

Country	k_j	n	$\hat{b}=b^2$ $t(n)$ (confidence)	$\hat{b}^*=b^{*2}$ $t(n)$ (confidence)	Test statistic[3] (confidence) MG^*	MG	DP^*	DP	Tax rates constrained $F(4k_j+2,n)$ (confidence)	Tax rates free $F(4k_j,n)$ (confidence)	Tax rates constrained current inflation free $F(4k_j,n)$ (confidence)	Tax rates and current inflation free $F(4k_j-2,n)$ (confidence)	R^2
Canada	8	51	1.392	1.154	1.096	1.646	2.292	2.022	1.667	1.771	1.578	1.612	0.65
			(0.836)	(0.752)	(0.619)	(0.865)	(0.965)*	(0.938)*	(0.952)*	(0.967)*	(0.929)*	(0.935)*	
Germany	8	51	−0.199	−0.305	1.779	1.727	2.556	2.881	2.222	2.319	1.895	1.815	0.76
			(0.158)	(0.240)	(0.897)	(0.885)	(0.980)*	(0.990)*	(0.995)*	(0.996)*	(0.980)*	(0.970)*	
Italy	4	71	1.547	2.270	0.435	0.699	1.413	1.840	1.320	1.278	1.220	0.819	0.58
			(0.878)	(0.977)*	(0.216)	(0.405)	(0.762)	(0.869)	(0.798)	(0.764)	(0.725)	(0.355)	
Netherlands	5	66	1.144	1.876	0.279	0.935	2.199	0.681	1.518	1.417	1.488	1.263	0.61
			(0.747)	(0.939)*	(0.077)	(0.536)	(0.935)*	(0.361)	(0.901)*	(0.854)	(0.884)	(0.759)	
UK	1	86	−0.942	−0.416	−0.347	−1.137	2.801	0.079	1.905	2.318	0.498	0.734	0.41
			(0.654)	(0.322)	(0.277)	(0.741)	(0.994)*	(0.063)	(0.911)*	(0.937)*	(0.263)	(0.517)	

C. Tests based on current and past inflation rates

Country	k_j	n	$\hat{b}=b$[2] $t(n)$ (confidence)	$\hat{b}^*=b^*$[2] $t(n)$ (confidence)	Test statistic[3] (confidence) DP^*	Test statistic[3] (confidence) DP	Tax rates constrained $F(2k_j+2,n)$ (confidence)	Tax rates free $F(2k_j,n)$ (confidence)	Tax rates constrained current inflation free $F(2k_j,n)$ (confidence)	Tax rates and current inflation free $F(2k_j-2,n)$ (confidence)	R^2
Canada	2	85	1.238 (0.784)	0.765 (0.556)	2.109 (0.872)	7.072 (0.999)*	2.598 (0.977)*	3.887 (0.994)*	3.256 (0.984)*	6.335 (0.997)*	0.35
Germany	7	70	0.716 (0.526)	0.000 (0.000)	2.294 (0.963)*	2.576 (0.980)*	2.138 (0.984)*	2.344 (0.990)*	1.661 (0.916)*	1.487 (0.850)	0.59
Italy	1	88	1.000 (0.683)	2.778 (0.995)*	2.892 (0.997)*	−1.198 (0.769)	3.284 (0.985)*	4.706 (0.989)*	—	—	0.49
Netherlands	2	85	0.708 (0.521)	1.840 (0.941)*	3.216 (0.955)*	0.343 (0.289)	2.143 (0.943)*	1.981 (0.895)	2.943 (0.975)*	3.412 (0.962)*	0.47
UK	1	88	−0.679 (0.503)	−0.498 (0.382)	2.806 (0.995)*	−0.041 (0.033)	2.505 (0.952)*	3.926 (0.977)	—	—	0.40

[1] Confidence level of the test that estimated tax rate equals corporate tax rate.
[2] t-test that estimated tax rate equals corporate tax rate.
[3] t-statistics when $k_j=1$, and $F(k_j,n)$ otherwise. This is the joint test that coefficients on these variables are zero.
* Marginal significance level <0.1.

be able to account for all the effects of taxes by the use of corporate rates. For example, a problem associated with the above approach concerns the double taxation of distributed profits. In the UK, Germany, and Canada, some sort of allowance is made to reduce the double taxation of dividends, but no such allowance is made in the US, Italy, or the Netherlands. What is critical is the assumption underlying the motives of the firm. If managers seek to maximize shareholder wealth, to the extent that dividends are paid out of profits made on interest income, the effective tax rate will exceed the corporate rate. By employing flat corporate tax rates, we assume that managers disregard taxes on shareholders in their decision-making.

Since the corporation tax rate may not accurately reflect the true tax burden, it might be desirable to view $\hat{b}$ and $\hat{b}^*$ as estimates of the effective rate of taxation, and to implement the less restrictive test of zero restrictions on the coefficients on the elements in Z_t, while leaving $\hat{b}$ and $\hat{b}^*$ unconstrained. In what follows, the results of both the constrained and unconstained tests are reported.

Tests of the equality of after tax real rates appear in Table 3. In panel A, current and past real differentials make up the information set. In panel B, we use money growth and inflation rates. In panel C, we employ only current and past inflation rates.

Consider panel A. Estimates of tax rates do not significantly deviate from their hypothesized values. In the joint test with estimated tax rates constrained to corporate rates, equality of net of tax real rates is rejected at the 10% level for all but the German case, and at the 5% level of Italy, the Netherlands, and the UK. When tax rates are unconstrained, the equality hypothesis is rejected at the 5% level in Canada, Italy, and the UK with marginal significance levels 0.015, 0.035, and 0.006 respectively. Tests which account for possible time averaging in the CPIs were not performed here, but notice that zero coefficients at the first lag for Canada and the Netherlands can be rejected with marginal significance levels of 0.009 and 0.033 respectively, so that if time averaging is present, equality can be rejected for these two cases when tax rate estimates are unconstrained.

In panel B, coefficients on the foreign nominal interest rate differ significantly from the corporate tax rate for Italy and the Netherlands. The net of tax equality hypothesis is rejected for all but Italy at the 10% level when coefficients on nominal rates are constrained. When they are unconstrained, one can no longer reject for the Netherlands. Dropping the coefficients on current inflation rates from the set of zero restrictions to account for possible time averaging, we can reject equality only for Canada and Germany at the 10% level, and only for Germany at the 5% level. Money growth variables appear to lose importance once taxes are considered.

When money growth is eliminated (panel C), foreign tax rate estimates continue to differ from foreign corporate tax rates in Italy and the Netherlands. Net of tax equality can be rejected at the 5% level for all countries except the Netherlands whether tax estimates are constained or not, and at the 10% level in the Netherlands when tax estimates are constrained (m.s.l. = 0.057). Omitting coefficients on current inflation from the set of zero restrictions permits rejection of net of tax equality at the 10% level for Canada, Germany and the Netherlands when tax estimates are constrained, and in Canada and the Netherlands (at the 5% level) when tax estimates are unconstrained.

Thus, it appears on the basis of the results reported in this section, the hypothesis that monthly net of tax real Eurorates are equal across countries is not well supported. The evidence is considerably more mixed if the time averaging problem is present.

V. On the Equality of Quarterly Real Interest Rates

In this section, we examine the hypothesis of equal quarterly real interest rates across countries by estimation of equations ⟨10⟩ and ⟨15⟩ when $k=3$. Making full use of the data here causes the forecast error, w_{t+3}, to be serially correlated because the sampling interval is finer than the forecast horizon. As is well known, when the regressors are not strictly (econometrically) exogenous, corrections for residual serial correlation using generalized least squares can yield inconsistent estimates. However, since the forecast error is uncorrelated with elements in the Z_t set, ordinary least squares yield consistent estimates. The assumptions made in section III concerning the joint stationarity and ergodicity of the time series employed and the linearity of agent's prediction rules permit us to exploit results derived by Hansen (1982) to obtain the correct standard errors for least squares estimates.[12]

In Table 4, the results of tests of before and after tax equality employing current and past real differentials as regressors are reported. Tests employing money growth rates and inflation as regressors appear in Table 5.

Consider Table 4, where current and past ex-post real rate differentials form the information set. The joint test that $a_0=a_1=0$ is rejected at the 5% level for Canada and the Netherlands, and at the 10% level for the UK (m.s.l.=0.001, 0.006, and 0.086 respectively). This contrasts somewhat with the results in Table 1 where rejections of equality occurred for all countries in the sample at the 10% level. Moreover, equality cannot be rejected here if time averaging in the CPI is considered important because $\hat{a}_1$ does not differ significantly from zero in any of the regressions.

Considering net of tax equality, estimates of tax rates differ significantly from corporate rates only for the Netherlands. In the joint tests with estimated tax rates constrained to the corporate rates, equality of net of tax real rates is rejected at the 5% level in all but the German case. When tax estimates are unconstrained, rejection at the 5% level still occurs for Canada, Italy, and the Netherlands. When $a_0=0$ is dropped from the set of exclusion restrictions, equality can be rejected only in the Netherlands and Italian cases when tax coefficients are constrained, and only in the Italian case when unconstrained.

In Table 5, for test results on pre-tax equality based on regressions employing money growth and inflation variables, the equality hypothesis is rejected at the 5% level for all countries. When coefficients on current lags of inflation rates are omitted from the set of exclusion restrictions, rejections occur at the 5% level of Germany, the Netherlands, and the UK and at the 10% level for Italy.

Finally, when considering net of tax equality, estimates of the US tax rate differs significantly from the US corporate tax rate in the German, Italian, and Netherlands regressions, and the estimated Netherlands tax rate differs from its corporate rate. In contrast to results on monthly net of tax real rates, money growth rates remain important even when taxes are considered. Whether tax coefficients are constrained or not, net of tax equality is rejected at the 10% level in Canada and for the remaining countries at the 5% level. Eliminating coefficients on current lags of inflation rates from the set of zero restrictions result in rejections at the 5% level in all but the Canadian case whether estimated tax rates are constrained or not.

TABLE 4. Tests of before and after tax equality of quarterly real interest rates based on current and past real interest differentials

Country	$\hat{b}^1$ (s.e.) (confidence)	$\hat{h}^{*1}$ (s.e.) (confidence)	$\hat{a}_0$ (s.e.) (confidence)	$\hat{a}_1$ (s.e.) (confidence)	Pre-tax equality $\chi^2(2)$ (confidence)	Tax rates constrained $\chi^2(4)$ (confidence)	Tax rates free $\chi^2(2)$ (confidence)	$\chi^2(3)$ (confidence)^2	R^2
Canada	—	—	0.403 (0.127) (0.998)*	−0.055 (0.150) (0.286)	13.947 (0.999)*	—	—	—	0.19
Germany	—	—	0.052 (0.182) (0.225)	−0.072 (0.183) (0.306)	0.145 (0.070)	—	—	—	0.41
Italy	—	—	0.144 (0.164) (0.620)	−0.178 (0.131) (0.825)	1.928 (0.619)	—	—	—	0.16
Netherlands	—	—	0.329 (0.131) (0.988)*	0.109 (0.124) (0.621)	10.171 (0.994)*	—	—	—	0.35
UK	—	—	0.206 (0.151) (0.827)	−0.016 (0.061) (0.209)	4.895 (0.914)*	—	—	—	0.26

Canada	0.384 (0.196) (0.376)	0.368 (0.241) (0.358)	0.385 (0.128) (0.997)*	−0.082 (0.150) (0.415)	—	13.565 (0.991)*	13.464 (0.999)*	0.902 (0.175)	0.24
Germany	0.633 (0.172) (0.626)	0.646 (0.218) (0.307)	0.143 (0.147) (0.669)	−0.036 (0.149) (0.191)	—	2.570 (0.368)	1.473 (0.521)	0.953 (0.187)	0.51
Italy	0.877 (0.249) (0.889)	−0.608 (0.148) (0.613)	0.084 (0.125) (0.498)	−0.209 (0.040) (1.000)*	—	34.626 (1.000)*	27.077 (1.000)*	30.122 (1.000)*	0.53
Netherlands	0.800 (0.156) (0.960)*	0.859 (0.157) (0.984)*	0.300 (0.102) (0.997)*	0.149 (0.119) (0.789)	—	18.283 (0.999)*	12.202 (0.998)*	8.150 (0.957)*	0.55
UK	0.540 (0.254) (0.187)	0.230 (0.221) (0.811)	0.236 (0.150) (0.884)	−0.012 (0.072) (0.132)	—	9.677 (0.954)*	3.501 (0.826)	2.877 (0.589)	0.31

[1] Confidence level of test that estimated tax rate equals corporate tax rate.
[2] Joint test that estimated tax rates equal corporate tax rates and $a_1=0$.
* Marginal significance level <0.1.

TABLE 5. Tests of before and after tax equality of quarterly real interest rates based on current and past money growth and inflation rates.

Country	$\hat{b}=b$[1] N(0,1) (confidence)	$\hat{b}^*=b^*$[1] N(0,1) (confidence)	$\chi^2(4)$ (confidence)[2] MG^*	MG	DP^*	DP	Pre-tax equality $\chi^2(16)$ (confidence)	Pre-tax equality current inflation free $\chi^2(14)$ (confidence)	Tax rates constrained $\chi^2(18)$ (confidence)	Tax rates free $\chi^2(16)$ (confidence)	Tax rates constrained current inflation free $\chi^2(16)$ (confidence)	Tax rates and current inflation free $\chi^2(14)$ (confidence)	R^2
Canada	—	—	13.690	4.085	9.322	0.720	34.636	19.092	—	—	—	—	0.32
			(0.992)*	(0.605)	(0.946)	(0.051)	(0.996)*	(0.839)					
Germany	—	—	5.441	43.002	0.400	9.029	110.079	90.487	—	—	—	—	0.58
			(0.755)	(1.000)*	(0.018)	(0.940)*	(1.000)*	(1.000)*					
Italy	—	—	4.259	8.107	10.681	41.649	65.763	23.230	—	—	—	—	0.50
			(0.628)	(0.912)*	(0.970)*	(1.000)*	(1.000)*	(0.943)*					
Netherlands	—	—	11.545	9.430	19.624	10.972	92.217	87.506					0.58
			(0.979)*	(0.949)*	(0.999)*	(0.973)*	(1.000)*	(1.000)*					
UK	—	—	11.169	6.621	2.446	6.227	41.252	34.603					0.43
			(0.975)*	(0.843)	(0.346)	(0.817)	(0.999)*	(0.998)*					
Canada	0.424	0.345	9.733	5.831	6.774	3.737	—	—	26.581	25.102	21.918	20.456	0.37
	(0.329)	(0.270)	(0.955)*	(0.788)	(0.852)	(0.558)			(0.913)*	(0.932)*	(0.854)	(0.884)	
Germany	2.164	1.131	8.010	25.137	3.416	28.050	—	—	121.436	112.399	90.365	44.467	0.67
	(0.970)*	(0.742)	(0.909)*	(1.000)*	(0.509)	(1.000)*			(1.000)*	(1.000)*	(1.000)*	(1.000)*	
Italy	2.305	1.264	2.682	3.520	3.442	5.793	—	—	51.183	30.544	39.734	26.516	0.66
	(0.979)*	(0.794)	(0.388)	(0.525)	(0.513)	(0.785)			(1.000)*	(0.985)*	(0.999)*	(0.978)*	
Netherlands	2.114	2.897	6.097	13.925	29.795	24.581	—	—	125.042	113.278	107.906	91.287	0.74
	(0.965)*	(0.996)*	(0.808)	(0.992)*	(1.000)*	(1.000)*			(1.000)*	(1.000)*	(1.000)*	(1.000)*	
UK	−0.108	−1.280	10.756	5.565	2.997	2.310	—	—	44.781	39.000	38.139	34.295	0.45
	(0.086)	(0.800)	(0.971)*	(0.766)	(0.442)	(0.321)			(1.000)*	(0.999)*	(0.999)*	(0.998)*	

[1] Test that estimated tax rate equals corporate tax rate. [2] Joint test that coefficients on these variables are zero. * Marginal significance level 0.01.

VI. Conclusion

This paper has undertaken econometric tests of the hypothesis that short term ex-ante real interest rates are equal across countries which operate under a regime of floating exchange rates and whose capital markets are intimately linked to one another. Disregarding taxes, this hypothesis can be derived as a consequence of parity conditions frequently employed to represent international asset and commodity market equilibrium. Rejections of pre-tax equality of real interest rates can be thought of as rejections of the joint validity of UIP and EAPPP. Moreover, while previous investigators have focused exclusively on the equality of pre-tax real rates, we have attempted to extend the analysis to test for the equality of net of tax real rates. By and large, the hypothesis of net of tax real rate equality has not been well supported. Finally, we have not been able to rule out the efficacy of monetary policy in influencing the open economy through the real interest rate channel, since a necessary condition for such policy to operate is that real rate equality not hold.

Notes

1. See also Flood and Marion (1982), Cox (1980), and Mundell (1968).
2. Theoretically, it is not clear why the real exchange rate would follow a non-stationary process such as a random walk; nevertheless, this appears to have been an important empirical fact over much of the recent experience with floating exchange rates. For evidence on this point, see Frenkel (1981).
3. We note that equations ⟨4⟩ and ⟨5⟩ are sufficient, but may not be necessary to deliver the equalization of real interest rates, as there may exist other ways to derive this condition. The present discussion is merely intended as one way of developing the null hypothesis of real rate equality.
4. See Frenkel and Levich (1975, 1977) for empirical support of the covered interest arbitrage condition in the Eurocurrency markets.
5. Let m be the number of regressors; m_T^*, the largest model being entertained; $\hat{\sigma}_m^2$, any consistent estimator of the error variance σ^2; and $g(T)$, a non-negative function of sample size, T. Geweke and Meese (1981) show when the true model is assumed to be one of a sequence of nested alternatives, minimization of an estimation criterion function, $EC(m,T)=\hat{\sigma}_m^2+m\ g(T)$ $(m=1,2,\ldots,m_T^*)$, asymptotically leads to the choice of the correct model under some fairly general regularity conditions in the normal regression model. When $g(T)>0$, it acts as a penalty function for increasing the size of the model. The particular form used here was $g(T)=2m\hat{\sigma}^2_{m_T^*}(\ln(T_m)/(T_m-m_T^*))$ where $T_m=t-m$. However, m was taken to denote the order of the regression so that in regressions employing lags of more than one variable, a smaller marginal penalty is ascribed to increasing the size of the model, but does not alter the asymptotic properties of the estimation criteria.
6. Tests for serial correlation in the residuals were made by calculating Box–Ljung (1978) statistics for sample autocorrelations of the residuals (not reported). These statistics are distributed as chi-squared under the hypothesis of white noise residuals.
7. See Working (1960). The problem he addresses is when averaging a sequence whose first difference is serially uncorrelated, a random walk for example, and then taking the difference of these averages, one obtains a series which will behave like an $MA(1)$. As an illustration, let $p_t=\frac{1}{2}(x_t+x_{t-1})$ where $x_t=x_{t-1}+e_t$ with e_t *i.i.d.* with mean 0, variance s_e^2. Then it can be seen that $DP_t\equiv p_t-p_{t-2}=\frac{1}{2}(e_t+2e_{t-1}+e_{t-2})$, and $E(Dp_t\,Dp_{t-2})=\frac{1}{4}s_e^2$. Thus averaging a random series and taking first differences results in a serially correlated series. This may not be as serious a problem as it first appears to be, since the CPI is an average not of a single commodity price, but of different prices. Suppose each month is comprised of two periods, and the CPI is composed of the prices of two goods, x and z, with x sampled at the end of the period and z sampled at mid-month. Now suppose that $x_t=x_{t-1}+e_t$, and $z_t=z_{t-1}+u_t$ with $(e_t,u_t)'$ *i.i.d.* with zero mean and variance matrix $[s_e^2\ s_{ue}/s_{eu}\ s_u^2]$. The CPI at time t would be $p_t=\frac{1}{2}(x_t+z_{t-1})$ where each good is assumed to have a budget share of $\frac{1}{2}$. Now it can be seen that $Dp_t\equiv p_t-p_{t-2}=\frac{1}{2}(e_t+u_t+e_{t-1}+u_{t-1})$, $Dp_{t-2}\equiv p_{t-2}-p_{t-4}=\frac{1}{2}(e_{t-2}+u_{t-2}+e_{t-3}+u_{t-3})$. Now, $E(Dp_t\ Dp_{t-j})=0; j\geq 2$.

8. Since items in the Canadian CPI are by and large sampled monthly, there is no compelling reason for omitting shelter from the Canadian CPI. It is done here only to put the two real rates on an equal footing.
9. It remains an open question as to whether sampling practices in the construction of the CPI in other countries may have biased the results, as detailed documentation and disaggregated CPI data for these countries were not readily available.
10. There is an empirical literature (Levich, 1978 and Meese and Rogoff, 1983) which suggests that the current spot-rate is the best predictor of the future spot-rate in that it has been shown to dominate more sophisticated models in terms of predictive performance. This is in accord with Mussa's (1976) observation that the exchange rate follows approximately a random walk. Thus, it seems reasonable when two countries are experiencing similar rates of inflation, that the *ex-ante* tax liability on foreign exchange capital gains (loses) will be small. This omission is potentially important, however. Levi (1977) has shown that the tax treatment on hedging operations can exert real effects.
11. The corporate tax rate in Italy is actually 38%, but due to applicable municipal taxes, does not accurately reflect the corporate tax burden there. See Stapleton and Burke (1978). In the constrained tests, we assume that these municipal tax rates are 10%, so that the 'effective ' corporate tax rate is 48% in Italy.
12. Consider the linear regression model, $y_{t+k} = x_t' b + u_t$; $t = 1,2, \ldots, T$, where T is sample size and $E(u_t : x_t, x_{t-1}, \ldots, u_{t-j}, \ldots) = 0$ for $j > k \geq 1$. Suppose that y_t and x_t are jointly stationary and ergodic, and let $\hat{u}_t$ and $\hat{b}$ be the least squares residual and coefficient vector respectively. Hansen has shown that $\sqrt{T}\,(\hat{b} - b)$ converges to a normally distributed zero mean random vector with covariance matrix $[E(x_t x_t')]^{-1}\ V[E(x_t x_t')]^{-1}$, where

$$V = \sum_{k=1-j}^{j-1} E(u_t x_t x_{t-k}' u_{t-k}).$$

This asymptotic covariance matrix is consistently estimated by

$$\left[(1/T)\sum_{t=1}^{T} x_t x_t'\right]^{-1} \hat{V}\left[(1/T\sum_{t=1}^{T} x_t x_t')\right]^{-1} \text{ where } \hat{V} = \sum_{k=1-j}^{j-1} (1/T)\cdot \sum_{t=k+1}^{T} \hat{u}_t x_t x_{t-k}' \hat{u}_{t-k}$$

Alternatively, one could employ quarterly data to coincide with the forecast horizon so that the forecast error would be serially uncorrelated. Hansen and Hodrick (1980) suggest that in the present circumstances, making full use of the data is asymptotically more efficient than a sampling procedure which uses quarterly data.

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