UNIVERSITÉ DE SHERBROOKE
Faculté des lettres et sciences humaines
Département d'économique

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GOVERNMENT DEBT, INTEREST RATES AND INTERNATIONAL CAPITAL FLOWS: EVIDENCE FROM COINTEGRATION

by

Pene Kalulumia*

Department of Economics
University of Sherbrooke
Sherbrooke (Qc), Canada J1K 2R1
Tel: (819) 821-7236
Fax: (819) 821-7237
E-mail: k.pene@afdb.org

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ABSTRACT

This paper examines the impact of government debt on interest rates in the United States, Germany, the United Kingdom and Canada. It builds on the general portfolio balance framework which allows for both direct and indirect tests of the link between public debt and interest rates, and uses the Johansen-Juselius multivariate cointegration techniques to perform these tests. Indirect tests in this model consist of investigating the debt impact on interest rates through the effects of debt on the exchange rate and money demand. The evidence indicates that both transitory and permanent changes in the level of government debt cause higher domestic interest rates and money demand and appreciate the exchange rate in all four countries under study.

I. INTRODUCTION

Since the early 1970s many industrialized countries have experienced large shifts in their government debt and deficits while the international capital flows have grown tremendously in world financial markets. These events raise a number of questions, particularly, whether such government debt and deficits affect interest rates and private demand in the presence of international capital flows. The macroeconomic literature has addressed this issue in two interesting and separate frameworks. According to various versions of Keynesian and Neoclassical models, an increase in government debt which makes households wealthier, stimulates both output and employment, and causes higher interest rates. The driving up of interest rates, however, crowds out private investment with deleterious impact on long-term growth.

In the last two decades, this traditional view has been cast into doubt through the reawakening of another theory of public debt's effects, the Ricardian overlapping generation model, suggesting that government debt implies future tax liabilities with a present value equal to the value of the debt. Rational agents recognize this debt/tax equivalence and act as if the debt does not exist, resulting in the debt and deficit having no effect on interest rates nor any of the effects attributed to them by traditional models. Which of those two views is appropriate bears an important implication for macroeconomic policy-making.

Empirical work, intended to quantify the relationship between government debt or deficits and interest rates, has reached disconcertingly different conclusions using US postwar data. First, Barth, Iden and Russek (1985), Zahid (1988), Cebula and Cock (1991, 1994), and Bahmani-Oskooee (1994), among others, concluded that public debt and deficits increase interest rates (long-run rates were mostly used). They tend to support the Keynesian hypothesis. Second, Evans(1985, 1987), Barro (1987), Deravi, Hegji and Morbely (1990), Seater (1993) and Gulley (1994), among others, found no evidence linking government debt or deficits and interest rates (short-run rates were often used). They then tend to support the Ricardian equivalence hypothesis.

At the heart of these two puzzles, noted in previous empirical results, are the statistical techniques employed to assess the interest rate effects of debt and deficits. These techniques can be divided into two categories. The first method, which has been widely utilized, is the interest rate regression approach. It consists in regressing some measure of the nominal interest rate on various independent variables, including measures of government debt. As pointed out by Deravi *et al*. (1990) and Gulley (1994), interest rate regressions lead to weak tests of debt effects on interest rates since any link between these two variables can be clouded either by an inflow of capital which occurs before interest rates respond to government deficit or by the monetary policy that often targets the nominal interest rates. The second approach, proposed in empirical literature, indirectly examines the effect of government debt or deficits on interest rates through the

impact of debt or deficits on any relevant variable which is closely related to the nominal interest rate. This helps to avoid problems caused by interest rate regressions, especially, the low variability in interest rates. Along this line of thought, Evans (1985), Deravi *et al.* (1990) and Gulley (1994) propose an approach which attempts to estimate money demand functions which include a measure of government debt to account for the wealth effect. This procedure indirectly analyzes the impact of government debt on interest rates through the effect of debt on money demand. In the same vein, Seater (1993) suggests a test of the relationship between public debt and exchange rates as a more direct assessment of the debt-interest rate linkage under higher international capital flows that may offset interest rate effects of public debt. Following Evans(1985) and Seater (1993), if the private sector is not Ricardian, both money demand and the exchange rate will be positively related to government debt even in the presence of early capital flows and/or that of a monetary policy that targets interest rates.

Although the indirect methodology is adequate, previous studies which applied it analyzed the effect of deficits on interest rate-related variables, using the Engle and Granger (1987) procedure or single-equation models that produce inefficient estimation and test results [see Johansen and Juselius (1992) for discussion]. Furthermore, tests of theories of government debt's effects developed in earlier research do not correctly distinguish between the behavior of Keynesian, Neoclassical and Ricardian individuals according to the persistence of debt and deficits. When confronted with temporary deficits, Keynesian agents who are myopic and/or liquidity constrained will view the resulting new government debt as net wealth, while both Neoclassical and Ricardian individuals who are rational, will fully discount this new debt. If the deficit is permanent, however, Neoclassical individuals will only partially discount the debt due to their finite lifetimes and hence, debt and deficits will affect real variables as in a Keynesian economy. Therefore, a way to clearly differentiate between the Ricardian hypothesis and the Keynesian-Neoclassical models is to investigate the effects of permanent shifts in government debt on interest rates. The rejection of the temporary debt-interest rate linkage cannot necessarily be an indication of the Ricardian equivalence, since the individuals may also behave in a Neoclassical manner.

The objective of this paper is to examine the effects of government debt on interest rates in order to contribute to the debate on the Ricardian equivalence hypothesis. The paper improves upon previous studies in three important ways. Firstly, it uses a general portfolio balance model which embodies indirect tests of interest rate effects of government debt à la Evans (1985) and Seater(1993), as well as a standard direct test from interest rate equations. Secondly, tests of government debt's effects are preformed under the Johansen (1988) and Johansen and Juselius (1990) multivariate cointegration model, which determines all the possible

cointegrating vectors and allows one to distinctively test for the effects of temporary and permanent changes in series. Transitory effects are obtained from the short-run part of the model, while tests for long-run restrictions such as the Ricardian equivalence, are to be derived from individual cointegrating or long-run relationships, which provide the impact of permanent shifts in model variables. Thirdly, while previous studies focus on the United States (US), the current paper includes other industrial countries comprising Canada, Germany and the United Kingdom (UK). We also use the broad definition of money, M2, instead of M1 [as suggested by Miller (1991), Arize and Shiff (1993), among others].

The paper is organized as follows. It starts with a description of the economic framework, then presents a brief discussion of the Johansen's procedure, followed by the outcome of preliminary unit root and model specification tests. Section 4 is devoted to the tests of identifying restrictions on cointegrating structures to assess the impact of permanent shifts in the level of debt on interest rates. The paper ends with concluding remarks.

II. THE ECONOMIC FRAMEWORK

Consider a general portfolio balance model which builds upon the following cross-country behavioral equations:

$$m_t^d - p_t = a_1 y_t - a_2 i_t, \quad m_t^d = m_t^s = m_t \tag{1.1}$$

$$m_{ft}^d - p_{ft} = a_1 y_{ft} - a_2 i_{ft}, \quad m_{ft}^d = m_{ft}^s = m_{ft} \tag{1.2}$$

$$\bar{s}_t = p_t - p_{ft} \tag{1.3}$$

$$i_t = r_t + \text{Ä}p_{t+1}^e \tag{1.4}$$

$$i_t = r_{ft} + \text{Ä}p_{f,t+1}^e \tag{1.5}$$

$$b_t - s_t - b_{ft} = \text{ä}\,(i_t - i_{ft} - \text{Ä}s_{t+1}^e), \quad \text{ä} > 0 \tag{1.6}$$

$$\text{Ä}s_{t+1}^e = \text{ö}(\bar{s} - s)_t + (\text{Ä}p_{t+1}^e - \text{Ä}p_{f,t+1}^e), \quad \text{ö} > 0 \tag{1.7}$$

where p is the price level, y is the real national income, i is the nominal interest rate, s is the bilateral exchange rate, r is the real interest rate, b is the stock of assets (government bonds), $\text{Ä}p^e$ is the expected inflation rate, and $\text{Ä}s^e$ is the expected change in the exchange rate. Subscripts f and upper bars denote, respectively, foreign variables and long-run values. All variables are expressed in natural logarithms except i, i_f, r, $\text{Ä}p^e$ and $\text{Ä}s^e$

which are in percentages. The first five equations constitute the basic monetary model comprising two typical Cagan demand for money functions with their associated equilibrium conditions for money markets; a purchasing power parity (PPP) relationship which is assumed to hold only in the long-run; and two Fisher conditions for the home and foreign countries, relating nominal to real interest rates and to the expected inflation rate. The portfolio equilibrium equation (1.6) relates the risk premium ($rp = i - i_f - \ddot{A}s^e$) to the relative supply of domestic and foreign assets or bonds. An increase in the relative supply of domestic bonds requires an increased risk premium of these assets (rp>0) to be willingly held in international portfolios. Thus, in order to diversify the foreign exchange risk, investors balance their portfolios between home and foreign countries' bonds in proportions that depend on the relative rate of return or risk premium. Equation (1.7) states that the expected change in the exchange rate ($\ddot{A}s^e$) is governed by a regressive expectation and the expected inflation differential.

Assuming that money supply is exogenously determined by the monetary authorities and real interest rates equalized across countries, the first five equations can be solved to provide the standard bilateral monetary exchange rate expression:

$$\bar{s}_t = m_t - m_{ft} - a_1(y_t - y_{ft}) + a_2(\ddot{A}p^e_{t+1} - \ddot{A}p^e_{f,t+1}) \tag{1.8}$$

Now, assume that long-run equilibrium values are given by their actual current values (e.g. $s_t = \bar{s}_t$) and substitute equation (1.8) into (1.6). Thus, one obtains the following equilibrium equation for nominal money demand

$$m_t - m_{ft} = a_1(y_t - y_{ft}) - a_2(\ddot{A}p^e_{t+1} - \ddot{A}p^e_{f,t+1}) - \ddot{a}(i_t - i_{ft} - \ddot{A}s^e_{f,t+1}) + b_t - b_{ft} \tag{1.9}$$

which is now a function of the risk premium and the relative stock of domestic assets. On the other hand equation (1.7) may be transformed by simultaneously subtracting the interest rate differential from both members and solving for s_t as follows:

$$s_t = \bar{s}_t - 1/\ddot{o}[(i_t - p^e_{t+1}) - (i_{ft} - \ddot{A}p^e_{f,t+1})] + 1/\ddot{o}(i_t - i_{ft} - \ddot{A}s^e_{t+1}) \tag{1.10}$$

According to this equation, the exchange rate deviates from its long-run value by an amount proportional to the real interest rate differential and the risk premium. By substituting respectively equations (1.8) into (1.10), and equation (1.6) for the risk premium into the last term of (1.10), one obtains the following generalized asset market representation of the exchange rate:

$$s_t = m_t - m_{ft} - a_1(y_t - y_{ft}) + (a_2 + 1/\ddot{o})(\ddot{A}p^e_{t+1} - \ddot{A}p^e_{f,t+1}) - 1/\ddot{o}(i_t - i_{ft}) + (1/\ddot{o}\ddot{a})(b_t - b_{ft}) \tag{1.11}$$

The Johansen's procedure will build upon variables included in the two equilibrium equations (1.9) and (1.11). In this multivariate setting, the finding of more than one cointegrating vector indicates additional long-run relationships which need to be identified. For the portfolio model outlined above, the following relationships of primary interest for the test of the Ricardian equivalence hypothesis can be identified (we drop all foreign variables except i_f):

(a) the equilibrium money demand equation

$$m_t = \ddot{a}_1 y_t + \ddot{a}_2 i_t + \ddot{a}_3 i_{ft} + \ddot{a}_4 b_t + \ddot{a}_5 \ddot{A} p^e_{t+1} + \ddot{a}_6 \ddot{A} s^e_{t+1}, \text{ where } \ddot{a}_4 = 1, -\ddot{a}_2 = \ddot{a}_3 = \ddot{a}_6, \ddot{a}_1 > 0 \text{ and } \ddot{a}_2, \ddot{a}_5 < 0 \tag{1.12}$$

(b) the equation for the domestic interest rate reflecting the relative demand for domestic assets (capital flows relationship)

$$i_t = \ddot{o}_1 i_{ft} + \ddot{o}_2 b_t + \ddot{o}_3 y_t + \ddot{o}_4 m_t + \ddot{o}_5 \ddot{A} p^e_{t+1} + \ddot{o}_6 \ddot{A} s^e_{t+1}, \text{ where } \ddot{o}_1 = \ddot{o}_6 = 1, \ddot{o}_5 < 0 \text{ and } \ddot{o}_2, \ddot{o}_3 > 0 \tag{1.13}$$

(b) and, the equation determining the exchange rate

$$e_t = \grave{e}_1 m_t + \grave{e}_2 y_t + \grave{e}_3 i_t + \grave{e}_4 i_{ft} + \grave{e}_5 b_t + \grave{e}_6 \ddot{A} p^e_{t+1}, \text{ where } \grave{e}_1 = 1, \grave{e}_3 = -\grave{e}_4, \grave{e}_3 < 0, \text{ and } \grave{e}_2, \grave{e}_5, \grave{e}_6 > 0 \tag{1.14}$$

If against the alternative in (1.12), (1.13) and (1.14), the following restrictions hold:

$$\ddot{a}_4 = \ddot{o}_2 = \ddot{o}_3 = \ddot{o}_4 = \ddot{o}_5 = \grave{e}_5 = 0; \ddot{o}_1 = \ddot{o}_6 = 1$$

then the UIP condition would hold from (1.13), money demand and exchange rates would not depend on the level of debt from (1.12) and (1.14), which provides strong evidence that debt influences have little bearing on the determination of domestic interest rates (an indication of the Ricardian equivalence). It would also provide a clear evidence of whether a general portfolio balance model of the exchange rate, in which the risk premium exists and debt affects domestic interest rates, holds.

III. THE STATISTICAL MODEL AND DIAGNOSTIC TESTS

The multivariate time series model that is used to implement the economic framework outlined above and to perform a set of tests for the debt/interest rate relationship in long-run, is now briefly described below. Consider the p-dimensional VAR(K) model in error-correction form

$$\ddot{A} X_t = \mu + \ddot{O} D_t + \text{Đ} X_{t-1} + \tilde{A}_1 \ddot{A} X_{t-1} + \ldots + \tilde{A}_{K-1} \ddot{A} X_{t-K+1} + \varepsilon_t, \; t = 1, \ldots, T, \tag{2.1}$$

where X_t is I(1), $X_0, X_{-1}, \ldots, X_{-K+1}$ are fixed, $\tilde{A}_1, \ldots, \tilde{A}_K$, $\ddot{O}$ and μ are the regression coefficients, ε_t are i.i.d $N(0, \acute{O})$, and D_t are centred seasonal dummies. The long-run impact matrix Đ contains information about the long-run relationships between the variables. Let r denote the rank of Đ. Then, the hypothesis of at most r

($0<r<p$) cointegration vectors is defined as a reduced rank condition $H_0(r)$: $\Pi = \alpha\beta'$, where β is a $p\times r$ matrix of cointegrating vectors and α is a $p\times r$ matrix of adjustment coefficients of individual series toward the r long-run relationships ($\beta' X_t$).

Johansen and Juselius (1990) define the maximum likelihood procedure for estimating the above model and propose the following likelihood ratio statistics for determining the rank (r) of the Π matrix:

$$\text{trace stat } (H_0(r)|H_1(p)) = -T \sum_{i=r+1}^{p} \ln(1-\hat{\lambda}_i) \tag{2.2}$$

$$\lambda\text{max stat } (H_0(r)|H_1(r+1)) = -T(1-\hat{\lambda}_{r+1}) \tag{2.3}$$

where $\hat{\lambda}_1,\ldots,\hat{\lambda}_p$ are the estimated squared canonical correlations between the first differences and the levels of variables. Cushman and *al.* (1995) find, however, that the power of the maximal eigenvalue test is higher than for the trace test, which confirms the expectation of Johansen and Juselius (1989, pp.19).

The Johansen procedure also provides an opportunity to formulate and test a set of interesting economic structural hypotheses about the cointegrating relations. Two of these restrictions are of interest for the current study. First, a linear hypothesis applying equally on all cointegrating vectors is formulated as: $\beta = H\varphi$, where H is $p\times s$, φ is $s\times r$ and tested by using the likelihood ratio (LR) statistic

$$LR1 = -T\sum_{i=1}^{r} \ln\{(1-\tilde{\lambda}_i)/(1-\hat{\lambda}_i)\} \sim \chi^2_{(r(p-s))} \tag{2.4}$$

where $\tilde{\lambda}_j$ are obtained from the restricted model. Second, for cases of more than one cointegrating vectors, Johansen (1995, pp. 70-120) presents some tests for restrictions on subsets of the vectors which permit to identify them as economically meaningful relations. We consider a more general hypothesis of the form $\beta = (H_1\varphi_1,\ldots,H_r\varphi_r)$, where H_i is $p\times s_i$ and φ_i is $s_i\times 1$, with $1\le s_i\le p$. The LR statistic for these constraints, obtained at the convergence of the switching algorithm, is asymptotically distributed as χ^2 with d.f. given by $\sum_{i=1}^{r}(p-r-s_i+1)$, provided β is identified. For special cases of two types of restrictions, the null hypothesis is $\beta = (H_1\varphi_1, H_2\varphi_2)$, where φ_i is $s_i\times r_i$ with $r_i\le s_i\le p$ and $r_1+r_2=r$, the LR test statistic is given by

$$LR2 = -T\{\sum_{i=1}^{r_1}\ln(1-\hat{\eta}_i)+\sum_{i=1}^{r_2}\ln(1-\tilde{\lambda}_i)-\sum_{i=1}^{r}\ln(1-\hat{\lambda}_i)\} \tag{2.5}$$

where $\hat{\eta}$ and $\tilde{\lambda}$ come from the constrained estimation at the convergence ($\hat{\beta}_1 = H_1\hat{\varphi}_1$ and $\hat{\beta}_2 = H_2\hat{\varphi}_2$) .

3.1 Time series Properties of the Data

We use quarterly and seasonally unadjusted data from 1957:1 to 1993:4 on the variables specified in the models (1.9) and (1.11) for Canada, Germany, the UK and the US. In long-run, however, one period ahead expected inflation rates and changes in the exchange rate do not matter since $\text{Ä}p^{e}_{t+1}=0$ and $\text{Ä}s^{e}_{t+1}=0$. Thus, the variables of interest comprise the nominal M2, the national income proxied by the real GNP, the total government outstanding debt (federal plus state liabilities), the nominal exchange rate and the domestic interest rate measured by long-term government bond yields. The series are stacked in a vector as follows: $X_t=[m_t,y_t,i_t,i_{ft},b_t,s_t]$. All the data are taken from the IMF *International Financial Statistics* (IFS) and the OECD *Main Economic Indicators.* Here, an increase in the exchange rate level is defined as a currency appreciation. In all cases the US interest rate is used as the foreign interest rate, whereas the German long-term government bond yield is used for the US model. Accordingly, the bilateral exchange rates are expressed in terms of national currency per unit of US dollar, while for the United States it is given by $US per unit of Mark. As indicated in the economic model, all the variables are expressed in natural logs except interest rates. It is worth noting that the choice of the 1957-1993 period was motivated by the will to provide more weight for the period during the Reagan-Bush regime, characterized by large and persistent public deficits.

The statistical model (2.1) requires that all the included variables are integrated of order one. Nevertheless, some economic series also exhibit substantial seasonality. We then use the test procedure developed by Hylleberg, Engle, Granger and Yoo (1990) [HEGY] for quarterly data, which allows one to jointly test for both seasonal and nonseasonal unit roots. This approach contains the Dickey-Fuller test as a special case. To perform the HEGY tests for each series x_t, one estimates by ordinary least squares the following equation:

$$\varphi^{*}(B)y_{4t}=\delta_1 y_{1t-1}+\delta_2 y_{2t-1}+\delta_3 y_{3t-2}+\delta_4 y_{3t-1}+\sum_{i=1}^{n}\phi_i y_{4t-i}+\sum_{s=2}^{4}\theta_s D_{st}+\mu_t+\varepsilon_t, \tag{3.1}$$

where

$$y_{1t}=(1+B+B^2+B^3)x_t,\ y_{2t}=-(1-B+B^2-B^3)x_t, y_{3t}=-(1-B^2)x_t,\ y_{4t}=(1-B^4)x_t,\ \text{and}\ \varphi^{*}(B)=1.$$

Equation (3.1) includes the linear trend μ_t, the seasonal dummies D_{st} and n additional lags of y_4 to whiten the error terms ε_t. Hypotheses about various unit roots are tested as follows. For frequencies 0 and π, one examines the t-statistics for $\delta_1=0$ and $\delta_2=0$ against $\delta_1<0$ and $\delta_2<0$, respectively. For the frequency $\pi/2$, one jointly tests $\delta_3=0$ and $\delta_4=0$ with an F-statistic. Critical values for the t- and F-statistics are provided in

HEGY. There will be nonseasonal unit roots if δ_2 and either δ_3 or δ_4 are different from zero. This therefore requires the rejection of both the test for $\delta_2 = 0$ and the joint test for $\delta_3 = 0$ and $\delta_4 = 0$. To conclude that the series exhibits no unit roots at all and is therefore stationary, we must find that each of the δ's is different from zero (except possibly either $\delta_3 = 0$ or $\delta_4 = 0$). Table 1 reports the results of HEGY tests for the four countries. As indicated in column 4, there is no series for which the t-statistic for δ_1 rejects at the 5 percent level the null hypothesis of a unit root at the nonseasonal frequency. Furthermore, the last five columns show that no series exhibits seasonal unit roots for all countries.

3.2 Model Specification and Estimation

The Johansen multivariate VAR model is also based on two other important assumptions which are independence over time and normality of disturbance terms. We therefore test for those assumptions to determine the optimal autoregressive order (*K*) of the model applied to each country. We start by estimating VAR models with lag lengths *K* running from 1 to 12 and retain the shortest lag length for which all equations are left with white noise residuals. To save space, we only report in Table 2 the diagnostic test results for the most satisfactory lag length specification for each country. The computed statistics indicate that all the estimated residuals are normally distributed, free from serial correlation problems, and do not exhibit conditional heteroscedasticity when *K*=9 for the US, *K*=6 for the UK, *K*=9 for Germany and *K*=8 for Canada.

Table 3 reports the LR statistics for the cointegration ranks. Since, for all cases, both rank tests did not contradict each other, we only present the results for the λ_{max} tests to save space (those for the trace tests are available upon request). The results support at the 5% significance level, *r*=2 for the US, *r*=1 for Germany, *r*=2 for the UK, and *r*=2 for Canada. The LR tests for deterministic trends, which are not reported, suggested an EC model including a constant term for Canada and the UK, a linear trend for Germany, and a time trend only in the cointegrating vectors for the US (see the first line of Table 3). The LR tests for deterministic trends, which are not reported, suggested an EC model including a constant term for Canada and the UK, a linear trend for Germany, and a time trend only in the cointegrating vectors for the US (see the first line of Table 3).

Table 4 gives the estimated cointegrating vectors for each country. They may embody some of the three relevant long-run relationships derived from our theoretical framework (steady state UIP, the exchange rate or money equation) since most of the cointegrating coefficients have expected signs: a positive relationship between debt and domestic interest rates, as well as between debt, the exchange rate and money. As pointed out by Johansen and Juselius (1990) however, it is quite difficult to find all the expected signs in each cointegration

vector since their choice is mainly based on the stationarity assumption.

Up to now we have been concerned with the specification and the estimation of an adequate ECM for each country. The reader, however, will recall that the main purpose of this study is to examine the debt-interest rate relationship directly, and indirectly through the effects of debt on the exchange rate and money demand. In what follows, the effects of permanent changes in the level of debt are tested from cointegrating vectors, whereas the short-run dynamics are studied from restricted error-correction models.

IV. EMPIRICAL RESULTS

4.1 Identification of Long-run Structure

In this section, we perform the Johansen tests on individual $\hat{a}$ (presented in section 3) to examine whether the cointegrating structure obtained for each country may be identified as the three interpretable economic relationships (1.12), (1.13) and (1.14) defined in section 2. Given the system $X' = (m, y, i, i_f, b, s)$, these individual identifying restrictions are formulated as follows.

A1: $\hat{a}_i' = (0,0,1,-1,0,0)$ or $\hat{a}_i' = \varphi_i' H_i'$, (4.1)

where $H_i' = [0 \quad 0 \quad 1 \quad -1 \quad 0 \quad 0]$

A2: $\hat{a}_i' = (1, *, \ddot{a}_3, -\ddot{a}_3, -1, 0)$ or $\hat{a}_i' = \varphi_i' H_i'$, (4.2)

where $H_i' = \begin{bmatrix} 1 & 0 & 0 & 0 & -1 & 0 \\ 0 & 1 & 0 & 0 & 0 & 0 \\ 0 & 0 & 1 & -1 & 0 & 0 \end{bmatrix}$

A3: $\hat{a}_i' = (1, *, \grave{e}_3, -\grave{e}_3, *, -1)$ or $\hat{a}_i' = \varphi_i' H_i'$, (4.3)

where $H_i' = \begin{bmatrix} 1 & 0 & 0 & 0 & 0 & -1 \\ 0 & 1 & 0 & 0 & 0 & 0 \\ 0 & 0 & 1 & -1 & 0 & 0 \\ 0 & 0 & 0 & 0 & 1 & 0 \end{bmatrix}$.

The design matrices H_i define, respectively, under assumptions A1, A2, and A3, the cointegrating vectors $\hat{a}_i$ which are interpreted as the interest rate differential (or the long-run UIP), the long-run money relation, and the

exchange rate equation. The acceptance of restriction A1 or that of A2 and A3 after excluding the debt variable implies that permanent shifts in the stock of debt have no wealth effect on private demand, as it does not influence asset prices (the nominal interest rate and the exchange rate). Such a finding will tend to support the Ricardian equivalence while an opposite result will support the Keynesian-Neoclassical point of view. Identifying test results for the space spanned by the $\hat{a}$ vectors are analyzed below for each country. We test various hypotheses and present only those which are accepted or are of primary interest for our analysis.

Test Results for Germany

The analysis is easier for Germany where we found only one cointegrating vector. The significance tests for cointegrating coefficients, reported in Table 4, indicate that only the exchange rate and money variables can be excluded from the cointegration relation at the 5% level. This implies that the interest rate differential cannot be stationary by itself, as in the restriction (4.1). We need, however, to identify whether the German cointegrating vector can be interpreted as a long-run interest rate equation (1.13) or money relation (1.14). The vector cannot be the exchange rate equation (1.12) because the money variable is excluded. In this regard, we first formulate an hypothesis that both money and the exchange rate are excluded from the cointegration relation and the coefficients of interest rates sum to zero, that is, $(0, *, \varphi_{13}, -\varphi_{13}, *, 0)$ is a cointegrating vector. The LR test is calculated to be 17.68 which rejects the null hypothesis in a $\div^2_{(3)}$ distribution, at the 5% significance level. The restrictions that the cointegration vector might be a long-run money demand relation, say $\hat{a}' = (1, *, -\varphi_{13}, \varphi_{13}, -1, 0)$ and $\hat{a}' = (1, *, *, *, -1, 0)$, were all rejected. The tests fail to reject only the hypothesis that the interest rate differential is stationary after including income and the relative supply of assets (b-s-b_f, b_f=0). That is,

$$\hat{a} = H\varphi \text{ with } H = \begin{bmatrix} 0 & 0 & 0 & 0 \\ 1 & 0 & 0 & 0 \\ 0 & 1 & 0 & 0 \\ 0 & -1 & 0 & 0 \\ 0 & 0 & 1 & 0 \\ 0 & 0 & 0 & 1 \end{bmatrix}. \tag{4.4}$$

The computed LR test for this restriction is 3.86, and should be evaluated in a $\div^2_{(2)}$ corresponding to a p-value of 15 percent. Thus, the hypothesis (4.4) is consistent with the data and we conclude the analysis of the model

for Germany by the following restricted vector: $\hat{\hat{a}}^{/}=(0.00,-1.97,0.07,-0.07,-1.23,1.00)$. This suggests that, in this country, permanent changes in government debt positively affect domestic interest rates and money demand.

The Test Results for the United States, the United Kingdom and Canada

We find a two-dimensional cointegration space for the US, the UK and Canada which seems to contain coefficients which correspond to the theoretical economic relations defined in equations (4.1), (4.2), and (4.3). We first test whether the interest rate differential is stationary by itself in each country. This hypothesis is formulated in Table 5 and the LR statistics are significant for all cases (The d.f. is given by $r_1(p-r)=1\times(6-2)=4$). This means that the interest rate differential as a stationary relationship alone is not supported by the data. It is an indication that permanent shifts in the level of government debt may affect domestic interest rates, money demand or the exchange rate. For a conclusive decision, however, we need to investigate whether the cointegration space contains the vectors which are interpreted as a money demand relation or an exchange rate equation including government debt. In what follows, we report for each country the estimation and test results for only $\hat{a}$ structures that are compatible with the data.

For the United States, the LR test for the restrictions defined in Table 6, $LR2(4)=8.10$, is not significant at standard levels (5, 2.5 and 1 per cent). Thus, the US data supports the existence of a long-run exchange rate equation which is homogeneous of degree one in money with the domestic and foreign interest rate coefficients summing to zero. The second cointegrating vector is compatible with a long-run equation of domestic interest rates which does not include government debt and the exchange rate. The economic interpretation of this finding is that the first cointegrating vector is consistent with the assumptions of the general portfolio balance model, while the second vector suggests the insignificance of permanent changes in public debt on interest rates. For a conclusive decision, however, these results need to be complemented by investigating the short-run dynamics of variables.

The results for the UK are given in Table 7. The LR test statistic (8.65) for the UK long-run structure, which is asymptotically distributed as a $\div^2_{(4)}$, fails to reject the hypothetical structure $\hat{a}=(H_1\varphi_1,H_2\varphi_2)$ at 5% significance level. Thus it is not against the data to assume that the first vector is a long-run portfolio equilibrium equation which links the interest rate differential to the relative supply of domestic assets (b-s-b_f, b_f=0), and the second vector is a long-run exchange rate equation derived from the portfolio balance framework. Data analysis for the UK suggests therefore that permanent shifts in the level of government debt have a

significant impact on interest rates, money demand and the exchange rate.

The structural hypothesis on â, for Canada, is defined in Table 8. The computed test statistic 9.21 to be compared with a $\chi^2_{(4)}$, indicates non rejection of the defined long-run structure at all standard significance levels. Thus, the cointegration space contains a long-run money relation with income homogeneity, and a long-run exchange rate equation with the coefficients to interest rates being equal with opposite signs. Consequently, the data for Canada is compatible with the general portfolio balance model which suggests that in the long-run, government debt through its wealth effect on private demand causes higher domestic interest rates and money demand and appreciates domestic currency.

4.2 Short-Run Dynamics of Cointegrated Variables

The short-run adjustment of variables is investigated by means of error-correction (EC) models where the long-run restrictions on â identified above, were imposed.As indicated by Kalulumia and Yourougou (1997), temporal causality in the restricted EC framework may ensue from two potential sources. The constrained error-correction terms which measure the impact of transitory deviations of variables from their long-run equilibria, and the past changes in series values (the standard Granger causality test). The coefficients to EC terms (á) represent the endogenous responses of real money balance growth rate, domestic interest rates and the exchange rate to market disequilibrium.

The results of the maximum likelihood estimation of the EC models for real money M2, domestic interest rates and the exchange rate are presented in Table 9, for all countries. To save space, we only present the coefficients of lagged debt variables that are of primary interest to the current study. For the US, the results indicate that only the first EC term, derived from the long-run exchange rate equation, is significant in the real M2 equation. The significance of this constrained EC term implies causality from all six variables (including government debt) to the real M2. The F-test with a value of 2.78 (p-value=0.015) also indicates joint significance of all lagged debt variables at the 5% level. The US nominal interest rate and exchange rate, however, hardly react to departures from the equilibrium as well as to the lagged values of public debt. The coefficients to EC terms and to lagged debt values were insignificant in both *i* and *s* equations. Overall, these US results suggest that transitory changes in government debt affect money demand in the short-run.

For Germany, the causality test results indicate that the EC term is significant in the real money and exchange rate equations, suggesting causality of government debt to both dependent variables. The interest rate, however, does not adjust since in its equation, the t-est fails to reject the null hypothesis of zero coefficient to the EC term and the lagged debt coefficients are jointly insignificant (F(8,54)=1.96, p-value=0.692)). For the

UK, the results clearly indicate that real money and the exchange rate strongly reacts to transitory changes in the level of debt since both EC terms are significant in their respective error-correction equations. Canada is the only case where both EC terms are significant in the interest rate equation. This result, combined with the fact that the second EC term is also significant in the real money and exchange rate equations, suggests the evidence of a short-run impact of public debt on all three interest-rate related variables in this country (money demand, interest rates and the exchange rate).

In general, the above findings advocate the use of indirect tests of the debt-interest rate relationship as, with the exception of Canada, the domestic interest rate hardly adjusts to the dynamics of government debt. This may be attributed to the low variability of the variable or to the offsetting impact of premature capital inflows on interest rates. The results also reveal that EC terms which measure market disequilibrium are the most important source of short-run causality between debt and interest-rate related variables. The past values of debt are not jointly significant in all the cases except the US money equation. Another noticeable result is that, the short-run impact of government debt appears to be weaker in the US as only one of the three relevant dependent variables adjusts to changes in the level of public sector indebtedness.

V. CONCLUSION

This paper explored the impact of government debt on interest rates in the United States, Germany, the United Kingdom and Canada. One feature of the study was to address the issue in the Johansen-Juselius multivariate cointegration model and the general portfolio balance framework which allows for both direct and indirect tests of the relationship between debt and interest rates. Indirect tests of the debt effects on interest rates were investigated through the impact of debt on the exchange rate and money demand. The evidence indicated that both transitory and permanent changes in the level of government debt cause higher domestic interest rates and money demand and appreciate the exchange rate in all four countries under study. This suggested that a shift in the level of government indebtedness, *ceteris paribus*, will have all the traditional effects predicted by standard Keynesian models. Nevertheless, the fact that the short-run causality arises mainly from adjustment to asset market disequilibrium is an indication that households behave more as rational than myopic individuals. Consequently, one explanation for the finding of the short-tun causality between debt and interest-rates may be firstly, the asymmetry of market information which prevents rational individuals from evaluating exactly the government debt/tax values over time and sencondly, the uncertainty ensued from individual finite life spans.

Overall, our results are consistent with the previous findings supporting the Keynesian and Neoclassical hypotheses. The evidence also fits well with some aspects of the general portfolio approach suggesting that capital is not perfectly mobile in the international financial markets since there exists a non null risk premium

on asset holdings.

Table 1. Tests For Seasonal Unit Roots

Variables	Period	Lags	`t' : ð$_1$ 0	`t' : ð$_2$ ð	`t' : ð$_3$ ð / 2	`t' : ð$_4$ ð / 2	`F': ð$_3$ ∩ ð$_4$
				United States			
m	57:1-93:4	6	-1.11	-1.29	1.74	3.72*	8.74*
y	57:1-93:4	9	-1.56	-3.84*	4.76*	4.41*	21.44*
i	57:1-93:4	4	-1.39	-3.37*	3.80*	4.53*	19.66*
b	57:1-93:4	2	-1.85	-3.04*	3.40*	5.02*	20.71*
s	57:1-93:4	2	-2.86	-5.86*	5.85*	3.11*	21.95*
				Germany			
m	57:1-93:4	1	-2.37	-5.77*	5.49*	5.17*	28.54*
y	57:1-93:4	1	-2.10	-5.03*	5.91*	4.04*	24.90*
i	57:3-93:4	1	-1.83	-6.88**	4.59*	4.64*	21.33*
b	57:1-93:4	2	-1.16	-3.62*	4.27*	4.24*	22.38*
s	57:1-93:4	3	-1.64	-3.31*	4.27*	0.99	9.86*
				United Kingdom			
m	57:1-93:4	1	-2.30	-5.35*	6.58*	3.83*	29.01*
y	57:1-93:2	1	-2.32	-4.59*	5.77*	4.41*	26.31*
i	57:3-93:4	1	-1.14	-6.71*	5.71*	6.05*	30.83*
b	57:1-93:4	1	-2.70	-2.66*	4.29*	2.58*	12.44*
s	57:1-91:4	10	-2.95	-2.82*	4.66*	1.22	11.99*
				Canada			
m	57:1-93:4	1	-1.57	-5.46*	9.21*	2.99*	46.88*
y	57:1-93:4	1	-0.12	-7.19*	6.32*	6.93*	44.41*
i	57:1-93:4	17	-1.43	-2.25*	3.00*	3.89*	12.05*
b	57:1-93:4	1	-1.69	4.00*	4.07*	4.70*	19.36*
s	57:1-93:4	7	-2.94	-4.98*	3.14*	0.68	5.13*

Notes: 1. The superscripts * and ** indicate respectively the rejection at 5% and 10% level of the null hypothesis.
2. Lags is the number of lagged values of the dependant variable included in the regression.
3. We use critical values presented in Hylleberg et *al.* (1990) for the regressions including a constant, three seasonal dummies and a time trend.
4. The results from regressions including a constant and/or a time trend yield just slight changes in figures, without affecting our conclusions about unit root tests.

Table 2. Error term diagnostics

Equations	Normality	Ljung-Box Q(l)	GARCH	Normality	Ljund-Box Q(l)	GARCH
	United States (K=7)			**Germany (K=9)**		
m	1.24	20.38	5.58	5.58	8.76	1.67
y	1.00	18.95	3.79	3.79	6.60	12.11
i	5.61	19.40	2.07	2.07	7.08	17.64
i_f	2.12	17.57	4.01	4.01	6.28	9.83
b	1.30	21.96	3.22	3.22	7.21	15.35
s	5.17	12.05	3.91	3.91	6.54	14.68
	United Kingdom (K=6)			**Canada (K=8)**		
m	1.10	3.12	1.00	2.25	18.03	6.83
y	1.41	15.10	11.60	1.04	16.90	5.93
i	1.80	21.35	5.59	1.10	18.42	14.51
i_f	1.56	16.18	3.23	2.06	21.64	8.48
b	4.61	12.91	5.31	0.81	15.71	5.41
s	1.18	20.70	1.75	2.89	14.01	2.47

Notes: [a]The GARCH and normality tests are distributed, respectively, as $\div^2_{(K)}$ and $\div^2_{(2)}$ variables. At the 5% level, their respective critical values are $\div_{(2)} = 5.99$, $\div_{(6)} = 14.46$, $\div_{(8)} = 15.51$ and $\div_{(9)} = 16.93$.

[b]Ljung-Box Q(l)-statistics, are asymptotically distributed as a $\div^2_{(l)}$ whose critical value is 23.69 at the 5% level, for *l*=14.

Table 3: Tests for the Cointegration Rank

	United States		**Germany**		**United Kingdom**		**Canada**	
D-tr	$\mu_0 + á\hat{a}'(X'_{t-1}, t)'$		$\mu_0 + \mu_1 t + á\hat{a}'X_{t-1}$		$\mu_0 + á\hat{a}'X_{t-1}$		$\mu_0 + á\hat{a}'X_{t-1}$	
Ho	ëmax	ëcrit	ëmax	ëcrit	ëmax	ëcrit	ëmax	ëcrit
$r \le 0$	54.34*	43.97	54.62*	42.48	69.79*	39.37	50.02*	39.37
$r \le 1$	38.01*	37.52	32.70	36.41	36.45*	33.46	41.23*	33.46
$r \le 2$	30.11	31.16	19.46	30.33	26.34	27.07	19.94	27.07
$r \le 3$	17.87	25.54	14.03	23.78	19.96	20.97	17.76	20.07
$r \le 4$	16.28	18.96	12.68	18.87	13.07	14.07	10.45	14.07
$r \le 5$	8.75	12.25	2.55	3.74	2.29	3.76	0.40	3.76

Notes: [a]Superscript * indicates a rejection at the 5% level of the null hypothesis of at most *r* stationary linear combinations of the series against the alternative of at most *r*+1 such combinations. D-tr are the deterministic trends in EC Models.
[b]Critical values are from Osterwald-Lenum (1992).

Table 4: Estimated Cointegrating Vectors and Likelihood Ratio Statistics

Vectors& LR-Statistics	Variables						
	m	*y*	*i*	i_f	*b*	*s*	*t*
				United States (*r*=2)			
$\hat{\hat{a}}'$	1.000	7.921	-0.111	0.076	-2.017	-0.850	-7.256
	1.000	-1.613	0.028	-0.036	-0.424	0.922	0.833
LR1(2)	10.29	12.67	15.66	8.87	7.85	9.46	
Prob-v	(0.010)	(0.002)	(0.000)	(0.012)	(0.020)	(0.009)	
				Germany (*r*=1)			
$\hat{\hat{a}}'$	1.000	-1.224	-0.048	0.045	-0.814	-0.566	
LR1(1)	2.06	8.48	16.14	17.86	17.68	3.78	
Prob-v	(0.151)	(0.004)	(0.000)	(0.000)	(0.000)	(0.052)	
				United Kingdom (*r*=2)			
$\hat{\hat{a}}'$	1.000	-4.809	-0.124	0.153	1.168	7.878	
	1.000	-1.795	-0.040	0.062	0.083	1.465	
LR1(2)	7.48	11.19	7.01	13.80	14.87	13.29	
Prob-v	(0.024)	(0.004)	(0.03)	(0.001)	(0.001)	(0.001)	
				Canada (*r*=2)			
$\hat{\hat{a}}'$	1.000	-0.995	-0.233	0.163	-0.133	2.526	
	1.000	-2.041	-0.037	0.041	1.119	-1.821	
LR1(2)	4.04	8.75	17.42	14.87	20.54	17.92	
Prob-v	(0.133)	(0.013)	(0.000)	(0.001)	(0.000)	(0.000)	

Note: The number in parentheses for each country gives the significance level.

Table 5. Identification of the interest rate differential for all countries, $\hat{a}=(H_1\varphi_1,\varphi_2)$

$$\hat{a}=\begin{bmatrix} 0 & * \\ 0 & * \\ 1 & * \\ -1 & * \\ 0 & * \\ 0 & * \end{bmatrix} \qquad H_1=\begin{bmatrix} 0 \\ 0 \\ 1 \\ -1 \\ 0 \\ 0 \end{bmatrix}$$

<u>Tests of identifying restrictions</u>

United States	$LR2(4)=18.42$ *p-value*=0.00
United Kingdom	$LR2(4)=14.87$ *p-value*=0.00
Canada	$LR2(4)=26.54$ *p-value*=0.00

Table 6. US: Identification of the general structure $\hat{a}=(H_1\varphi_1,H_2\varphi_2)$

$$\hat{a}=\begin{bmatrix} 1 & * \\ * & 0 \\ \varphi_{31} & \varphi_{32} \\ -\varphi_{31} & -\varphi_{32} \\ * & 0 \\ -1 & 0 \\ * & * \end{bmatrix} \qquad H_1=\begin{bmatrix} 1 & 0 & 0 & 0 & 0 \\ 0 & 1 & 0 & 0 & 0 \\ 0 & 0 & 1 & 0 & 0 \\ 0 & 0 & -1 & 0 & 0 \\ 0 & 0 & 0 & 1 & 0 \\ -1 & 0 & 0 & 0 & 0 \\ 0 & 0 & 0 & 0 & 1 \end{bmatrix} \qquad H_2=\begin{bmatrix} 1 & 0 & 0 & 0 \\ 0 & 1 & 0 & 0 \\ 0 & 0 & 1 & 0 \\ 0 & 0 & -1 & 0 \\ 0 & 0 & 0 & 0 \\ 0 & 0 & 0 & 0 \\ 0 & 0 & 0 & 1 \end{bmatrix}$$

Tests of identifying restrictions $LR2(4)=8.10$ *p-value*=0.09

	Restricted estimates $\hat{\hat{a}}$	Restricted estimates $\hat{a}$	T-test values for $\acute{a}$
m	1.000 1.000	−0.003 −0.004	2.09 −0.39
y	0.169 1.564	−0.003 0.014	−1.64 1.48
i	−0.168 −0.047	−0.330 2.164	−1.59 1.84
i_f	0.0168 0.047	0.182 −0.648	1.16 −0.75
b	3.778 0.000	−0.011 0.069	−6.30 7.04
s	−1.000 0.000	−0.001 0.031	−0.20 1.29
t	−0.020 −0.017		

Table 7. UK: Identification of the general structure $\hat{a}=(H_1\varphi_1, H_2\varphi_2)$

$$\hat{a}=\begin{bmatrix} 1 & 0 \\ * & 0 \\ \varphi_{31} & \varphi_{32} \\ -\varphi_{31} & -\varphi_{32} \\ * & * \\ -1 & * \end{bmatrix} \qquad H_1=\begin{bmatrix} 0 & 0 & 0 \\ 0 & 0 & 0 \\ 1 & 0 & 0 \\ -1 & 0 & 0 \\ 0 & 1 & 0 \\ 0 & 0 & 1 \end{bmatrix} \qquad H_2=\begin{bmatrix} 1 & 0 & 0 & 0 \\ 0 & 1 & 0 & 0 \\ 0 & 0 & 1 & 0 \\ 0 & 0 & -1 & 0 \\ 0 & 0 & 0 & 1 \\ -1 & 0 & 0 & 0 \end{bmatrix}$$

Tests of identifying restrictions $LR2(4)=8.65$ $p\text{-}value=0.071$

	Restricted estimates $\hat{\hat{a}}$	Restricted estimates $\hat{a}$	T-test values for $\hat{a}$
m	0.000 1.000	0.155 −0.044	2.25 −2.43
y	0.000 −2.864	0.011 −0.002	0.58 −0.46
i	0.013 0.046	−2.552 0.598	−1.30 1.15
i_f	−0.013 −0.046	5.151 −1.310	2.56 2.47
b	−0.279 −1.319	−2.426 0.910	−3.44 −4.88
s	1.000 −1.000	−0.148 0.038	−3.53 3.40

Table 8. Canada: Identification of the general structure $\hat{a}=(H_1\varphi_1, H_2\varphi_2)$

$$\hat{a}=\begin{bmatrix} 1 & 1 \\ -1 & * \\ * & \varphi_{32} \\ * & -\varphi_{32} \\ * & * \\ 0 & -1 \end{bmatrix} \qquad H_1=\begin{bmatrix} 1 & 0 & 0 & 0 \\ -1 & 0 & 0 & 0 \\ 0 & 1 & 0 & 0 \\ 0 & 0 & 1 & 0 \\ 0 & 0 & 0 & 1 \\ 0 & 0 & 0 & 0 \end{bmatrix} \qquad H_2=\begin{bmatrix} 1 & 0 & 0 & 0 \\ 0 & 1 & 0 & 0 \\ 0 & 0 & 1 & 0 \\ 0 & 0 & -1 & 0 \\ 0 & 0 & 0 & 1 \\ -1 & 0 & 0 & 0 \end{bmatrix}$$

Tests of identifying restrictions $LR2(4)=9.21$ $p\text{-}value=0.056$

	Restricted estimates $\hat{\hat{a}}$	Restricted estimates $\hat{a}$	T-test values for $\hat{a}$
m	1.000 −1.000	0.029 0.173	1.18 5.20
y	−1.000 1.638	0.012 −0.001	1.33 −0.028
i	−0.165 0.040	4.498 4.394	4.49 3.25
i_f	0.118 −0.040	3.130 3.392	2.52 2.02
b	−0.159 −0.647	−0.066 −0.024	−3.91 −1.10
s	0.000 1.000	−0.036 −0.087	−1.32 −2.39

Table 9: Short-Run Dynamics derived from Constrained Error-Correction Models

Equations		*errc1*	*errc2*	*b(t-1)*	*b(t-2)*	*b(t-3)*	*b(t-4)*	*b(t-5)*	*b(t-6)*	*b(t-7)*	*b(t-8)*
US	*m*	-0.003*	-0.004	0.156*	0.074	0.107	0.111	0.110	-0.060		
		(-2.09)	(-0.39)	(2.32)	(1.16)	(1.82)	(1.87)	(1.77)	(-1.06)		
	i	-0.338	0.014	-13.89	-8.145	-9.093	-1.986	0.880	0.256		
		(-1.59)	(1.84)	(-1.57)	(-0.97)	(-1.18)	(-0.26)	(0.11)	(0.03)		
	s	-0.001	0.034	0.112	-0.296	-0.435*	-0.328	-0.183	0.029		
		(-0.20)	(1.29)	(0.57)	(-1.57)	(-2.52)	(-1.89)	(-1.00)	(0.17)		
Germany	*m*	0.067*	_	0.117	-0.050	-0.040	0.011	0.074	0.033	-0.002	0.066
		(2.79)		(1.68)	(-0.71)	(-0.59)	(0.16)	(0.98)	(0.44)	(-0.03)	(0.95)
	i	0.885	_	14.50*	0.060	0.201	-12.66*	0.843	5.153	9.010	-1.868
		(0.66)		(2.62)	(0.02)	(0.05)	(-2.33)	(0.20)	(1.22)	(1.61)	(-0.47)
	s	-0.205*	_	0.111	-0.062	0.191	-0.016	-0.445*	-0.039	-0.060	0.257
		(-4.11)		(0.76)	(-0.42)	(1.34)	(-0.11)	(-2.83)	(-0.25)	(-0.41)	(1.76)
UK	*m*	0.155*	-0.044*	-0.015	-0.016	-0.010	-0.001	-0.012			
		(2.25)	(-2.43)	(-1.71)	(-1.79)	(-1.20)	(-0.13)	(-1.51)			
	i	-2.552	0.598	0.047	-0.123	-0.045	0.030	0.185			
		(-1.30)	(1.15)	(0.22)	(-0.59)	(-0.24)	(0.17)	(0.10)			
	s	-0.148*	0.038*	0.013*	0.008	0.004	0.008	-0.004			
		(-3.54)	(3.40)	(2.52)	(1.45)	(0.91)	(1.75)	(-0.97)			
Canada	*m*	0.029	0.173*	-0.072	-0.085	-0.030	-0.169	0.082	0.030	-0.012	
		(1.18)	(5.19)	(-0.58)	(-0.67)	(-0.23)	(-1.35)	(0.64)	(0.25)	(-0.10)	
	i	4.498*	4.394*	-8.011	-1.947	1.103	6.978	5.490	-0.368	0.772	
		(4.49)	(3.25)	(-1.58)	(-0.38)	(0.21)	(1.38)	(1.05)	(-0.10)	(0.15)	
	s	-0.036	-0.087*	0.078	0239	-0.211	0.089	-0.001	-0.018	-0.079	
		(-1.32)	(-2.39)	(0.57)	(1.75)	(-1.49)	(0.66)	(-0.01)	(-0.13)	(-0.57)	

Notes: The numbers in parentheses are the values of the t-tests. Superscript * indicates significance at the 5% level.

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