Are International R&D Spillovers Costly for the US?

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Abstract: Coe and Helpman (1995) and others report positive and equivalent R&D spillovers across G7 countries. We argue that their homogeneity constraint on spillovers across G7 countries is inappropriate, and show that it is rejected by the data. Extending the data set and applying new empirical approaches, we find: (i) R&D spillovers are extremely heterogeneous across G7 countries; (ii) panel estimates do not correspond to country specific estimates and conceal important cross-country differences in knowledge diffusion; and (iii) the US is a net loser in terms of international R&D spillovers. Our interpretation is that when competitors 'catch-up' technologically, they challenge US market shares and investments worldwide and this has implications for US productivity.

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I. Related Literature

In a seminal paper, Coe and Helpman [1995; henceforth, CH] provide empirical evidence on trade related international R&D spillovers by using panel data for 21 OECD countries and Israel over the period 1971-1990. Their main findings are that the domestic (S^d) and foreign (S^f) R&D capital stocks affect domestic total factor productivity (TFP) positively; S^d has bigger effect than S^f on large countries whereas the opposite holds on smaller countries; and the more open the smaller countries are the more likely they are to benefit from S^f. According to Navaretti and Tarr (2000, p. 2) CH's work is the 'most quoted reference' in the field.

The finding of significant R&D spillovers across countries is consistent with the growth literature. The endogenous growth literature, in particular, posits endogenous innovations as key propagators of long-run economic growth[1]. In these models technology spills over through international trade and triggers productivity increases in importing countries so long as there exists a positive mark-up between the marginal product and the cost of imported intermediate goods.[2] Productivity transmissions of this kind are not only important for developed countries but they are also vital in promoting economic growth in developing countries. Indeed, Coe et al. (1997) report significant R&D spillovers from 22 OECD countries to the Group of 77.

CH's findings have been put under rigorous scrutiny. Engelbrecht (1997) re-examines the sensitivity of CH's results by including the measures of human capital and productivity 'catch-up' and finds that the significant R&D spillovers remain, albeit with a reduced magnitude. Keller (1998) scrutinizes the role of trade patterns in

[1] See, among others, Romer, 1990; Aghion and Howitt, 1998; and Grossman and Helpman, 1991.

[2] It should be noted that, besides international trade, knowledge is internationally diffused through a range of channels such as foreign direct investment, international alliances between firms, migration of scientists and engineers, international collaborative research, conferences and publications etc.

determining the extent of R&D spillovers. He focuses on the weights (actual import shares) used by CH to compute S^f and shows that randomly generated import ratios can lead to similar or even higher international spillovers. He further shows that ignoring the import ratios altogether and assigning equal weights to all trading partners' R&D capital stocks also leads to larger spillover effects than those reported by CH. In a recent paper, however, Coe and Hoffmaister (1999) show that Keller's random weights are technically 'not random' and they suggest alternative randomisations which re-confirm that trade patterns are important for knowledge diffusion. Lichtenberg and van Pottelsberghe (1998) show that CH's weighting scheme biases the measurement of S^f and that their indexation scheme also biases the estimates of spillovers coefficients. They propose an alternative weighing scheme but continue to find significant spillovers albeit with somewhat reduced magnitude.

CH used panel co-integration tests. Unfortunately, at the time of their writing the econometrics of panel co-integration was not fully developed. Kao et al. (1999) re-examine R&D spillovers using CH's data and specifications but addressing the econometrics of panel co-integration tests in a more formal and complete manner. Interestingly, Kao et al. (op. cit) do not find evidence of international spillovers - the effect of S^f on TFP appears insignificant – when they use a dynamic OLS (DOLS) estimator shown to have better power properties. Recently, van Pottelsberghe and Lichtenberg (2001) extended CH's analyses by treating foreign direct investment (FDI) as a channel of technology diffusion. They use only 13 of the CH's 22 sample countries and apply panel co-integration tests due to Pedroni (1999). They find evidence of significant R&D spillovers. To sum up, the general picture emerging from this strand of literature is supportive of positive and significant international R&D spillovers across countries[3].

II. Motivation

Studies reviewed above (CH; Engelbrecht, 1997; Keller, 1998; Lichtenberg and van Pottelsberghe, 1998; van Pottelsberghe and Lichtenberg, 2001; Kao et al., 1999), important though they are, impose two important constraints. First, they inadvertently impose homogeneity of domestic and foreign R&D elasticities of TFP across the G7 countries. Second, productivity gains from knowledge diffusion are considered non-rival. Thus, technology diffusions across countries lead to equivalent productivity gains irrespective of whether the country is a technological leader (e.g. US) or a follower (e.g. Canada). Growth theories, which advocate positive international knowledge spillovers, neglect the issues of technological and industrial rivalry. Unfortunately, technological rivalry is a world reality and knowledge diffusion, in principal, can be positive or negative. [4] If R&D strategy is designed to pre-empt competition then spillovers can be negative. Further, R&D competition may lead to duplicative R&D and resource wastage. Thus, whether international knowledge spillover is indeed positive for all countries irrespective of their stages of technological sophistication is an interesting empirical question, which needs to be examined at country level. Nadiri and Kim (1996) point out, and we concur, that "R&D spillovers are likely to be country specific even for the highly industrialised G7 countries". Such diversity can be attributed to hosts of country-specific factors

[3] Our purpose is to add to the spillover literature initiated by CH. We have, therefore, been restrained in summarising the main papers in this category. Griliches (1992 and 1994), Mohnen (1999) and Mairesse and Sassenon (1991), to name but a few, provide extensive surveys.

[4] A number of high profile rival R&D projects exist. EU's Galileo satellite programme, Eurofighter, the Airbus etc. are examples of competition between Europe and America.

including the heterogeneity of 'social capability' and technological infrastructure (see section IV). Panel tests, which are what all the studies reviewed above are, only provide estimates of 'average' spillover coefficients among a group of countries and are ill-equipped to offer any insight into country-specific spillover elasticities. Time series studies, which analyse these issues at country level, are conspicuously lacking in the R&D field.[5] This paper aims to fill this gap in the literature by offering, among other things, up-to-date country level analyses of knowledge spillovers across G7 countries. We apply tests that are capable of modelling the R&D dynamics both in time series and panel frameworks[6]. In so doing, the paper contributes the literature in following ways.

First, we extend the R&D data set in two directions: (i) the data are extended to 35 years (1965-1999) compared to the 20 years (1971-1990) analysed by other studies reviewed above; and (ii) our data set includes total R&D activity (i.e., total R&D expenditure incurred within a national boundary) as opposed to the business-sector-only R&D used in studies reviewed above. This is important because, besides the business sector, higher education, government and the private non-profit institutions also undertake R&D activities and command a significant share (above one-third) of the total R&D activities. Moreover, the share of non-business sector R&D appears to be even higher historically (see section III). Griliches (1994, p.2) remarked that the advances made in theory and econometric methods will be 'wasted' unless they are applied to the right data set. We hope that the new and extended data

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[5] Our EconLit Bid search under the search word 'R&D Spillovers' scored 141 hits (returns). All empirical papers used panel estimators and none were time series studies.

[6] Nadiri and Kim (1996) address the heterogeneity of R&D spillovers across G7 countries using a multi-product-translog-cost function and report that domestic and foreign R&D elasticities differ considerably across G7 countries. We follow a time series approach and assess these issues in the tradition of Coe and Helpman (1995). This paper complements the work of Nadiri and Kim.

set and the new econometric approaches that we employ to examine the knowledge spillover issues will prove to be a step in the right direction in addressing Griliches' point.

Second, we confront an interesting empirical issue of whether global technology diffusion is beneficial to US. A growing body of empirical literature, which is distinctly different from CH's approach, doubts that it does. International spillovers appear asymmetrical, flowing from large R&D-intensive nations to small and less R&D-intensive nations, but not vice versa (Park, 1995; Mohnen, 1999). The US and Japan trade heavily and Japan is a rich and technologically advanced country yet the bilateral spillovers between US and Japan are either greatly in favour of Japan or unidirectional from US to Japan (Bernstein and Mohnen, 1998). International R&D spillover exists from Canada to Japan but not vice versa (Bernstein and Yan, 1995). Only a few OECD countries (US, Germany and Japan) are major spillover generators (Eaton and Kortum, 1996). Inward FDI and Japanese new plant ('greenfield') investments do not contribute to US skills, nor do the imported inputs appear to upgrade the levels of U.S. productivity (Blonigen and Slaughter, 2001) [7]. These results, based on bilateral spillover analyses and / or disaggregated (micro) data, cast doubt on the thesis that international R&D spillovers favour the US. However, Griliches (1992), points out that macro effects of R&D cannot be directly inferred from micro estimates and the extent of R&D spillover may depend on the level of data aggregation. We address this issue at a wider level by using aggregate data and modelling R&D dynamics at country level. Since the US has been the technological leader of capitalist world since World War II and there exist ample grounds to believe that technological and industrial rivalry exist between the US, the EU and Japan, it is of interest to enquire whether the US benefits or loses when its competitors (G7 partners) accumulate their own R&D stocks.[8]

Third, we follow a novel and robust empirical approach. The long-run relationship between TFP, S^d and S^f is examined by employing Johansen's (1991) multivariate VAR, a well-established method in time series econometrics. This method also addresses the possibility of multiple co-integrating vectors between TFP, S^d and S^f and the validity of normalisation on TFP. The robustness of our results is further assessed by employing the fully modified OLS (FMOLS) estimator (Phillips and Hansen, 1990). An important dimension of our empirical approach is that these estimators are used to perform both time series and panel co-integration tests. Larsson et al. (2001) extend Johansen's multivariate vector error correction model to a panel framework. Pedroni (2000) develops panel co-integrations tests based on FMOLS. Thus, we offer time series results which shed some new light on the diversity of R&D dynamics across G7 countries while our panel results help reconcile our findings with the existing ones. Moreover, Johansen's ML based panel co-integration (rank) tests provide an appealing alternative to the two-step residual based tests applied by van Pottelsberghe and Lichtenberg (2001) and Koa et al. (1999).

Fourth, the issue of the stability of spillover elasticity has attracted considerable interest in the literature. CH address this issue by comparing the time varying elasticity of TFP with respect to S^f across different years (i.e., 1971, 1980 and

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[7] Caves (1996) and Aitken and Harrison (1999) also point out that inward FDI can be costly to the productivity of domestic firms.

[8] A number of points can be put forward as to why international R&D spillover may not be positive for US productivity. Foreigners heavily imitate the US so that the foreign R&D stock that the US faces may not be distinct from her own domestic R&D stock. Hence, it can be argued that the foreign R&D stock, duplicated from the US, does not enhance US productivity. If spillover from the US accrues to its product-market rivals then that may cost the US in terms of productivity loss. Further, the accumulation of R&D by the EU and Japan may gradually replace US investments both at home and abroad and reduce US productivity. For empirical evidence on this latter aspect see van Pottelsberghe and Lichtenberg (2001) and see also Dunning (1994).

1990). These elasticities are derived by multiplying the full sample coefficient of S^f by the import share for 1971, 1980 and 1990. Kao et al. (1999) follow the same approach. On the basis of the magnitude of elasticities thus derived they conclude that the impact of foreign R&D has risen 'substantially' from 1971 to 1980. The problem however is that their approach essentially fixes the coefficient of S^f to its full sample value and only picks up the variations in import shares. Moreover, CH and Kao et al. (1999) do not implement any formal tests of structural stability[9]. van Pottelsberghe and Lichtenberg (2001) conduct standard F tests by splitting the sample between 1971-80 and 1981-90 and report significant structural shifts in international spillover coefficients. Given the non-stationary nature of the data, the validity of standard F tests can be called into question however. In this paper we address this issue through the tests of the stability of cointegrating ranks and cointegrating parameters.

Fifth, our time series results also address some of the concerns surrounding the panel tests. Levine and Zervos (1996, p. 325) state that panel regressions mask important cross-country differences and suffer from 'measurement, statistical, and conceptual problems'. Quah (1993) shows the difficulty associated with the lack of balanced growth paths across countries when pooling the data; Pesaran and Smith (1995) point out the heterogeneity of coefficients across countries. Indeed, we find significant parameter heterogeneity of R&D dynamics across G7 countries (see section IV). Finally, we provide country-specific parameters (spillover elasticities, etc.), which are potentially of more policy significance than the cross-country 'average' parameters.

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[9] CH (pp. 884-5) also test for the stability of parameters by using dummy and trend variables and report instability. However, they do not report the standard errors, which makes it difficult to infer whether these shifts are indeed significant.

To preview our results, tests reveal that the R&D dynamics across G7 countries are heterogeneous and as such data cannot be pooled. This is consistent with the diversity of G7 countries in terms of their economic sizes, openness and R&D intensity and infrastructure shown in section IV. TFP, S^d and S^f are cointegrated and all but Germany could be normalised on TFP. We find extremely diverse elasticities of TFP with respect to S^d and S^f across G7 countries. Most importantly, we find international R&D spillovers to be significantly and robustly negative to the US. Thus, accumulation of R&D by G7 partners hurts US total factor productivity. Further, a comparison of our panel and time series results suggests that panel results indeed conceal the cross-country differences, a concern echoed by many. Formal tests show that cointegrating ranks and parameters are stable for a considerably long period, findings that go against those of CH, Kao et al. (1999) and van Pottelsberghe and Lichtenberg (2001).

The rest of the paper is organised as follows. Section III covers data issues; section IV discusses the issues of heterogeneity; section V discusses model specification and econometric methodology; section VI presents empirical results and compares and contrasts these with existing findings; and section VII summarises and concludes.

III. Data

Our sample consists of G7 countries viz., Canada, France, Germany, Italy, Japan, United Kingdom and the United States. Data frequency is annual and covers a period of 35 years (1965-1999). The data series required for the core analysis of this paper are TFP, S^d and S^f. Details of their construction as well as other relevant data and their sources are given in Appendix A. Figure 1 plots the total factor productivity[10].

Figure 1 about here

France, Italy and Japan show more or less smooth increases in their total factor productivity except for some reductions around 1974-75. Canada's total factor productivity shows a prolonged period of stagnation and/or decline from the early 1970s to mid-1980s and then again in the early periods of 1990s. Germany shows quite a sizeable downturn in total factor productivity after 1990, which may be attributed to her unification. UK productivity shows three episodes of decline: mid-seventies, early eighties and late-eighties that overlap well into the 1990s. US total factor productivity appears quite stagnant for a rather long period extending from mid-sixties to early eighties and shows improvements after 1984. In fact our plot closely mirrors the discussion contained in a voluminous literature about the slowdown in US productivity. Griliches (1994) argues that the decline in US productivity might have started as early as the mid-sixties rather than in the mid 1970s, the aftermath of the first oil price shock, as is widely claimed; and productivity might not have recovered until mid-1980s. Our plot of US total factor productivity echoes Griliches' explanation.

Figure 2 about here

Figure 2 plots domestic R&D capital stocks. Canada, France, Germany, Japan and Italy show rises in their stocks of domestic R&D. The UK's plot is smooth but rather flat indicating a slow rate of accumulation. The US's stock of domestic R&D suffers around the first oil shock and recuperates only after 1984; since then the trend is clearly upward.

[10] All plots are normalised at 1995=1. This is done for the ease of comparison with CH's data set. Our

Figure 3 about here

Figure 3 plots S^f. It is interesting to note that the S^f of Japan and Germany are pretty flat since 1975 while their TFP and S^d are rising. This pattern in data is puzzling given the common belief that Japan, in particular, has increasingly benefited from international R&D spillovers. The US, on the other hand, shows an upward trend in S^f (due to a rise in other countries' S^d), but a flat TFP during most of the sample period. This does raise the question whether the build-up of R&D outside of the US is at all helpful to US productivity. We take up these issues in the empirical section. For the remaining countries S^f and TFP are both trending upwards.

Table 1 about here

In table 1 we report the relative importance of various sectors involved in R&D activities. Business sector R&D is dominant in US and Japan and accounts for 75 and 72 percent respectively of their total national R&D expenditure in 1998; however this ratio is only 54% for Italy and 62% for Canada and France. For other countries, the business sector accounts for about two-thirds of total R&D activities. A comparison of R&D expenditures across the 1980s and 1990s indicates that the share of non-business sector R&D activities might have been quite high historically. This pattern of R&D activities clearly shows the importance of total R&D stocks.

IV. Heterogeneity

The heterogeneity of international R&D spillovers has long been emphasised by technology-gap theorists.[11] They argue that technology or 'know-how' is very much entrenched in the organisational structures of a country and carries a distinct

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data plots appear pretty close to those of CH. However, we do not use indexed data in our econometric analysis.

[11] Ames and Rosenberg (1963); Nelson and Wright (1992); Dosi (1988); Nelson (1993) are to name but a few.

‘national flavour', which makes technology transfer often difficult and costly. Each country is perceived as separate technological entity characterised by its own R&D dynamics and differing ‘social capability’ to absorb international innovations. ‘Social capability’ is defined in terms of technical, industrial, economic, financial and political ability of the country concerned. Abramovitz (1993), for example, argues that the lack of ‘technological congruence’ might have significantly delayed the adoption of US technology by the European countries.

In Table 2 we present some aggregate statistics. It is evident that the US is by far the most dominant country, producing around 46% of total G7-wide GDP, followed by Japan (17%) and Germany (11%). Canada is the smallest in terms of GDP (4.2%). Major disparities also exist in the ownership of R&D capital stock and R&D intensity. The US owns 53% of G7-wide R&D stocks, followed by Japan (16%), Germany (11%), UK (8.2%) and France (7.2%). Canada and Italy own around 3% each. The average R&D intensity has remained lowest in Canada (1.4%) and highest in Japan (2.6%) during the sample period. US R&D intensity has been more or less stable during the last 35 years (2.5% of GDP), whereas Japan shows some increments. The intra-G7 trade-flows show that Canada is the most open (42% of her GDP), followed by the UK (18%), Italy (17%), France (16.7%) and Germany (16.7%). Japan’s trade with her G7 partners is lowest (6.8%) and the US trades the equivalent of 7.3% of GDP. Given this heterogeneity in economic size, openness, stocks and intensity of R&D etc., empirical investigations that lump G7 countries together and constrain the elasticity of TFP with respect to S^d and S^f to be equal across them, raise some doubts.

Formal tests of dynamic heterogeneity of the TFP relationship across G7 countries are conducted as follows. First, we estimate a second order autoregressive and distributed lag model, ADL(2), for level of TFP conditioning on levels of S^d and S^f and test for the equality of parameters across G7 countries. Second, we estimate ADL(2) on growth rates and perform tests of parameter equality. Chow type F tests under the null of parameter equality across G7 countries are reported in Table 3; tests reject the null. Thus, the elasticity of TFP with respect to S^d and S^f across G7 are not homogenous. This contradicts the constraint maintained by CH and others. Finally, we test if error variances across groups are homoskedastic as another measure of dynamic heterogeneity. Both the LM-test and the White-test of group-wise heteroskedasticity are reported. The LM test is equivalent to the LR-test and assumes normality whereas White's test is robust to non-normality. Both tests confirm that error variances across G7 countries are significantly different. Thus, the elasticity of TFP with respect to S^d and S^f as well as their dynamics across G7 countries are significantly different and hence the data set cannot be pooled.

V. Specification and Econometric Methods

Specification

We adopt the behavioural specification of CH, followed by numerous studies cited above, in order to examine the effects of domestic and foreign stocks of R&D on domestic TFP. Their basic econometric specification is:

$$LogTFP_t = \beta_1 + \beta_1^d \log S^d{}_t + \beta_1^f \log S^f{}_t + \varepsilon_t \qquad (1)$$

where TFP is total factor productivity, S^d and S^f are domestic and foreign R&D capital stocks, and β^d and β^f are (unknown) parameters which directly measure the elasticity of TFP with respect to S^d and S^f. Equation (1) states that domestic total factor productivity is a function of domestic and foreign R&D capital stocks. In order to evaluate the role of trade patterns on international R&D spillovers, CH interact the import ratio with the stock of foreign R&D and specify the following equation:

$$LogTFP_t = \beta_2 + \beta_2{}^d \log S^d{}_t + \beta_2{}^f m_t \log S^f{}_t + \varepsilon_t \qquad (2)$$

where 'm_t' is the time-varying import ratio. We estimate the long-run relationship between TFP_t, S^d_t and S^f_t using specifications (1) and (2). The variable, $(m \log S^f)_t$, captures the effect of trade patterns in international R&D spillovers.

Methods

Johansen's (1988) maximum likelihood (ML) method re-parameterises a k-dimensional and p^{th} order vector (X) to a vector error-correction model (VECM):

$$\Delta X_t = \mu + \Gamma_1 \Delta X_{t-1} + \Gamma_2 \Delta X_{t-2} + ... + \Gamma_{p-1} \Delta X_{t-p+1} + \Pi X_{t-p} + \varphi D_t + \varepsilon_t \qquad (3)$$

In our analysis $X_t = [TFP, S^d, S^f]_t$ is a 3x1 vector of the first order integrated [I(1)] variables; Γ_i are (3x3) short-run coefficient matrices; $\Pi_{(3x3)}$ is a matrix of long-run (level) parameters; D_t capture the usual deterministic components; μ is a constant term and ε_t is a vector of Gaussian errors. The steady-state of (3) is given by the rank of Π which is tested by the well known Maximal Eigen-value and Trace tests (Johansen, 1988). Asymptotic critical values for these test statistics are tabulated by Osterwald-Lenum (1992). A co-integrated system, X_t, implies that: (i) $\Pi = \alpha_{(3 \times r)}\beta'_{(r \times 3)}$ is rank deficient, i.e., r< k (r = number of distinct co-integrating vectors); and (ii)$\{\alpha_\perp \Gamma \beta_\perp\}$ has full rank, (k-r), where $\alpha_\perp$ and $\beta_\perp$ are (3 x (3-r)) orthogonal matrices to α and β.

A number of issues are important in the specification and testing of VAR models. The power of co-integration tests depends on the time span of data rather than on the number of observations (Campbell and Perron, 1991) [12]. Our data extend 35 years, which in our view is sufficiently long to capture the long-run relationship between TFP, S^d and S^f. Further, in order to allow for finite samples, degrees of freedom adjustments are suggested by, among others, Reimers (1992) and we adjust the test statistics accordingly [13]. The VAR lengths (p) are specified such that the VAR residuals are rendered non-autocorrelated[14]. Since variables in the VAR have non-zero mean we restrict a constant term in the co-integrating space. Our trivariate VAR can at most have two co-integrating vectors. If multiple co-integrating vectors are found in the system then Johansen (1991) suggests identification through exactly identifying restrictions whereas Pesaran and Shin (2002) suggest tests of over-identifying restrictions. We follow the latter approach if two co-integrating vectors are found. The stock of foreign R&D for each country, a key conditioning variable, is a weighted sum of the rest of the world's (i.e., other G7 countries') domestic R&D. Therefore, S^f may be weakly exogenous to the system. We subject S^f to weak-exogeneity tests and, where found, impose it in further estimations for it improves the efficiency of the estimated co-integrating vectors.

Recently, there have been significant advances in the econometrics of panel cointegration. Larsson et al. (2001) develop a panel co-integration test based on

[12] Hakkio and Rush (1991) point out that, unlike in the univariate tests, shorter sample is acceptable in a multivariate VAR since it yields additional observations on the long-run fluctuations. Luintel and Khan (1999) further elaborate on this issue.

[13] Reinsel and Ahan (1992) and Cheung and Lai (1993) also suggest for (equivalent) degree of freedom correction for small samples.

[14] It is common to specify lag lengths following some information criteria (for example, Akaike, 1973; Schwarz, 1978). However, Johansen (1992) suggest that the lag length in the VAR should be specified such that the VAR residuals are empirically uncorrelated. Cheung and Lai (1993) show that the lag length selection based on information criteria may not be adequate when errors contain moving average terms. Hence, we specify lag-length based on the test of serial correlation in VAR residuals.

Johansen's ML approach. The panel cointegration rank trace test (P_{TR}) is shown to be the standardized mean of individual trace statistics (TR_{VT}) estimated for each member of the panel and is given by:

$$P_{TR} = \frac{\sqrt{V}[TR_{VT} - E(Z_w)]}{\sqrt{\operatorname{var}(Z_w)}} \Rightarrow N(0,1) \qquad (4)$$

where V is the number of countries in the panel, i.e., (i=1,…,V); T is the time dimension (t=1, …,T); $E(Z_w)$ and var (Z_w) are the expected mean and variance of the asymptotic trace statistics which are tabulated by the authors through stochastic simulations. Under the null all the countries of the heterogeneous panel have at most r co-integrating vectors among k variables i.e., H_0: rank (Π_i) = $r_i \leq r$ against the alternative hypothesis H_1: rank (Π) = k. P_{TR} is shown to be asymptotically standard normal. For $T \geq 25$ and $V \geq 5$, the power of this test is near unit. Since we have a panel of V = 7 and T= 35, this test should be sufficiently powerful.

Monte Carlo simulations of Kao and Chiang (1998) show that the pooled (within-dimension) dynamic OLS (DOLS) panel cointegration estimator outperforms those based on OLS and Fully Modified OLS (FMOLS; Phillips and Hansen, 1990). Koa et al. (1999) employ these tests in order to assess the R&D spillovers using CH's model and data. More recently, Pedroni (2000) shows that the group mean (between-dimension) panel cointegration tests based on FMOLS are preferred over the pooled estimators as the former shows relatively minor size distortions even in small samples. In particular, the group mean estimators allow for the heterogeneity of co-integrating vectors across the countries in the panel under the alternative hypothesis whereas the pooled estimators constraint the value of the cointegrating vector to be same for all countries. The null hypothesis is H_0: $\beta_i = \beta_0$ for all i (i = 1,…,V) against the alternative hypothesis H_1: $\beta_i \neq \beta_0$; thus the values for β_i are not constrained to be the same under the alternative hypothesis. There is a clear advantage in using group mean estimators in this application as we expect the elasticities of TFP with respect to S^d and S^f to be heterogeneous across G7 countries. The group mean estimators are robust to heterogeneity of residual dynamics around the cointegrating vectors whereas the pooled panel estimators are not. Further, group mean estimators do not constrain the transitional (short-run) dynamics to be similar across countries and are robust to the fixed effects. Finally, the point estimates of group mean estimators can be interpreted as the mean value of the cointegrating vectors which is however not true in the case of pooled estimators. Pedroni (2000) shows that the group mean panel FMOLS estimator, $\hat{\beta}_{GFM}$, can be constructed as:

$$\hat{\beta}_{GFM} = V^{-1}\sum_{i=1}^{V}\beta_{FM,i} \tag{5}$$

where $\beta_{FM,i}$ is the conventional FMOLS estimator applied to the ith member of the panel. The associated t-statistic for the between dimension estimator is given by

$t_{\hat{\beta}} = V^{-1/2}\sum_{i=1}^{V} t_{\hat{\beta}FM,i}$; where $t_{\hat{\beta}FM,i}$ is the conventional t-ratio associated with the $\beta_{FM,i}$.

The group mean t-statistics are shown to be standard normal and so long as the T > V, which is what we have, this panel test is extremely powerful and exhibits remarkably small size distortion even in a small sample.

VI. Empirical Results

CH (1995) reported that TFP, S^d and S^f are clearly trended and contained unit roots. Plots of our data set in figures 1 to 3 also confirm this trending pattern. Nevertheless, we implement the univariate KPSS test (Kwiatkowski et al., 1992), which tests the null of stationarity, in order to formally evaluate the time series properties of data.[15] Results are reported in Table 4. As expected, in most cases, tests reject the null of stationatity of TFP, S^d and S^f at very high precision (1%). In some cases however the precision is not as high; nevertheless, the null of stationarity is rejected for all but one at 10% or better. The only exception is UK TFP which appears trend stationary; however its level stationarity is clearly rejected. In view of its level non-stationarity and the slowly decaying autocorrelation functions, we decided to treat UK TFP as a non-stationary process. All series are stationary in their first differences. Overall, KPSS tests confirm that TFP, S^d and S^f are unit root processes, a result consistent with earlier findings (e.g., CH).

Johansen rank tests and a range of VAR diagnostics pertaining to specifications (1) and (2) are reported in Table 5. Tests show that S^f is weakly exogenous in five countries (viz., Canada, France, Germany, Japan and the UK) in specification (1) whereas it holds for all but the U.S in specification (2). Where identified, the weak exogeneity of S^f are imposed for they improve the efficiency of the estimates.

Trace statistics, adjusted for the finite samples, show that TFP, S^d and S^f are co-integrated in all sample countries and exhibit a single co-integrating vector under both specifications. [16] For a valid normalisation and error-correction representation, the associated loading factors (α_s) must be negatively signed and significant. On this basis, we could normalise all countries except Germany on TFP; their associated

[15] Kwiatkowski et al. (1992) show that these tests are more powerful than the usual DF/ADF tests. Recently however Caner and Kilian (2001) warn against these power gains specially when data are high frequency. Our data are low frequency.

[16] Under specification (1) Italy shows two cointegrating vectors unless S^f is treated weakly exogeneous. Since the identification of multiple cointegrating vectors in a panel setting is not yet fully developed we circumvent this by imposing weak exogeneity of S^f for Italy in specification (1). In any case, even

loading factors are negatively signed and significant at 5% or better. Germany, on the other hand, showed perversely (positively) signed and insignificant loading factors and hence could not be normalised on TFP. Instead, Germany's co-integrating vector is normalised on S^d, which now has a correctly signed and significant loading factor[17]. LM tests show absence of serial correlation in VAR residuals. Residuals also pass normality tests. A second or third order lag length is sufficient to render the VAR residual uncorrelated. This is plausible in view of the low frequency (annual) of data.

The estimated co-integrating vectors, obtained from the Johansen and FMOLS methods are reported in table 6. Most importantly, we find that the international R&D spillover for the U.S. is significantly negative (-0.16 under Johansen and –0.06 under FMOLS). This result is extremely robust to VAR length (1-4), estimation methods and specifications. Thus, it appears that R&D accumulation by competitors hurts US TFP. This supports our conjecture above and reinforces the findings of Bernstein and Mohnen (1998) and Blonigen and Slaughter (2001) from a macro perspective. Japanese results, on the other hand, are puzzling. International R&D spillovers appear insignificant for Japan in all but one estimate (i.e., FMOLS under specification (1)). We imposed a positive coefficient on Japanese S^f and tested for its sustainability. LR tests do not reject the proposition that the elasticity of TFP with respect to S^f is positive for Japan (at most 0.02 in specification (1)). The remaining countries - Canada, France, Italy and UK - show positive and significant effects of S^f on TFP.

Interacting S^f with the import ratio magnifies the spillover coefficients (semi-elasticities) of all countries but Canada. The negative spillover coefficient for the US becomes –0.30 under the Johansen method and it becomes –0.20 under the FMOLS.

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when weak exogeneity is not imposed the exactly identified Italian TFP relationship appears qualitatively similar to the one reported in table 6.

[17] Normalisation on S^f is conceptually problematic nonetheless it produces an insignificant loading factor (α) for Germany.

Increases in these coefficients are due to the low import ratios and are consistent with the patterns found by CH and others. Japan continues to show insignificant spillover coefficients but it does not statistically reject the null of a positive coefficient (with a maximum of 0.15).

The effect of S^d on TFP is more prevalent. Of the six countries normalised on TFP all but Canada exhibit a positive and significant elasticity of TFP with respect to S^d. For Canada the effect of S^d on TFP appears insignificant, which, again is rather surprising.

Our results vividly show that the estimated point elasticities of S^d, S^f and $m*S^f$ exhibit considerable cross-country heterogeneity. Starting with the Johansen approach, the unrestricted domestic R&D elasticity of TFP ranges from (statistically) zero (Canada) to 0.55 (UK) whereas the foreign R&D elasticity of TFP ranges between –0.16 (US) to 0.10 (France) in specification (1). In specification (2), the coefficient of S^d hovers between (statistically) zero (Canada) to 0.62 (UK) and that of $m*S^f$ between (statistically) zero (Japan) to –0.30 (US). FMOLS shows very similar pattern, which establishes the robustness of our results. This multiplicity in point elasticity across G7 countries is consistent with the heterogeneity of R&D dynamics shown in section IV as well as the arguments that R&D spillovers are likely to be country specific.

The co-integrating vectors for Germany deserve some comment. The Johansen method explicitly allows for the test of the validity of normalisation. Based on this, the German co-integrating vector could only be normalised on S^d. Specification (1) shows significant effects of S^f and TFP on S^d whereas in specification (2) the coefficient of $m*S^f$ becomes insignificant. Thus, in this tri-variate framework, the

R&D spillovers for Germany can be described, at best, as complementarities between S^d and S^f.

Finally, the last row of table 6 reports panel trace statistics and panel group mean cointegrating parameters. Since Germany cannot be normalised on TFP we exclude it from the panel analysis. The panel null H_0: r=0 is strongly rejected at very high precision whereas H_0: $r \leq 1$ is not. Thus, panel trace tests show that six of the G7 countries share a common cointegrating vector. Panel estimates show positive and significant effects of S^d and S^f on the TFP which are consistent with the extant panel tests. Both estimators (Johansen and FMOLS) produce pretty close panel estimates. The point elasticity of TFP with respect to S^d ranges between 0.23 and 0.28 which are also close to those reported by CH (0.22) and Engelbrecht (0.26) but larger than those of Keller (0.13) and van Pottelsberghe and Lichtenberg (0.15). However, our coefficients of foreign R&D (0.03 to 0.06) are smaller than those of other studies. For example, CH's estimates are 0.06 and 0.29 for S^f and $m*S^f$ while Keller's random weight estimates are 0.13 and 0.33, respectively. It is instructive to note that while our panel results are qualitatively similar to those in the literature, country level results are extremely diverse and acutely different from panel results.

In table 7 we report statistical evidence on the extent to which panel estimates correspond to the country specific estimates. Two sets of results are reported. Panel A contains p-values of the LR tests under the null that each individual country specific parameter is equal to its respective panel (between dimension) estimates. Tests (low p-values) show that the panel estimate of the elasticity of TFP with respect to S^d is statistically different from its corresponding country specific point estimates in all but one country (US). Likewise, tests reject the equality of the panel and country specific estimates of the elasticity (semi-elasticity) of TFP with respect to S^f ($m*S^f$) for

France, Italy and the US. Panel B reports the joint test that all country specific estimates of a parameter are equal to the corresponding group mean parameter. This test is in the spirit of Hague et al. (1999). It involves conducting a Wald or LR test for the restriction that each country-specific coefficient is equal to the group mean value and summing up the individual (χ^2) statistics. The test is distributed as (χ^2 (V)). The joint test strongly (at a very high precision) rejects the null that country specific coefficients are equal to their respective group mean values. Thus, statistical evidence suggests that panel estimates do not correspond to country specific estimates. An important insight is that panel tests indeed conceal important cross-country differences and any generalisation of panel results is fraught with the risk of deriving wrong inferences with respect to certain member(s) of the panel. This seems to be true with the majority of countries in this study and the US, in particular, appears distinct from others.

Stability

The stability of cointegrating ranks and parameters are examined following the approach of Hansen and Johansen (1998) which compares the recursively-computed ranks of the $\prod$ matrix with its full sample estimate. If the sub-sample rank of $\prod$ differs significantly from its full sample counterpart, then that implies structural shifts in the co-integrating rank. Likewise, conditional on the identified co-integrating vectors, significantly different sub-sample parameters from their full sample counterparts signify instability of cointegrating parameters. The LR test for these hypotheses is asymptotically χ^2, with kr-r^2 degrees of freedom. Tests are carried out in two settings: (i) allowing both short-run and long-run parameters to vary (*Z-model*); and (ii) short-run parameters are concentrated out and only long-run parameters are allowed to vary (*R-model*).

The stability tests are carried out over a period of 15 years (1985-1999) which leaves the first 20 observations as the initial sample. Figure 4 plots the normalised LR statistics that tests the stability of ranks under specification (1) using the *R-model.*[18] All LR statistics are scaled by the 5% critical value; hence the values greater than unity imply rejection of the null of stability and *vice versa*. In these plots the rank, r, is stable if rank, r-1, is rejected.

Figure 4 about here

The time path of the scaled LR statistics show that the null of non-cointegration (H_0: r=0) is clearly rejected as all plots which test this hypothesis are above unity. The plots that test H_0: r≤1 are always below unity (i.e., less than the 5% critical value) except for the US which shows a short period of rank instability around the 1990s. Thus, all in all, tests reveal that the reported cointegrating ranks are stable.

Figure 5 about here

Figure 5 plots the normalised LR statistics, which test for the stability of co-integrating parameters. Both R- and Z-models show co-integrating parameters to be stable as all recursive plots appear well below the unity threshold [19]. Thus, we find that co-integrating ranks and parameters were remarkably stable over the 15-year period which is in sharp contrast to what CH, Kao et al. (1999) and van Pottelsberghe and Lichtenberg (2001) reported. As pointed out before, it is important to note that CH and Kao et al. do not implement any formal tests of stability and their approach

[18] Results from the Z-model are similar to those of the R-model and hence only R-models are reported. In fact, the R-model is more suitable for testing the stability of the long-run parameters (Hansen and Johansen, 1998).

[19] Minor exceptions to this are that the Z-model indicates significant LR statistics (parameter instability) for Japan and the US in the first year of recursion. These, however, die out quickly.

essentially captures the variability of import ratio alone. van Pottelsberghe and Lichtenberg's standard F test may also be questioned due to the non-stationary nature of data. We address these issues and find that the long-run relationships between TFP, S^d and S^f are remarkably stable[20].

Bilateral and Multilateral Spillover elasticities

The estimates of the bilateral international R&D spillovers based on the aggregate point elasticities of table 6 (specification (1)) are reported in Table 8. Each entry is the estimated elasticity of TFP of country i (reported in columns) with respect to the domestic R&D capital stock of country j (reported in row). These bilateral spillover elasticities are calculated as:

$$\beta^f{}_{ij} = \alpha^f{}_i \frac{m_{ij}}{y_j} . \frac{S^d{}_j}{S^f{}_i} \qquad (6)$$

where $\beta^f{}_{ij}$ is the bilateral spillover elasticity of TFP of country i with respect to the S^d of country j; $\alpha^f{}_i$ is country i's elasticity of TFP with respect to S^f; other variables are as already defined. Table 8 shows that a 1% increase in US R&D would increase Japanese output by 0.018%; however, a 1% rise in Japanese R&D would reduce US output by 0.056%. US R&D shows the highest productivity effect on Canada (0.048%), followed by France (0.022%) and Japan (0.018%). Given the negative elasticity of US TFP with respect to foreign R&D, all bilateral spillover elasticities are negative. The accumulation of R&D by Japan hurts US productivity the most (-0.056%) whereas that by Italy hurts the US the least (-0.006%).

[20] Tests of the stability of co-integrating parameters (β^d and β^f) are also conducted under FMOLS by computing recursive Wald tests over the period of 1985-1999. The null hypotheses are that the recursively computed sub-sample and full-sample parameters are equal. Canada, Italy, Japan and the UK did not reject the null for the whole period, whereas France, Germany and the US show robust stability over 1990s only. Overall, FMOLS corroborates the stability found in the VECM. To conserve space, we do not report these results here, but they are available on request.

The mean international productivity effects of domestic R&D are reported in the last row of Table 8. Calculations show that US R&D has the biggest output effect across other members of G7 (a 1% increase in US R&D increases international output by 0.112%), followed by Germany (0.067%). German R&D appears to enhance the productivity of France and Italy in an important way whereas its effect on Japanese output is almost one-tenth of that of the US. The German cointegrating vector, on the other hand, could not be normalised on TFP.

The mean elasticity of domestic output with respect to foreign R&D is reported in the last column of the table. Our calculations show that a 1% rise in the R&D of other G7 countries would reduce the US output by 0.165%. French output appears to benefit most (0.106%) from the rise in the R&D of fellow G7 countries. Japan's major productivity gains come from the US.

The average own rate of return of domestic R&D shows tremendous variation across G7 countries.[21] The US has the highest rate of return (165%) followed by the UK (140%), Italy (106%), Japan (100%), France (58%) and Canada (-0.23%). van Pottelsberghe and Lichtenberg (2001) estimate the average rates of return for G7 to be 68% which is somewhat lower than our mean estimate of 94% for six of the G7 countries (US excluded) [22]. However, both of these estimates are smaller than those reported by CH for G7 countries (123%).

Now we compare our results with those observed in the existing empirical literature. One of the stylised findings is that international R&D spillovers are positive

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[21] The own rate of return from domestic R&D, $\theta_{jj} = \alpha^d_j \dfrac{y_j}{S^d_j}$, where α^d_j is the elasticity of TFP of country j with respect to its own domestic R&D capital Stock, S^d.

[22] In calculating the mean rate of return we set the Canadian rate of return to zero, as the coefficient is statistically insignificant.

and they are not importantly different across G7 countries (CH, footnote 10). Our country level results vividly show this is not the case. The long-run spillover elasticity differs widely across G7 countries. Moreover, the system approach of Johansen does not allow the German cointegrating vector to be normalised on TFP, thus raising doubt about the ad hoc normalisations followed elsewhere. Further, the heterogeneity of co-integrating parameters suggests that results from pooled regressions should be treated with great caution.

A second stylised finding is that foreign R&D contributes more to the productivity of smaller countries than that of large countries. Our country level results exhibit no such robust pattern. Canada, the smallest in the sample, does not exhibit a noticeably large elasticity of TFP with respect to foreign R&D.

A third observation is that the more open the smaller countries are the more they are likely to benefit from international R&D spillovers. We do not find such a pattern across G7 countries. Canada, the most open in terms of intra-G7 trade, which imports 17% from G7 partners shows the smallest spillover elasticity. On the other hand, the magnitude of spillover elasticity of the US is amongst the biggest, although US is not a small country.

A fourth observation is that the output elasticity of S^d tends to be higher than that of S^f for large countries. This is universally corroborated by our results.

A fifth observation is that the US is the main R&D spillover generator, but a weak receiver. Our results reinforce this. We find the US and Germany to be the main spillover generators and Germany generates almost half that of the US. An interesting finding is that the US not only is a weak receiver but a net loser. 1% rises in US and German R&D increase the output of G7 partners by 0.112% and 0.067%, respectively. However, a 1% rise in the R&D stocks of G7 partners reduces US output by 0.165%. This finding is robust to specifications and estimation methods. We attribute this finding to our empirical approach, which investigates R&D spillovers at country level and allows parameters to vary across countries. This is also consistent with our argument that the US, the technology leader, may loose if competitors become technologically more sophisticated and grab increased world market share.

Finally, it is also commonly observed that Japan benefits from spillovers a lot but she generates a little. Our results go a step further in confirming this. We find that Japanese R&D benefits all members of G7 except the US but only in a marginal way. A 1% rise in Japanese R&D stocks increases the output of all members of G7 except the US by 0.015% but it hurts US output by 0.056%. Thus, the net spillover generation from Japan is negative (-0.041%), whereas Japan benefits mainly from the US. Our finding that the US and Germany are the main spillover generators is consistent with those of Eaton and Kortum (1996); however, our finding on Japan differs from theirs.

VII. **Conclusion and Implications**

Coe and Helpman (1995) and a number of subsequent studies provide empirical evidence in support of significant international R&D spillovers in a panel framework. However, they inadvertently constrain the elasticity of TFP with respect to domestic and foreign R&D stocks to be equal across G7 countries. Equivalent knowledge diffusion across countries also implies that technology diffusion is non-rival. We argue that in view of the profound differences across G7 countries in terms of their economic sizes, openness, R&D stocks and intensity etc. the assumption of homogeneous spillovers is untenable. We further argue that technological and industrial rivalry is a world reality. Concerns over national competitiveness and world market shares encourage countries to pursue aggressive policies of acquiring and maintaining technological leadership and to pre-empt any competitor. The EU's resolve to launch the Galileo satellite in competition with the US Global Positioning System (GPS) is a case in point and several such rival R&D projects are well-known. Thus, knowledge diffusion, in principle, could be positive or negative. In this context, we raise an interesting but hitherto unaddressed empirical question: whether the accumulation of R&D by the US's competitors (i.e. G7 partners of US) is costly for US productivity. US productivity may suffer as a result of the gradual replacement of US investment worldwide and / or a gradual depletion of US world market share following the acquisition and build up of R&D by its competitors.

Empirically we adopt the behavioural specification of CH, as modified by Lichtenberg and van Pottelsberghe (1998), but take forward the empirical analysis by using extended data and new econometric approaches. The R&D data set is extended to 35 years and it encompasses total R&D activity as opposed to the 20 years' business-sector-only R&D data analysed by CH and others. It is shown that non-business-sector activity comprises a significant proportion of total R&D activities.

We follow a novel and robust empirical approach. We employ multivariate VAR (Johansen, 1988) and the fully modified OLS (FMOLS; Phillips and Hansen, 1990) estimators both of which are capable of providing country-by-country time series estimates as well as panel estimates of R&D spillovers. Larsson et al. (2001) extend Johansen's multivariate vector error correction model to a panel setting whereas Pedroni (2000) extends the FMOLS.

Time series and panel tests both show that TFP, S^d and S^f are cointegrated. Thus, long-run relationships between TFP, S^d and S^f exist irrespective of the methods of estimation. Four points stand out. First, our panel results show positive and significant effects of S^d and S^f on TFP, which establish the results of existing panel tests in the literature. Second, country level results reveal that the long-run elasticities of S^d and S^f differ widely across G7 countries. Most importantly, we find significantly negative elasticity of TFP with respect to foreign R&D stocks for the US. This result is robust to VAR lengths, specifications and estimation methods. Thus, our country level results suggest that accumulation of R&D by the rest of the world hurts US productivity. All countries (normalised on TFP) except Canada show significantly positive domestic R&D elasticity of TFP. For Canada the effect of S^d on TFP is insignificant which is puzzling. For Germany S^f appears to complement S^d .

Third, a comparison of time series and panel results reveal that panel tests indeed conceal important cross-country differences and any generalisation of panel results requires utmost care. Inferences drawn from panel results may go contrary to those drawn from individual members of the panel and the US happens to be one of such prime cases in this analysis. We found that, in most cases, the country specific estimates (point elasticities of S^d and S^f) are significantly different from their panel counterparts. It is also shown that data on G7 countries could not be pooled at least in order to analyse the R&D dynamics.

Finally, our results go some way forward reconciling two seemingly conflicting findings. CH and others (Keller, 1998; Lichtenberg and van Pottelsberghe, 1998 etc.) report positive and equivalent R&D spillovers across G7 countries. However, Park (1995) and Mohnen (1999) report international R&D spillovers to be asymmetrical, flowing from large R&D intensive nations to small and less R&D intensive nations. Bernstein and Mohnen (1998) report that R&D spillovers could also to be unidirectional. Our panel results - methodologically close to that of the CH-approach - show that lumping together G7 data and suppressing the country specific heterogeneity and / or estimating mean elasticities in the panel produce positive and equivalent spillovers across G7 countries. However, relaxing such constraints through country-by-country analyses produce multiplicity of R&D dynamics across G7 countries which is closer to the findings of Park (1995), Mohnen (1999), Bernstein and Mohnen (1998) and Blonigen and Slaughter (2001). Further, our results support the findings of Nadiri and Kim (1996) from a different analytical perspective.

The main implications of this study are two fold. First, the extent and the dynamics of knowledge diffusion may differ depending on the stage of technological sophistication of the country concerned. Second, as bilateral spillover elasticities indicate (Table 8), the distribution of knowledge diffusion is hardly uniform. For example, the US is the sole spillover generator for Canada; France and Italy mostly receive knowledge diffusion from Germany; Japan mainly receives from the US, whereas Germany and US both appear equally important for the UK. This indicates bonding between nations that have congruent technology and / or geographical proximity.

Appendix A: Sources and construction of data

The relevant data series and their sources are as follows. Gross domestic product (Y), gross fixed capital formation (I), level of employment (L) and GDP deflator (P) are obtained from OECD's Analytical database; total gross domestic expenditure on research and development (E^{RD}) is obtained from OECD's R&D database; exports (X) and imports (M) of goods and services are obtained from OECD's International Trade and Commodity Statistics (ITCS) database; bilateral exchange rates with US dollars are obtained from International Financial Statistics (IFS) published by the International Monetary Fund.

A consistent series on physical capital stock (K) for the whole sample period was lacking, therefore we constructed total capital stock for each country in the sample from respective gross fixed investment series in constant prices using the perpetual inventory method[23]. A depreciation rate of eight percent and the sample-average real GDP growth rate are used to generate the initial capital stock. Likewise, following common practice (CH, 1995) the domestic R&D capital stock (S^d) is calculated from E^{RD} using the perpetual inventory method. E^{RD} covers all the R&D expenditure carried out within the national territory of each sample country, which is converted to constant prices by deflating by the GDP deflator. The initial domestic R&D capital stock (S_0^d) is calculated as (see CH, 1995):

$$S_0^d = \frac{E_0^R}{g+\delta} \qquad (7)$$

where δ is depreciation rate, assumed to be eight percent,[24] g is the average annual growth rate of E^{RD} over the sample, E^R_0 is the initial value of E^{RD} in the sample. In

[23] Gross fixed investment was converted to constant prices by deflating by the GDP deflator.

[24] Following Coe and Helpmen (1995), we also computed R&D capital stocks using 15% depreciation rate. Our econometric results remain qualitatively same to this alternative measure of S^d.

order to compute the foreign R&D capital stock (S^f) we follow the insights of Lichtenberg and van Pottelsberghe (1998) and compute as:

$$S_i^f = \sum_{j \neq i} \frac{m_{ij} S_j^d}{y_j} \qquad (8)$$

where m_{ij} is imports of goods and services of country i from country j and y_j is country j's GDP.[25] Finally, we compute total factor productivity (TFP) in the usual way (c.f. CH,1995):

$$\log TFP = \log Y - \gamma \log K - (1-\gamma) \log L \qquad (9)$$

Following the literature in this tradition we set the value of the γ coefficient to 0.3.

[25] Note that S^d_j are converted to common currency (US dollars) using PPP equivalent exchange rates while calculating S_i^f.

Appendix

Figure 1: Total factor productivity

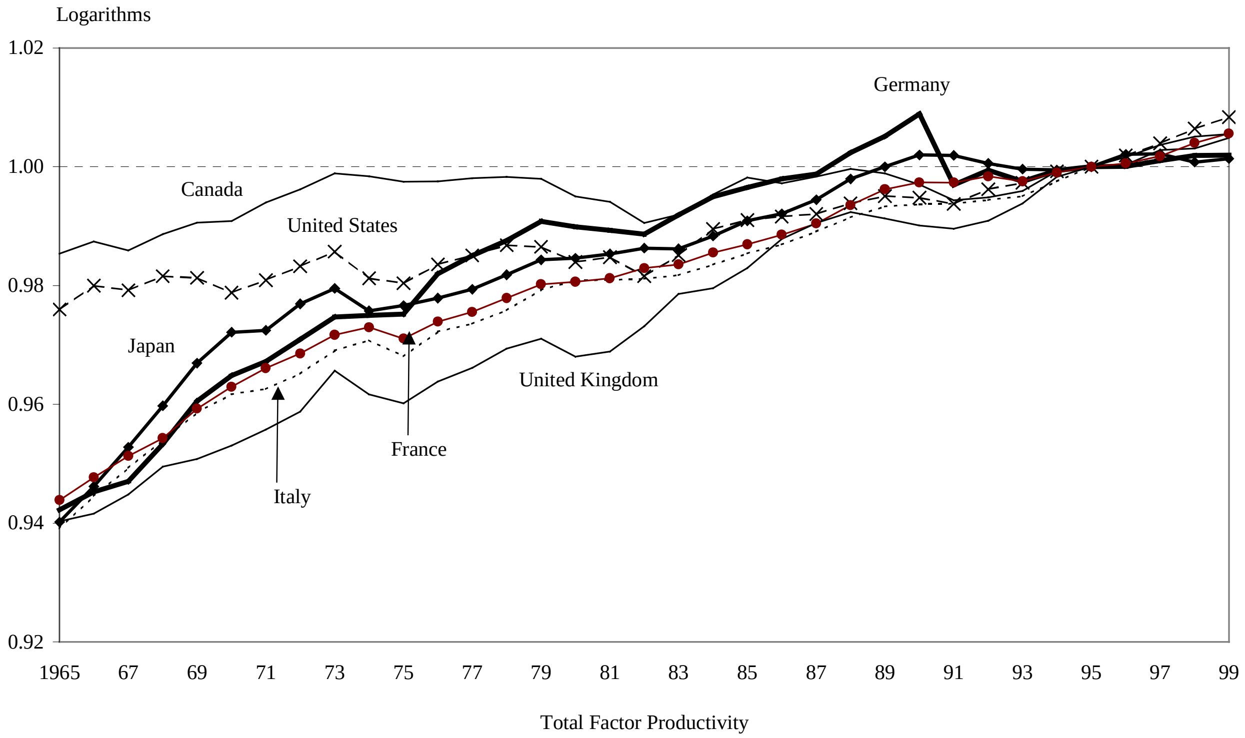


Figure 2: Domestic R&D capital stocks

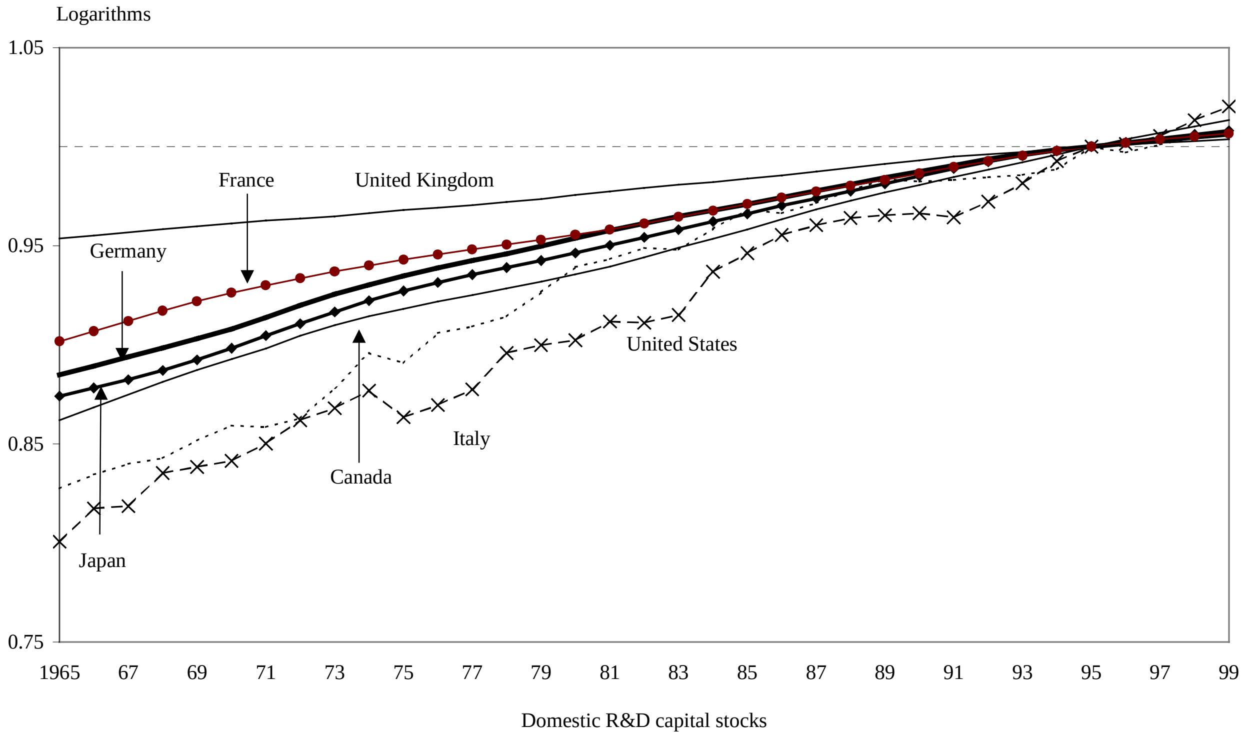


Figure 3: Foreign R&D capital stocks

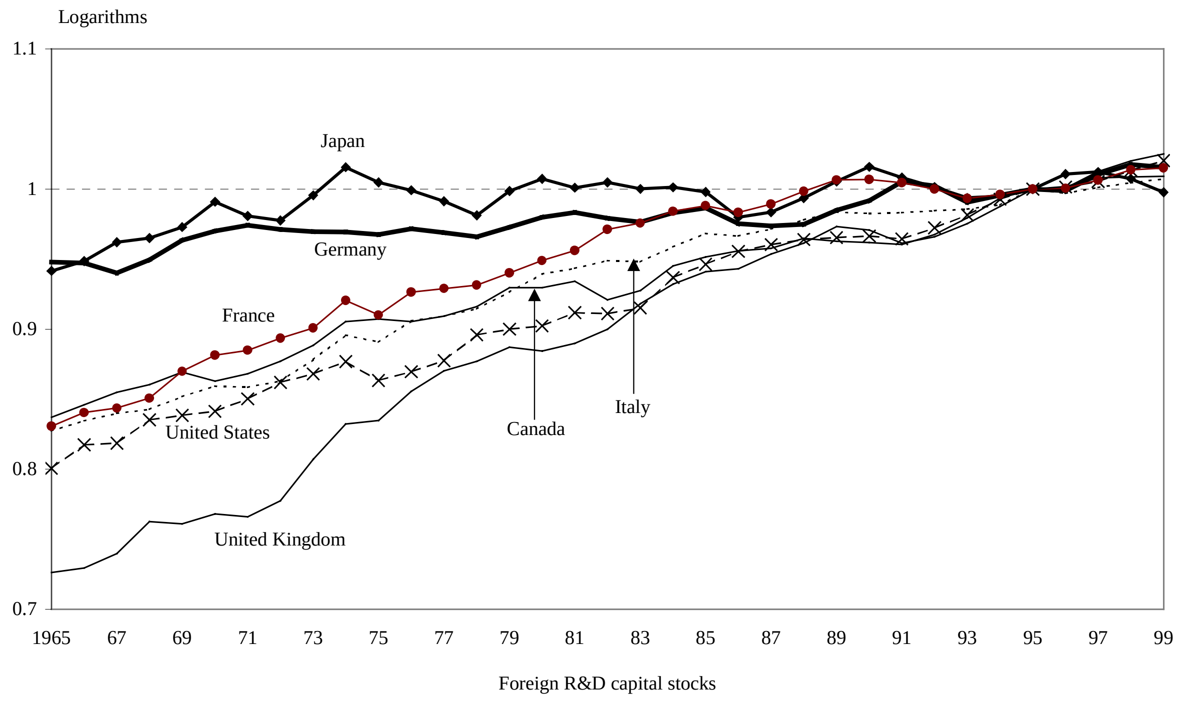
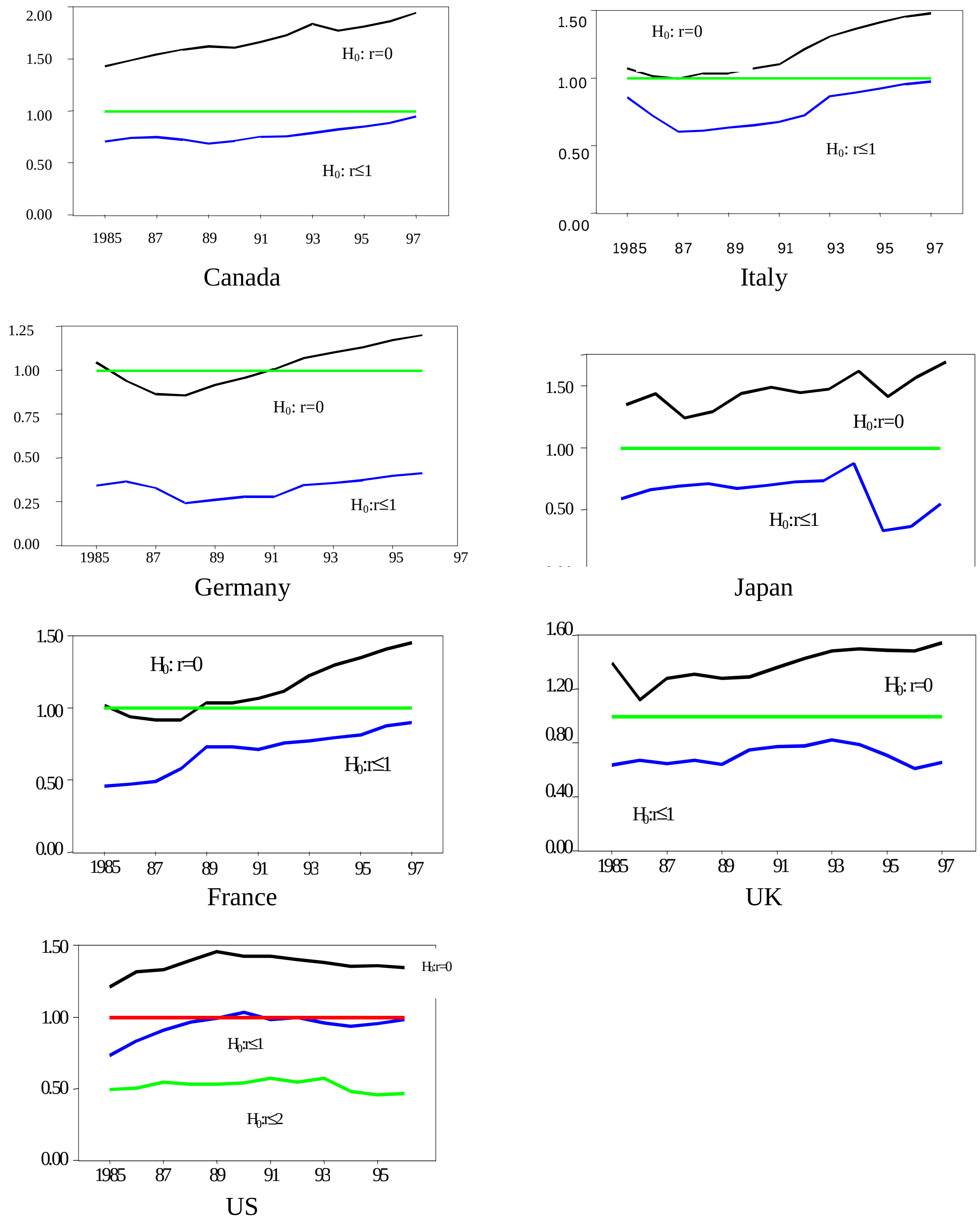
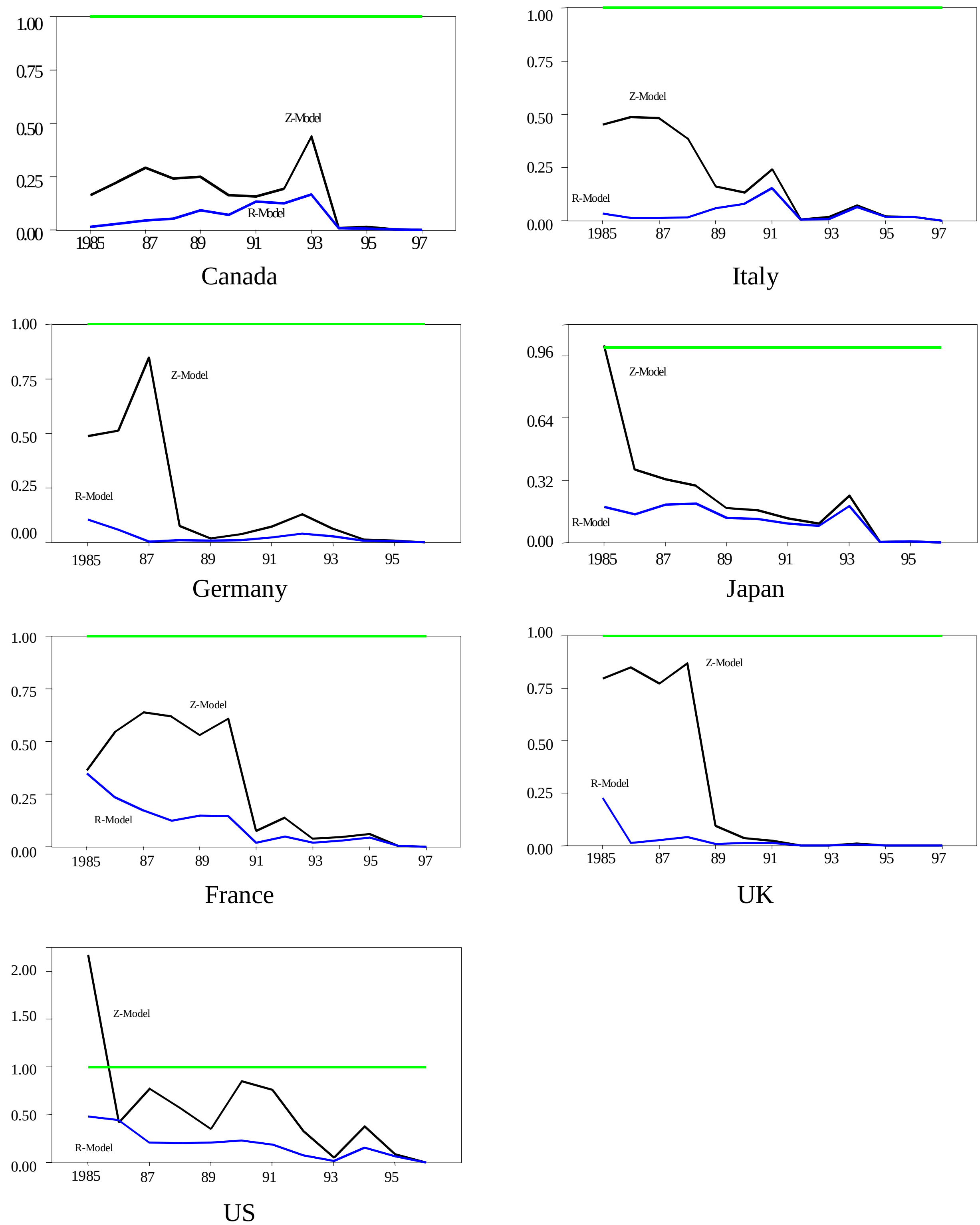


Figure 4: Plots of Scaled Recursive LR-Statistics (Rank Stability Tests)



Recursive LR statistics of rank stability are scaled by appropriate 5% critical value. Hence, plots above unity imply rejection of the null.

Figure 5: Parameter Stability Tests (Plots of Scaled Recursive LR-statistics)



Since the plots of recursive LR-statistics are scaled by the appropriate 5% critical value, the horizontal line at unity indicates the critical threshold.

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Table 1: Distribution of domestic R&D expenditure by sector of performance (%)

	Business enterprise sector						Other sectors					
Countries	1981	1985	1991	1995	1997	1998	1981	1985	1991	1995	1997	1998
Canada	48	53	52	60	61	62	52	47	48	40	39	38
France	59	59	62	61	61	62	41	41	39	39	39	38
Germany	69	72	70	66	68	68	31	28	31	34	33	32
Italy	56	57	56	53	53	54	44	43	44	47	47	46
Japan	66	72	75	70	73	72	34	28	25	30	27	28
United Kingdom	63	64	67	65	65	66	37	36	33	35	35	34
United States	70	73	73	72	74	75	30	27	27	28	26	25

Source: Main Science and Technology Indicators (MSTI) database, OECD.

Table 2: Some stylised aggregate statistics

GDP share in G7 total[a]	Canada	France	Germany	Italy	Japan	UK	US
1965-69	4.0	8.1	11.4	7.5	12.8	8.6	47.6
1970-79	4.2	8.4	11.1	7.7	16.0	7.8	44.7
1980-89	4.4	8.1	10.2	7.7	17.5	7.0	45.1
1990-99	4.2	7.4	10.6	7.1	18.1	6.7	45.8
Mean	4.2	7.9	10.7	7.4	17.0	7.2	45.5
R&D stock share in G7 total[a]							
1965-69	1.8	6.4	8.0	2.2	6.9	12.1	62.5
1970-79	2.1	7.1	9.8	2.7	10.9	10.0	57.5
1980-89	2.4	7.3	11.2	3.0	15.8	8.3	52.0
1990-99	2.6	7.3	11.0	3.4	19.6	6.7	49.5
Mean	2.4	7.2	10.6	3.0	15.8	8.2	52.8
R&D intensity[b]							
1965-69	1.2	2.0	1.7	0.7	1.6	2.3	2.7
1970-79	1.1	1.7	2.1	0.8	2.0	2.2	2.2
1980-89	1.4	2.1	2.6	1.1	2.6	2.2	2.6
1990-99	1.6	2.3	2.4	1.1	2.9	2.0	2.6
Mean	1.4	2.2	2.4	1.1	2.6	2.1	2.5
Trade intensity[c]							
1965-69	27.5	9.0	12.0	11.1	7.0	9.7	3.7
1970-79	32.7	13.6	13.8	16.9	6.8	14.6	5.8
1980-89	36.1	16.7	19.2	17.1	7.7	18.3	7.1
1990-99	47.5	17.4	16.5	17.1	6.2	18.6	8.1
Mean	41.6	16.6	16.7	17.1	6.8	18.0	7.3

a: based on constant price (1995 PPP dollars).
b: R&D expenditures as a percentage of GDP.
c: Exports plus imports from other G7 countries as a percentage of GDP.

Source: MSTI database, OECD

Table 3: Heterogeneity of R&D and TFP dynamics across G7 countries

Panel: A			Panel: B		
Equality of θ	LM test	WH test	Equality of λ	LM test	WH test
6.855[a] F(7,175)	30.528[a] $\chi 2(6)$	28.000[a] $\chi 2(7)$	17.757[a] F(7,182)	27.757[a] $\chi 2(6)$	33.485[a] $\chi 2(7)$

The specification for panel A is: $\Delta tfp = \theta_0 + \sum_{i=1}^{2} \theta_{1i} \Delta tfp_{t-i} + \sum_{i=1}^{2} \theta_{2i} \Delta S^d{}_{t-i} + \sum_{i=1}^{2} \theta_{3i} \Delta S^f{}_{t-i} + \varepsilon_t$.

The specification for Panel B is: $tfp = \lambda_0 + \sum_{i=1}^{2} \lambda_{1i} tfp_{t-i} + \sum_{i=1}^{2} \lambda_{2i} S^d{}_{t-i} + \sum_{i=1}^{2} \lambda_{3i} S^f{}_{t-i} + \varepsilon_t$.

Equality of θ and λ are standard (Chow type) F-tests of parameter equality across G7 countries. Lagrange Multiplier (LM) and White's (WH) tests both reject that error variances are homoscedastic across G7 countries. They are computed by regressing the square of residuals on original regressors, their squares and cross products.

Table 4: KPSS unit root tests

	Canada		France		Germany		Italy		Japan		UK		US	
Variables	η_μ	ι_μ	η_μ	ι_μ	η_μ	ι_μ	η_μ	ι_μ	η_μ	ι_μ	η_μ	ι_μ	η_μ	ι_μ
TFP	0.835[a]	0.130[c]	1.228[a]	0.228[a]	1.106[a]	0.290[a]	1.274[a]	0.264[a]	1.187[a]	0.225[a]	1.245[a]	0.074	1.353[a]	0.134[c]
S^d	1.323[a]	0.222[a]	1.256[a]	0.198[b]	1.242[a]	0.306[a]	1.316[a]	0.211[a]	1.280[a]	0.301[a]	1.263[a]	0.144[b]	1.447[a]	0.130[c]
S^f	1.172[a]	0.233[a]	1.236[a]	0.137[c]	1.103[a]	0.137[c]	1.237[a]	0.261[a]	0.670[b]	0.147[b]	1.225[a]	0.139[c]	1.202[a]	0.286[a]

The Critical Values for η_μ are 0.739 and 0.463 at 1% and 5%; the respective critical values for ι_μ are 0.216 and 0.146. In their first differences all series are stationary. The latter set of results is not reported here to conserve space, however are available on request. Superscripts a, b, and c indicate rejection of the null of stationarity at 1%, 5% and 10%, respectively.

Table 5: Co-integration tests and VAR diagnostics between TFP, S^d and S^f (Johansen Method)

	$LogTFP_t = \beta_1 + \beta_1^{\,d} \log S^d{}_t + \beta_2^{\,f} \log S^f{}_t + \varepsilon_t$ (1)						$LogTFP_t = \beta_2 + \beta_2^{\,d} \log S^d{}_t + \beta_2^{\,f} m_t * \log S^f{}_t + \varepsilon_t$ (2)					
	Trace Statistics: Ho: rank=r r=0 r≤1 r≤2	α: loading factor	Wexo	LM (3)	NOR	LAG	Trace Statistics: Ho: rank=r r=0 r≤1 r≤2	α: loading factors	Wexo	LM (3)	NOR	LAG
CA	37.84[a] 8.85 - [19.96] [9.24]	-0.242[b] (-1.96)	0.519	0.399	0.891	3	31.5[b] 8.79 - [19.96] [9.24]	-0.190[c] (-1.90)	0.081	0.414	0.989	2
FR	26.83[a] 6.59 - [19.96] [9.24]	-0.386[a] (-3.61)	0.963	0.354	0.383	2	29.37[a] 5.772 - [19.96] [9.24]	-0.356[a] (-4.05)	0.362	0.181	0.377	2
DE	20.83[b] 4.06 - [20.0] [9.2]	-0.035[a] (-4.36)	0.873	0.236	0.237	2	19.00[c] 6.002 - [19.96] [9.24]	-0.018[a] (-3.60)	0.311	0.801	0.004[a]	3
IT	30.99[a] 9.13 - [19.96] [9.24]	-0.602[a] (-5.68)	0.001[a]	0.137	0.115	3	35.46[a] 8.49 - [19.96] [9.24]	-0.629[a] (-5.62)	0.424	0.169	0.110	2
JP	21.00[b] 4.37 - [19.96] [9.24]	-0.526[a] (-4.49)	0.090	0.501	0.587	3	21.67[b] 5.25 - [19.96] [9.24]	-0.461[a] (-4.70)	0.228	0.367	0.818	3
UK	24.52[b] 7.56 - [19.96] [9.24]	-0.523[a] (-3.87)	0.236	0.409	0.289	2	23.53[b] 5.381 - [19.96] [9.24]	-0.666[a] (-4.93)	0.372	0.203	0.400	3
US	33.81[c] 15.7 3.24 [34.91][19.96][9.24]	-0.299[b] (-2.05)	0.008[a]	0.832	0.323	3	33.76[b] 14.1 3.12 [34.91][19.96] [9.24]	-0.396[a] (-2.26)	0.003[a]	0.733	0.448	3

The country mnemonics are: CA=Canada; FR=France; DE=Germany; IT=Italy; JP=Japan; UK= United Kingdom; US=United States. Reported trace statistics are adjusted for finite sample following Reimers (1992). [.] are the appropriate 5% critical values of trace statistics. Germany is normalised on S^d. Figures within parenthesis (.) are t-ratios. Wexo are P-values of weak-exogeneity test of S^f which are $\chi^2(1)$ distributed. LM(3) are P-values of third order LM test of serial correlation in VAR residuals. NOR are P-values of Bera-Jarque normality tests of VAR residuals, which are $\chi^2(2)$, distributed. LAG indicates VAR lag lengths. Superscript a, b, and c indicates significance at 1%, 5% and 10% respectively. Under the Johansen approach the UK and US required impulse dummy for first oil shock (1973); Germany required impulse dummy for unification (1990-91) in order to improve normality and/or auto-correlation. These dummies were entered unrestricted. No dummies are entered for the FMOLS estimates reported in (table 6).

Table 6: Estimated cointegrating parameters (Johnsen and FMLOS methods)

	Johansen			FMOLS		
	S^d	S^f	$m*S^f$	S^d	S^f	$m*S^f$
CA	-0.035 (1.094) 0.002 (0.125)	0.055 (1.719)^{b}	 0.036(2.118)^{b}	-0.090 (-1.408) 0.005 (0.254)	0.124 (2.903)^{a}	 0.046 (2.102)^{b}
FR	0.155 (6.739)^{a} 0.220 (14.667)^{a}	0.101(6.313)^{a}	 0.187(6.448)^{a}	0.218 (7.616)^{a} 0.254 (14.977)^{a}	0.071 (3.762)^{a}	 0.149 (4.346)^{a}
DE	S^d = 1.545^{a}TFP + 0.414^{b}S^f; S^d = 0.982^{b} TFP +0.060 $m*S^f$ (6.058) (2.029) (1.889) (0.149)			S^d = 2.306^{a} TFP + 0.962^{a} S^f ; S^d = 3.825^{a} TFP - 0.575 $m*S^f$ (6.946) (4.276) (8.019) (1.226)		
IT	0.174 (4.244)^{a} 0.342 (34.20)^{a}	0.092 (4.00)^{a}	 0.113(6.278)^{a}	0.262 (4.482)^{a} 0.339 (32.296)^{a}	0.069 (2.158)^{b}	 0.150 (5.828)^{a}
JP	0.221 (13.00)^{a} 0.224 (11.789)^{a}	0.024 (0.667)	 0.032(0.294)	0.195 (13.858)^{a} 0.234 (14.280)^{a}	0.138 (3.018)^{a}	 0.063 (0.396)
UK	0.554 (5.711)^{a} 0.619 (12.137)^{a}	0.034(1.619)	 0.077(1.974)^{b}	0.481 (4.501)^{a} 0.539 (10.943)^{a}	0.050 (2.221)^{b}	 0.141(3.648)^{a}
US	0.508 (4.269)^{a} 0.281 (6.690)^{a}	-0.159 (-3.00)^{a}	 -0.295 (-6.854)^{a}	0.339 (3.796)^{a} 0.281 (6.831)^{a}	-0.063 (-1.558)	 -0.197 (-2.014)^{b}
	r=0 r=1 Panel trace (P_{TR}): 6.905^{a} 1.489 TFP = 0.263^{a} S^d +0.025^{a} S^f ; TFP= 0.281^{a} S^d +0.025^{a} $m*S^f$ (14.309) (4.619) (32.493) (4.187)			TFP= 0.233^{a} S^d + 0.065^{a} S^f; TFP =0.276^{a} S^d + 0.059^{a} $m*S^f$ (13.406) (4.773) (32.482) (5.839)		

The country mnemonics are: CA=Canada; FR=France; DE=Germany; IT=Italy; JP=Japan; UK= United Kingdom; US=United States. (.) are respective t-ratios. The mean trace statistics for the panel H_0: r=0 and H_0:r=1 are 27.979 and 8.034; their $E(Z_w)$ are 14.995 and 6.086 and var (Z_w) are 24.995 and 10.535, respectively. Panel coefficients are computed as $\hat{\beta}_{Panel} = V^{-1}\sum_{i=1}^{V}\beta_i$ and the associated t-ratios as: $t_{\hat{\beta},panel} = V^{-1/2}\sum_{i=1}^{V} t_{\hat{\beta}i}$. Panel results exclude Germany. Bartlet window of second order is used for FMOLS.

Table 7: Tests for the heterogeneity of co-integrating parameters across countries (Johansen estimates)

	CA			FR			IT			JP			UK			US		
	S^d	S^f	$m*S^f$	S^d	S^f	$m*S^f$	S^d	S^f	$m*S^f$	S^d	S^f	$m*^f$	S^d	S^f	$m*S^f$	S^d	S^f	$M*S^f$
Panel:A	0.00[a]	0.32		0.00[a]	0.00[a]		0.10[c]	0.04[b]		0.02[b]	0.97		0.04[b]	0.72		0.17	0.05[b]	
	0.00[a]		0.54	0.00[a]		0.00[a]	0.00[a]		0.00[a]	0.00[a]		0.95	0.00[a]		0.22	0.71		0.02[b]

Panel:B		S^d	S^f;	S^d	$m*S^f$
		45.391	20.076	56.526	30.894
	d.f.	(6)	(6)	(6)	(6)
	Critical Value (1%)	16.816	16.816	16.816	16.816

The country mnemonics are: CA=Canada; FR=France; IT=Italy; JP=Japan; UK= United Kingdom; US=United States.
Panel: A reports the p-values of LR tests under the null that each country specific parameter is equal to its panel counterpart. The first and second rows in panel (A) pertain to specifications, $LogTFP_t = \beta_0 + \beta^d \log S^d{}_t + \beta^f \log S^f{}_t + \varepsilon_t$ and $LogTFP_t = \beta_0 + \beta^d \log S^d{}_t + \beta^f m_t \log S^f{}_t + \varepsilon_t$, respectively. Tests are reported only for the slope parameters. Panel B tests the joint null that all country specific parameters associated with a particular variable (e.g., S^d) are equal to the respective panel (group mean) coefficient. Thus, panel A reports country-wise and parameter wise tests of equality whereas panel B reports the joint tests.

Table 8: International output elasticities of domestic R&D capital stocks, 1965-1999

	CA	FR	DE	IT	JP	UK	US	Average
CA	-	0.001	0.001	0.000	0.003	0.003	0.048	0.057
FR	0.001	-	0.039	0.014	0.006	0.025	0.022	0.106
IT	0.001	0.027	0.035	-	0.003	0.016	0.013	0.095
JP	0.001	0.001	0.002	0.001	-	0.002	0.018	0.026
UK	0.001	0.007	0.011	0.003	0.003	-	0.011	0.035
US	-0.046	-0.011	-0.021	-0.006	-0.056	-0.025	-	-0.165
Average	-0.041	0.026	0.067	0.012	-0.041	0.020	0.112	

Bilateral output elasticities are calculated using equation (6) in the text. Their interpretation is as follows. The output elasticity of Japan with respect to US R&D is 0.018.The average figures in the last row show that a 1% increase in Japan's R&D would on average reduce other-G7 output by 0.041%. Likewise, the last column shows that Japan's output will increase by 0.026% following a 1% rise in the domestic R&D of other six members of G7 group.