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Kurt A. Hafner

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International Patent Pattern and Technology Diffusion*

Kurt A. Hafner**

May 2005

Department of Economics; University of Bamberg, Germany

Abstract

The paper focuses on the impact of business related R&D spending on input factor productivity (IFP) using international patent applications as a technology diffusion channel. Considering the relationship between research and productivity, international patent pattern reflects the link between the source (R&D) and the use (IFP). To estimate patent related spillover effects, I use the estimation techniques developed and proposed by Kao and Chiang (1998) in order to deal with nonstationary and cointegration and to obtain reliable coefficients. I find that patent related foreign R&D spillover effects are present and that the impact on labor productivity for Non-G7 countries is higher due to foreign rather than domestic R&D activities.

JEL-Classifaction: C12; C23; O30; O40

Keywords: Productivity, R&D, Technology Diffusion, Nonstationary Panels

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**Graduate School: “Märkte und Sozialräume in Europa”, Lichtenhaidestrasse 11, 96052 Bamberg, Germany. Email: kurt.hafner@sowi.uni-bamberg.de

1. Introduction

Are international R&D spillovers trade related and technology mainly embodied in intermediate goods? While Coe and Helpman (1995) among others confront this question by relating the direction of technology diffusion to bilateral trade shares, Keller (1998) shows by Monte Carlo Simulations that randomly created bilateral trade patterns explain more of the variation in total factor productivity (TFP) than those empirically observed. Additionally, the use of appropriate estimation techniques is crucial to deal with time-trended variables and to avoid spurious regression results. In applying a more sophisticated estimation technique on the data set of Coe and Helpman (1995) and through re-examining their econometric findings, Kao, Chiang and Chen (1999) confirm the impact of domestic R&D expenditure on TFP but reject the influence of foreign R&D expenditure weighted by bilateral trade shares. While there might be a theoretical consensus about trade related spillover effects and the importance of a country's openness to trade, empirically it seems to be difficult to quantify the extent and direction of technology diffusion by international trade.

The same applies to the second strand discussed by the literature in considering foreign direct investments (FDI) as an adequate channel for technology diffusion through the relationship of multinational parents and their subsidiaries abroad. Following Keller (2004), such subsidiaries might pick up new technologies from their host countries (outward FDI technology sourcing) or provide technology to domestic firms (inward FDI technology transfer). Again, the evidence is not straightforward as Xu and Wang (2000) mentioned and the impact of technology transfer either from or to host countries still needs to be examined. Case studies, such as the paper by Larrain, Lopez-Calva and Rodriguez-Claré (2000) analyzing the impact of Intel's FDI into Costa Rica in the 1990s, may offer some fruitful insights on how to determine firm specific technology transfer. However, particular case studies do not overcome the lack of quantitative evidence and general understanding.

I propose to use international patent pattern in order to analyze technology diffusion and follow the argumentation of Eaton and Kortum (1999): "*we think that patenting abroad is a much more direct, albeit imperfect, indicator of where ideas are going*". The idea is that patenting domestic research efforts abroad determines the transfer of technology. Firms of the host countries may take legal advantage of the transferred foreign knowledge by paying royalties. Adding such knowledge to a country's own research activities, even in the case of limited domestic R&D spending, impacts on the use and on the efficiency of input factors are likely to result and therefore yield to an increase of input factor productivity (IFP). Considering the

relationship between research and productivity, international patent patterns reflect the link between the source (R&D) and the use (IFP). Hence, international spillover effects may be patent related.

A patent holder receives a temporary legal monopoly at the cost of public disclosure of the underlying technical information. In order to protect themselves from imitators, inventors have to patent their innovations at home and abroad. The inventor's choice is to relate the costs of filing a patent application and of technical disclosure to the likelihood of imitation and the monopoly rents in specific markets. As a result, only the best and most valuable innovations are patented. However, patent figures show that most of the patents are filed at home rather than abroad. This might be the result of either technological immobility or less foreign protection as mentioned by Eaton and Kortum (1999). Given the tight distribution of productivity levels across countries in relation to the skewness of domestic research activity, Eaton and Kortum reject technology immobility and point to a lesser protection provided by foreign patents. Moreover, international patent statistics by the World Intellectual Property Organization (WIPO) and OECD provide only count numbers, whereas information about the value and importance of patents are not given. This may not measure technology diffusion properly because some patents are more valuable and their impact may differ between countries. Hence, using patent count data may serve mainly to determine the direction of international technology diffusion.

The bulk of foreign patent applications are filed and received by the five leading research nations: United States, Japan, Germany, Great Britain and France. The United States is the dominating source of foreign patents followed by Germany and Japan. Concurrently, the same pattern is observed regarding domestic business related R&D (BERD) spending. Thus, we expect higher foreign technology spillovers to smaller and/or less advanced countries from the five leading nations than vice versa in order to explain the small variation in productivity levels across different countries.

To sum up, we examine the effects of domestic and foreign business related R&D expenditure on IFP by the use of patent patterns as the technology diffusion channel. The structure of the paper is described as follows. In the next section, we review the underlying theoretical framework and introduce patent related technology diffusion channels. In section 3 we turn to a brief discussion of the pooled data and its use for estimating technology diffusion. We analyze nonstationary issues in section 4. The results of the testing procedures as well as of the empirical estimations are given in section 5 to show the direction and extent of patent related spillover effects. Section 6 concludes.

2. Theory

Consider the following aggregated production function:

$$Y = A * F(K, L), \quad (1)$$

where Y is aggregate output, K as capital and L as workforce are input factors respectively and A represents technical change. There are two ways to achieve output growth: first to augment the use of input factors by higher capital investments and labor efforts and second to increase the efficiency of input factors and therefore A . Coe and Helpman (1995) regard output growth as driven by innovation in the production of intermediate goods based on the Grossman and Helpman (1991) model. In a simple form, final output Y is produced by an aggregate of intermediate inputs which itself is the result of the use of primary input factors and research activity. Intermediate inputs can be either horizontally differentiated which leads to output growth proportional to the measure of available intermediate goods, or vertically differentiated which affects input productivity by their different qualities. In both cases, aggregate output increases with the usage of intermediate goods. Thus, the part of output growth, which is not attributable to the accumulation of primary inputs, is due to the R&D investment in the intermediate goods production. Therefore, international trade with intermediate goods provides countries with technology embedded goods and creates access to foreign technology knowledge.

Because of the mixed empirical results when analyzing trade related spillover effects, I analyze the impact of an increase in efficiency due to international patent application and choose the second method. In this case, an increase of R&D investments augments the efficiency of input factors used in final output production. As mentioned, foreign patent application transfers technology at the cost of public disclosure. In addition to domestic research, countries gain access to foreign financed know how in using and/or modifying legally foreign patented technology. This leads to an increase of national input productivity and output growth.

Following Coe and Helpman (1995), we define TFP using a Cobb-Douglas functional form as:

$$TFP = Y / [K^\beta L^{(1-\beta)}] = A, \quad (2)$$

where Y is value-added in the business sector, K and L are capital stock and workforce in the business sector respectively, A is technical change and, as usual, β is the production elasticity of capital. Taking the log of equation (2) leads to: $\log TFP = \log Y - \beta \log K - (1 - \beta) \log L = \log A$. However, we try to explain productivity variation of single rather than total input factors through technical change. Taking into account the time and cross section dimension, the estimation equation for IFP is:

$$\log IFP_{i,t} = \alpha_i + \alpha^d \log S_{i,t}^d + \alpha^f \log S_{i,t}^f + \varepsilon_{i,t} \quad i = 1, \dots, N \text{ and } t = 1, \dots, T \quad (3)$$

where i is country and t is time index, $S_{i,t}^d$ represents domestic R&D capital stock, $S_{i,t}^f$ is defined as the patent weighted average of domestic R&D capital stocks from abroad and $\varepsilon_{i,t}$ is the error term. Thus, we obtain for $S_{i,t}^f$:

$$S_{i,t}^f = \sum_{j \neq i} a_{ji,t} S_{j,t}^d \quad j = 1, \dots, N \quad (4)$$

with $a_{ji,t} = b_{ji,t} / \sum_{j \neq i} b_{ji,t}$ as the weight of foreign patent application of country j in country i to total foreign patent application of country i . In equation (3) we mainly modeled the direction of foreign technology transfer by international patent count data. Now, we should also take into account the intensity of technology diffusion and rewrite equation (3) to:

$$\log IFP_{i,t} = \alpha_i + \alpha^d \log S_{i,t}^d + \alpha^f m_{i,t} \log S_{i,t}^f + \varepsilon_{i,t}, \quad (5)$$

with $m_{i,t}$ as an appropriate weight to express technology intensity. In principle, there are two possible weights. First, the use of patent related foreign technology should be more efficient in countries with own research activity and higher domestic R&D spending. If we relate business related R&D expenditure, $R \& D_{i,t}$, to domestic GDP, $Y_{i,t}$, we obtain for the technology intensity:

$$m_{i,t} = R \& D_{i,t} / Y_{i,t}. \quad (6)$$

However, equation (6) aggravates the business cycle problem, which might be inherent to patent data, even further as domestic GDP is now in the denominator. Second, the higher the ratio of non-resident to total patent applications within a country, the more important is the

transfer of foreign technology. For this reason, relating foreign patent applications to total patent applications we write technology intensity as:

$$m_{i,t} = \sum_{j \neq i} b_{ji,t} / \sum_j b_{ji,t}, \quad (7)$$

with $\sum_j b_{ji,t}$ as the sum of total patent applications in country i.

3. Data

We are interested in the change of IFP due to the impact of international technology diffusion. Owing to the more reliable data on labor input and to a lack of data for an adequate stock of business sector capital either for distinctive countries or for a specific period we focus on labor factor productivity (LFP). For this reason, we specify equation (3) and equation (5) by $IFP = LFP$. Calculating productivity figures, there is a qualitative difference between the usage of the number of persons engaged and the number of hours actually worked as labor input factors. As proposed in the literature, we rely on the latter. The figures on labor productivity per hour worked in constant US\$ (PPP) are taken from the *Total Economy Database* provided by the Groningen Growth and Development Center.

Data on business related R&D expenditure have been published by the OECD since about 1965 (*R&D and TBP Database*) for some countries: mainly for the G7 countries as well as for Switzerland. In order to get a complete (balanced) data set for all OECD countries from the beginning of 1965, one is left with the task of estimating the missing R&D expenditure figures. Coe and Helpman (1995) estimated such missing figures by relating real R&D expenditure to real output and investment.¹ However, we limit our observation period from 1981 to 2001 and to 18 OECD countries² partly because of the lack of R&D data as well as the restriction of adequate patent numbers. Converting R&D expenditure flows into R&D capital stocks we use the perpetual inventory method and follow the procedure suggested by Griliches (1979) for calculating the R&D benchmark capital stock for each country. The R&D expenditure data are from the *OECD Main Science and Technology Database*.

The OECD also has been publishing patent figures since the early 1980s. As discussed, we are interested in country specific patent data and in an international pattern of foreign patent applications: i.e. domestic innovations seeking protection abroad by foreign patent appli-

¹ The reader is referred to the cited paper for further details and discussions.

² The 18 OECD countries are respectively: Australia, Belgium, Canada, Denmark, Finland, France, Germany, Greece, Iceland, Ireland, Italy, Japan, Netherlands, Norway, Spain, Sweden, United Kingdom and USA.

cations. Such differentiated data are not provided by the OECD. We therefore rely on the patent statistics published by the WIPO. Patent data have been collecting by the WIPO for more than 150 years (for at least some countries) and – in addition – for poorer and less developed countries. Moreover, since 1975 the WIPO has published figures on foreign patent application and grants broken down by and for each country in the *Industrial Property Statistics Publication B Part I*. We use this data to calculate the weights of foreign patent applications in equation (4) as well as the intensity measures in equation (7).

Indexation and Aggregation Bias

We follow Coe and Helpman (1995) in using indexed figures (1995=1) for LFP. However, we do not calculate the right hand side regressors as indexed data due to the indexation bias mentioned by Lichtenberg and Van Pottelsberghe (1998) when criticizing the Coe and Helpman (1995) approach. Instead, we use levels to express the R&D variables. Moreover, technology diffusion weights calculated by foreign patent applications sum up to one and might indeed have an aggregation bias: the more the foreign patent applications in a single country are, the higher the foreign R&D capital stock is. In this context, a merger between countries would always increase the foreign R&D capital stock. An alternative approach – as proposed by Lichtenberg and Van Pottelsberghe (1998) for the case of trade related spillover effects – would be a ratio of foreign patent application to foreign GDP instead:

$$S_{i,t}^f = \sum_{j \neq i} \frac{b_{j,i,t}}{Y_{j,t}} S_{j,t}^d \quad (8)$$

This formulation would reflect the intensity as well as the direction of ideas. However, equation (8) will also aggravate the problem with the business cycle since foreign output is in the denominator on the right-hand side of the equation (3). We therefore do not change our foreign R&D capital stock and rely on equation (4).

4. Nonstationary Panels

In general, productivity and R&D expenditure data exhibit a clear trend and unit root tests indicate that both sides of the regression are nonstationary, whereas the error term may or may not be stationary. If the error term is not stationary spurious estimations result. In order to avoid spurious correlation among the variables but without differencing and discarding long-run information in the level terms, I follow Kao, Chiang and Chen (1999) as well as Kao

and Chiang (1998) and adopt their proposed techniques as in Funk (2001). As Coe and Helpman (1995) have pointed out, the advantage of not transforming the variables in differences but of relying on level terms is to make use of the embedded information about common trends and long-run equilibrium properties.

Unit Roots and Cointegration Tests

In testing pooled balanced data for unit roots one usually comes across with the Levin, Lin Chu (2002) (LLC) tests, Im, Pesaran & Shin (2003) (IPS) tests and/or the residual-based Lagrange multiplier test by Hadri (2000) (LMH).³ Consider the following determination of variable y :

$$y_{i,t} = \rho_i y_{i,t-1} + \gamma_i z_{i,t} + r_{i,t} \quad i = 1, \dots, N \text{ and } t = 1, \dots, T, \quad (9)$$

where ρ_i is an autoregressive term with lag 1, $z_{i,t}$ is the deterministic component and $r_{i,t}$ is the error term. The deterministic component $z_{i,t}$ could be zero, one, units and/or time effects.⁴

The LLC test assumes that each autoregressive (AR) coefficient is the same for all units, $\rho_i = \rho$, that the error term $r_{i,t}$ is a stationary process and that units are independent across sections.⁵ LLC proposes three different specifications of the deterministic component $z_{i,t}$ and therefore of the series $\{y_{i,t}\}$: (1) $y_{i,t}$ in equation (9) has no deterministic component ($z_{i,t} = 0$), (2) has an individual-specific component ($z_{i,t} = 1$), or (3) has an individual-specific component as well as a time trend ($z_{i,t} = (1, t)$). Testing for unit roots, the LLC test proposes a null hypothesis $H_0 : \rho = 1$ of a unit root (i.e. nonstationary) against the alternative hypothesis that all individual series in the panel data are stationary, $H_1 : \rho < 1$.

Relaxing the restrictive assumption of a homogeneous ρ across units assumed by the LLC tests, the IPS test allows for heterogeneous autoregressive coefficients: ρ_i . The general IPS setting is based on averaging individual unit roots test statistics and assumes that the error

term is serially correlated across cross-sectional units: $r_{i,t} = \sum_{j=1}^{\rho_i} \phi_{ij} r_{i,t-j} + \varepsilon_{i,t}$, whereas $r_{i,t}$ is IID $(0, \sigma_r^2)$ and $\varepsilon_{i,t}$ is a stationary process. Moreover, there might be an individual-specific inter-

³ For the case of pooled unbalanced data, Maddala and Wu (1999) and Choi (2001) modified a Fisher type test.

⁴ Please refer to the cited papers or to Baltagi (2001), who provides an excellent overview for panel as well as cointegration issues.

⁵ To allow for a limited degree of dependence across units, cross sectional averages are subtracted from the observed data without affecting the limit distribution of the panel unit root test, see Levin, Lin and Chu (2002).

ception ($z_{i,t} = 1$) or, in addition, a time trend ($z_{i,t} = (1, t)$). The IPS testing procedure examines the null hypothesis $H_0 : \rho_i = 1$ of each series has a unit root against the alternative hypothesis $H_1 : \rho_i < 1$ of at least one individual series in the panel is stationary.

Finally, LMH limits the determination of $y_{i,t}$ in equation (9) to a random walk of part of the error term $r_{i,t} = \sum_{j=1}^t u_{i,j} + \varepsilon_{i,t}$, with $u_{i,t} \sim \text{IID}(0, \sigma_u^2)$ and $\varepsilon_{i,t}$ as a stationary process, and to a deterministic component $z_{i,t}$, which could be one or a time trend. Note, there is no lagged autoregressive term of $y_{i,t}$ in equation (9): $\rho_i = 0$. Moreover, the stationary error process $\varepsilon_{i,t}$ is assumed to be either homogenous ($\varepsilon_t \sim \text{IID}(0, \sigma_\varepsilon^2)$) or heterogeneous ($\varepsilon_{i,t} \sim \text{IID}(0, \sigma_{i,\varepsilon}^2)$) across units or – relaxing the assumption of being IID – to be serial correlated. The LMH test assumes that each time series is level or trend stationary ($H_0 : \text{stationary}$) against the alternative hypothesis of a unit root in panel data ($H_1 : \text{nonstationary}$).

All the discussed test procedures have in common that their adjusted test statistics asymptotically obey the standard normal distribution. However, LLC and IPS require that $N \rightarrow \infty$ such that $N/T \rightarrow 0$. As a result, in finite samples there are size distortions if N is small or N is large relative to T. Moreover, both tests suffer a dramatic loss of power if time trends are included. Given the fact that classical hypothesis testing ensures that the null hypothesis is accepted unless there is no strong evidence against it, we try to overcome this lack of power by testing both nonstationary as well as stationary for the null hypothesis.

However, having confirmed that variables are non-stationary and exhibit unit roots, a regression containing all variables is cointegrated, if the remaining error term is stationary. Hence, there is a long-run relation and a common trend binding all variables, which leads to steady state equilibrium. If the error term is not stationary, the estimated relationship is spurious and no long-run relationship between the variables will exist.

In order to test for the long-run cointegration relationship (i.e. stationary of the error term), one can either use the resulting error terms from the error correction (EC) model or the proposed cointegration tests presented by Kao (1999), McCoskey and Kao (1998) and Pedroni (1995). Turning to the EC model, the first step is to estimate long-run equilibrium values in levels by removing units as well as time effects (transformation for a two-way fixed effects model). The resulting residuals (i.e. error correction term) are used in the second step to estimate the EC model, which is the first difference of the long-run values augmented by the lagged error correction term as well as the lagged differentiated endogenous variable. The t-

statistic of the lagged error correction term now indicates whether it is significantly different than zero which would mean that a cointegration relationship amongst the variables exists. Another option is to extract the residual term from the EC model and to apply one of the described unit root tests. Alternatively to the error correction procedure, cointegration tests can be used as proposed by the literature. Such tests analyze either the null hypothesis of no cointegration as the Dickey-Fuller and the augmented Dickey-Fuller type tests proposed by Kao (1999) as well as the Phillips and Perron type tests of Pedroni (1995) do or the null hypothesis of cointegration as the residual-based Lagrange Multiplier test by McCoskey and Kao (1998)⁶ does. All tests have in common that residuals are derived by estimating the cointegration variables. However, only for tests presented by Kao (1999) and Pedroni (1995) can residuals be derived from OLS estimation whereas an efficient estimation technique other than OLS is needed for those of McCoskey and Kao (1998).

Estimation Techniques: panel, fully modified and dynamic OLS

The presence of cointegration and unit roots considerably affects the asymptotic distributions in time series as well as in panel analysis. However, cointegration equations have attractive properties: as the number of observations increase in T and N, the OLS estimation of the cointegrated variables converges in the long-run equilibrium to the true value. Nevertheless, for moderate sample size, the estimation bias may remain substantial. Kao and Chiang (1998) found the following limiting distribution: while the OLS estimator is normal distributed with non-zero mean, the fully modified (FM) and dynamic (D) OLS estimators are asymptotically normal with zero mean. They find that the OLS estimator has a non-negligible bias in finite samples and that the DOLS estimator performs better in estimating the panel equations than does the OLS estimator with bias correction or the FM-OLS estimator. As a result, they propose using the DOLS estimator when dealing with cointegration and unit roots.

Consider the following fixed effects panel regression as outlined in Kao, Chiang and Chen (1999):⁷

$$y_{i,t} = \alpha_i + x'_{i,t}\beta + r_{i,t} \quad i = 1, \dots, N \text{ and } t = 1, \dots, T, \quad (10)$$

⁶ Where economic theory predicts a long-run equilibrium, tests based on the null hypothesis of cointegration seems to perform better. Moreover, the asymptotics of the null of cointegration depend only on the cointegration relationship and not on spurious regression properties, see for example Baltagi (2001).

⁷ For further information as well as for analytical derivation see Kao and Chiang (1998).

where $r_{i,t} = u_{i,t}$ is a stationary process and $x_{i,t} = x_{i,t-1} + \varepsilon_{i,t}$. The remaining part of the error term $\varepsilon_{i,t}$ is IID with $(0, \sigma_\varepsilon^2)$. Note that both sides of the regression (10) are assumed to be independent across cross-sectional units. Thus, equation (10) describes a panel regression of cointegration and unit roots, where $y_{i,t}$ is cointegrated in $x_{i,t}$, which is determined by its lagged value. As usual, the ordinary OLS estimator $\hat{\beta}_{OLS}$ is:

$$\hat{\beta}_{OLS} = \left[\sum_{i=1}^N \sum_{t=1}^T (x_{i,t} - \bar{x}_i)(x_{i,t} - \bar{x}_i) \right]^{-1} \left[\sum_{i=1}^N \sum_{t=1}^T (x_{i,t} - \bar{x}_i)(y_{i,t} - \bar{y}_i) \right]. \quad (11)$$

Moreover, the cointegration literature does not assume strict exogenous regressors. There might also be feedback from $y_{i,t}$ to $x_{i,t}$. As a result, the $\hat{\beta}_{OLS}$ estimator is driven by endogeneity as well as by serial correlation.

The FM-OLS estimator corrects for endogeneity by modifying variable $y_{i,t}$ and for serial correlation by adjusting \bar{y}_i in equation (11). This leads to:

$$\hat{\beta}_{FM} = \left[\sum_{i=1}^N \sum_{t=1}^T (x_{i,t} - \bar{x}_i)(x_{i,t} - \bar{x}_i) \right]^{-1} \left[\sum_{i=1}^N \sum_{t=1}^T (x_{i,t} - \bar{x}_i) \hat{y}_{i,t}^+ - T \hat{\Delta}_{\varepsilon\mu}^+ \right], \quad (12)$$

with $\hat{y}_{i,t}^+$ as the endogeneity and $\hat{\Delta}_{\varepsilon\mu}^+$ as the serial correction.

The DOLS estimator however considers leads and lags in order to account for endogeneity and serial correlation. In using the past and future values of the differentiated $x_{i,t}$ as additional regressors, we obtain the following panel regression:

$$y_{i,t} = \alpha_i + x'_{i,t} \beta + \sum_{j=-q_1}^{q_2} c_{ij} \Delta x_{i,t+j} + r_{i,t}, \quad i = 1, \dots, N \text{ and } t = 1, \dots, T, \quad (13)$$

with $r_{i,t} = u_{i,t}$ as the stationary process with zero mean. By running panel regression (13), we get the $\hat{\beta}_{DOLS}$ estimator.

To compare, we first estimate our equations by panel-OLS and, extracting unreported country specific effects, by fixed effects and random effects. Due to the discussed estimation bias, we do not report any of the coefficient signification tests and levels. However when testing for cointegration, we list test statistics based on the residuals of the EC model as well as on the testing procedure of Pedroni (1995). Secondly, we use the estimation techniques pro-

posed by Kao and Chiang (1998)⁸ to estimate reliable coefficients for home and foreign R&D business capital stocks.

5. Empirical Results

To start with, we have to confirm that our data exhibit unit roots and follow a nonstationary path. The null hypothesis testing is nonstationary for LLC and IPS and stationary for LMH.

Turning to the null hypothesis of nonstationary and assuming two different lags, Table 1a shows test statistics and p-values from LLC and IPS for (1) an individual constant trend and (2) a time- and individual constant trend.

Table 1a: Unit Root Tests by Levin, Lin and Chu (2002) and Im, Peseran and Shin (2003)
(Annual data for 18 countries 1981-2001; Observations: 342 (Lag:1); Observations: 324 (Lag:2))

| | LLC (Lag:1) | LLC (Lag:2) | IPS (Lag:1) | IPS(Lag:2) |
|-----------------------------|----------------|----------------|----------------|----------------|
| Constants | | | | |
| log LFP | -0.844 (0.199) | -0.375 (0.354) | 4.625 (1) | 4.225 (1) |
| log S^d | -1.282 (0.1) | 0.495 (0.69) | 2.891 (0.998) | 2.583 (0.995) |
| log S^f | -8.753 (0) | 0.94 (0.827) | -1.161 (0.123) | -1.015 (0.155) |
| (6): m log S^f | 3.972 (1) | 3.735 (1) | 4.175 (1) | 3.498 (1) |
| (7): m log S^f | -3.368 (0) | -4.958 (0) | -1.131 (0.871) | -0.984 (0.163) |
| Trends and Constants | | | | |
| Log LFP | -1.841 (0.033) | -0.175 (0.431) | -0.155 (0.439) | 0.401 (0.656) |
| log S^d | -6.784 (0) | -0.015 (0.494) | -5.985 (0) | -0.684 (0.247) |
| log S^f | -10.91 (0) | 3.082 (1) | 3.760 (1) | 3.291 (1) |
| (6): m log S^f | -0.716 (0.237) | -0.507 (0.306) | 2.749 (0.997) | 2.069 (0.981) |
| (7): m log S^f | -3.208 (0) | -3.7 (0) | -0.089 (0.464) | -0.608 (0.271) |

Notes: The null hypothesis is nonstationary while the alternative hypothesis for LLC (IPS) is that all (some) individual series are stationary with identical (individual) first order autoregressive coefficient. Both, the adjusted test statistics for LLC (t-star) and IPS (w[t-bar]) convergence asymptotically to a standard normal distribution. The p-values are in parenthesis.

In the first case, both testing procedures confirm unit roots for labor productivity and domestic R&D capital stock at a 10% level or less and therefore would not reject the H_0 of nonstationary. Considering foreign R&D capital stocks, IPS confirms unit roots for all specifications of foreign R&D capital stocks while LLC only confirms unit roots for the R&D-expenditure weighted foreign R&D capital stock by equation (6) for both lags. Adding a time trend to the test procedures, the IPS still confirms unit roots for all variables except domestic R&D capital

⁸ A GAUSS code for the proposed estimation techniques is provided freely on the homepage of Chihwa Kao at Syracuse University, NY: <http://web.syr.edu/~cdkao/>.

stock with lag 1. The LLC have to reject the H_0 at the 5% level for all variables except $\log S^f$ by equation (6), if variables are lagged by one period. However, with two lags only $\log S^f$ by equation (7) is stationary by LLC.

Given these somehow mixed results for time trends by the LLC unit roots test procedure, we now analyze the null hypothesis of stationary by LMH. Test statistics and p-values for level stationary – deterministic constant – and for trend stationary – deterministic constant and time trend – are given in Table 1b.

Table 1b: Stationary Tests by Hadri (2000)

(Annual data for 18 countries 1981-2001; Observations: 378; Lag: 1)

| | Homo. Disturbance | Hetero. Disturbance | Ser. Corr. Disturbance |
|-------------------------|-------------------|---------------------|------------------------|
| Level Stationary | | | |
| log LFP | 50.428 (0) | 48.542 (0) | 7.021 (0) |
| $\log S^d$ | 51.782 (0) | 51.587 (0) | 7.035 (0) |
| $\log S^f$ | 46.829 (0) | 49.906 (0) | 6.789 (0) |
| (6): $m \log S^f$ | 43.212 (0) | 34.240 (0) | 6.712 (0) |
| (7): $m \log S^f$ | 40.158 (0) | 41.417 (0) | 6.357 (0) |
| Trend Stationary | | | |
| Log LFP | 18.036 (0) | 15.111 (0) | 8.147 (0) |
| $\log S^d$ | 33.659 (0) | 32.083 (0) | 7.092 (0) |
| $\log S^f$ | 13.310 (0) | 25.254 (0) | 8.571 (0) |
| (6): $m \log S^f$ | 25.678 (0) | 22.802 (0) | 7.508 (0) |
| (7): $m \log S^f$ | 20.457 (0) | 17.985 (0) | 7.779 (0) |

Notes: The null hypothesis is level or trend stationary while the alternative hypothesis is non-stationary. The test statistics for the null of level stationary (Z_μ) as well as for the null of trend stationary (Z_τ) converge asymptotically to a standard normal distribution. The p-values are in parenthesis. Test statistics and p-values are given for three distinctive assumptions about the disturbance terms: homogenous, heterogeneous and temporally serially correlated.

The overall result is that the null hypothesis of stationary is rejected for every variable and for each specification of the disturbance term. The LMH test confirms unit roots and nonstationary for the whole data set. Bearing in mind the lack of power of unit root test since time trends are included and given the results in Table 1b, we conclude that the variables are nonstationary. Once confirmed that the variables are nonstationary and before turning to the empirical results, we have to be certain, that a regression containing all variables will have a stationary error term.

Two different testing procedures based on the EC model as well as on test statistics from Pedroni (1995) are listed in Table 2.

Table 2: Cointegration Tests based on the EC Model and on Pedroni (1995)
(Pooled data for 18 countries 1981-2001)

| Equation: | (3) | (5) with (6) | (5) with (7) |
|--|----------------|---------------------|---------------------|
| EC-Model ^a | | | |
| t-statistics of the EC- modell ^b | -3.62 (0) | -3.74 (0) | -3.63 (0) |
| LLC ^c | -6.449 (0) | -6.383 (0) | -6.448 (0) |
| IPS ^c | -6.602 (0) | -6.547 (0) | -6.601(0) |
| LMH ^d | -0.045 (0.518) | -0.025 (0.51) | -0.041 (0.517) |
| Pedroni (1995) ^e | | | |
| PC_1 | -7.579 (0) | - 6.286 (0) | -7.510 (0) |
| PC_2 | -7.397 (0) | - 6.135 (0) | -7.329 (0) |

Notes: Test statistics converge asymptotically to a standard normal distribution. The p-values are given in parenthesis.

^a The first step is to estimate long-run equilibrium values in levels by removing units as well as time effects (transformation for a two-way fixed effects model). The resulting residuals (i.e. error correction term) are used in the second step to estimate the EC model.

^b The t-statistic from the EC model indicates weather the lagged error correction term is significantly different than zero which means stationary of the residual term. Observations: 342 (Lag:1).

^c Test statistics and p-values are based on the residual term of the EC model. Individual-specific intercepts and time trends. Observations: 306 (Lag:1). The null hypothesis is nonstationary of the residual term of the EC model.

^d Test statistics and p-values are based on the residual term of the EC model. Trend stationary and heterogeneous disturbances across units. Observations: 342 (Lag:1). The null hypothesis is stationary of the residual term of the EC model.

^e Two test statistics are given by Pedroni (1995) based on a pooled Phillips and Perron type test. Regressors are assumed to be strictly exogenous and the null hypothesis is no cointegration.

Taking the EC model, the first testing procedure uses the lagged error correction term and analyzes statistical significance by means of the usual t-statistics of the EC model. The t-statistics are significantly different from zero for the three model specifications – equation (3) and equation (5) in combination with (6) or (7) – which means that the error term is stationary. The second testing procedure uses the residual term extracted from the EC model and applies the discussed unit roots and stationary tests. Again, for all model specifications the LLC and IPS test have to reject nonstationary and the LHM test has to confirm stationary for the residual term. Turning to the cointegration tests by Pedroni (1995), both test statistics reject the null hypothesis and confirm cointegration. Hence, we conclude that the variables are cointegrated.

Finally, with nonstationary and cointegrated data we turn to the empirical results. In Table 3 estimation results from pooled-OLS as well as fixed effects and random effects are given. As mentioned, the estimation results in finite sample are biased due to endogeneity and correlation issues and the bias remains substantial even for moderate sample size. We do not report any test statistics and signification levels and are only interested in the estimated coefficient

of the underlying economic variable and in its sign. Estimations for the three model specifications are shown in Table 3.

Table 3: Labor Factor Productivity by Pooled-OLS, Fixed Effects and Random Effects
(Pooled data for 18 countries 1981-2001; Observations: 378)

| Equation: | (3) | (5) with (6) | (5) with (7) |
|-----------------------|------------|---------------------|---------------------|
| Pooled OLS: | | | |
| $\log S^d$ | -0.001 | -0.01 | 0.015 |
| $\log S^f$ | 0.182 | | |
| $m \log S^f$ | | 0.572 | 0.026 |
| Fixed Effects | | | |
| $\log S^d$ | 0.113 | 0.203 | 0.143 |
| $\log S^f$ | 0.168 | | |
| $m \log S^f$ | | 0.293 | 0.031 |
| Random Effects | | | |
| $\log S^d$ | 0.006 | 0.01 | 0.036 |
| $\log S^f$ | 0.260 | | |
| $m \log S^f$ | | 1.224 | 0.046 |

Notes: We do not report any t-statistics and signification levels due to the estimation bias as a result of endogeneity and serial correlation in the small samples. The Hausman (1978) type test indicates for all model specifications that differences in coefficients are systematic.

The Hausman (1978) specification type test confirms – as expected – that for all model specifications that differences in coefficients are systematic and estimation results obtained by a fixed effects estimator seem to perform best. However, these test statistics may be bias driven as well. Regarding the fixed effect estimator, all variables are correctly signed which indicate a positive impact of domestic and foreign R&D capital stock on domestic labor productivity. However, only the estimated coefficient for domestic R&D capital stock for equation (3) is still plausible given comparable studies such as in Coe and Helpman (1995), whereas higher estimations of domestic R&D elasticities as for equation (5) in combination with (6) or (7) are certainly not. Given equation (3), the impact of foreign R&D capital stock on domestic labor productivity is 16.8 percent.

Now, we turn to the estimation results using the estimation techniques proposed by Kao and Chiang (1998) and correct for endogeneity and serial correlation. Table 4 lists coefficients and their test statistics in parenthesis obtained by OLS with bias correction, FM-OLS and DOLS for the three model specifications. Again, starting with the impact of domestic R&D capital stock on labor productivity, the estimated coefficients in equation (3) seem to be the most plausible and are comparable to the results from Kao, Chiang and Chen (1999) re-estimating

Coe and Helpman's 1995 paper.⁹ While the estimated coefficient by OLS with bias correction and FM-OLS are quite similar, the coefficient from the DOLS estimator is about two percent higher. This might be the result of the two different ways of removing the nuisance parameter. As described in equation (12), the FM-OLS estimator corrects the dependent variable by the long-run covariance and applies usual OLS. In contrast, the DOLS estimator introduces leads and lags by equation (13) in order to deal with endogeneity.

Table 4: Labor Factor Productivity by OLS with Bias Correction, FM-OLS and DOLS
(Pooled data for 18 countries 1981-2001; Observations: 378)

| Equation: | (3) | (5) with (6) | (5) with (7) |
|----------------------------------|-----------------|---------------------|---------------------|
| OLS with Bias Correction: | | | |
| $\log S^d$ | 0.081 (2.796)** | 0.194 (6.436)** | 0.119 (4.519)** |
| $\log S^f$ | 0.221 (6.496)** | | |
| $m \log S^f$ | | 0.430 (1.353) | 0.041 (6.322)** |
| R^2 | 0.674 | 0.616 | 0.6735 |
| FM-OLS: | | | |
| $\log S^d$ | 0.078 (2.558)** | 0.2 (6.299)** | 0.116 (4.210)** |
| $\log S^f$ | 0.219 (6.133)** | | |
| $m \log S^f$ | | 0.505 (1.514) | 0.041 (5.947)** |
| R^2 | 0.668 | 0.613 | 0.667 |
| DOLS: | | | |
| $\log S^d$ | 0.099 (2.58)** | 0.19 (4.792)** | 0.129 (3.737)** |
| $\log S^f$ | 0.174 (3.90)** | | |
| $m \log S^f$ | | 0.449 (1.078) | 0.036 (4.248)** |
| R^2 | 0.652 | 0.6 | 0.631 |

Notes: The bias corrected t-statistics are reported in parenthesis. * (**) denotes that the coefficient is significantly different from zero at a 10 percent (5 percent) level. All equations include unreported, country-specific constants. Assumptions for Dynamic OLS: 2 lags and 2 leads.

Turning to equation (5) in combination with either equation (6) or equation (7), estimations suggests higher elasticities for domestic R&D capital stock, which vary between 11.6 and 20 percent. As a first result, the estimated coefficients for domestic R&D capital stock differ largely depending on the estimation technique and on the assumptions on foreign R&D capital stocks. However, the t-statistics are significantly large and domestic R&D capital stock is significant at the 5 percent level.

⁹ Both papers estimate the impact of domestic R&D – amongst others – on total factor productivity. However, the impact of domestic R&D activity should not vary largely regarding either total or labor productivity figures.

Next, we consider foreign technology spillover effects. Depending on the specification of the intensity, coefficients for foreign R&D capital stocks are either very high for the case of domestic R&D expenditure weighted intensity by equation (6) or very low for the case of patent weighted intensity by equation (7). Such extreme values for foreign technology diffusion relative to the impact of domestic research activity are not very plausible, otherwise one could not explain –as mentioned – the small variation in productivity levels across different countries. However, without any additional specification for the intensity, the impact of foreign R&D capital stock on domestic factor productivity has reasonable values of about 22 percent for OLS with bias correction/ FM-OLS and 17.4 percent for DOLS and all of them are at a 5 percent significance level.

Given the superiority of the DOLS estimator over OLS with bias correction and FM-OLS and considering equation (3) as an adequate approximation of technological spillover effect due to patent application, we conclude that foreign patent related spillover effects are present. As a result, a one percent increase in domestic or foreign R&D capital stock leads to a 10 percent or 17.4 percent increase in domestic labor productivity respectively.

Labor Productivity for G7 and Non-G7 Countries

Since the benefits of domestic research activity depend on domestic markets and traded volumes, the impact on IFP due to domestic R&D spending may differ to a great extent between G7 and Non-G7 countries. Introducing a dummy variable $G7$, which is equal to one for the major 7 countries and zero otherwise, equation (3) and (5) can be easily modified by an additional regressor: $G7 \log S_{i,t}^d$. Table 5 shows the estimation results for the modified equations (3') and (5') by the DOLS estimator as well as the cointegration test statistics by Pedroni (1995).

As before, the variables are confirmed to be cointegrated and we drop equation (5') in combination with (6) or (7) due to the discussed plausibility reasons. Indeed, the impact of domestic research activity for the G7 countries rises to 25 percent while for the Non-G7 countries it remains nearly unchanged. Again, both coefficients are comparable to Kao, Chiang and Chen (1999) and to Coe and Helpman (1995). As a result, the elasticity figure of foreign R&D capital stock is reduced to 13.9 percent but is still significant at a 5 percent level.

Table 5: Labor Factor Productivity for G7 and Non-G7 Countries by DOLS
(Pooled data for 18 countries 1981-2001; Observations 378)

| Equation: | (3') | (5') with (6) | (5') with (7) |
|----------------------------|-----------------|----------------------|----------------------|
| DOLS: | | | |
| $\log S^d$ | 0.107 (2.89)** | 0.158 (4.217)** | 0.117 (3.613)** |
| $G7 \log S^d$ | 0.144 (1.568) | 0.204 (2.25)** | 0.196 (2.294)** |
| $\log S^f$ | 0.139 (3.072)** | | |
| $m \log S^f$ | | 0.629 (1.644) | 0.032 (4.014)** |
| R^2 | 0.683 | 0.657 | 0.68 |
| Cointegration-Test: | | | |
| Pedroni (1995) | | | |
| PC_1 | -7.314 (0) | -6.664 (0) | -7.783 (0) |
| PC_2 | -7.138 (0) | -6.504 (0) | -7.596 (0) |

Notes: The bias corrected t-statistics are reported in parenthesis. * (**) denotes that the coefficient is significantly different from zero at a 10 percent (5 percent) level. All equations include unreported, country-specific constants. The variable $G7$ acts as a dummy variable, which is equal to one for the seven mayor countries and zero for the non-G7 countries. Assumption: 2 lags and 2 leads.

6. Conclusion

I propose to use international patent applications as a diffusion channel to question the impact of technology spillover effects on input factor productivity. Considering the relationship between research and productivity, international patent pattern reflects the link between the source and the use of transferred technology. Analyzing a panel data set with 18 OECD countries from 1981 to 2001 by estimation techniques, which deal with nonstationary and cointegration issues, I find evidence that: (1) patent related foreign spillover effects are present, (2) foreign technology diffusion on labor productivity is about 14 percent and (3) in Non-G7 countries the impact on labor productivity is higher due to foreign rather than domestic R&D activities.

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