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PRODUCT VARIETY AND ECONOMIC GROWTH – EMPIRICAL EVIDENCE FOR THE OECD COUNTRIES

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Abstract

Utilizing panel data for 19 OECD countries we find support for the hypothesis that a greater degree of product variety relative to the US helps to explain relative per capita GDP levels. The empirical work relies upon some direct measures of product variety calculated from 6-digit OECD export and import data. Although the issue is still far from being settled, the merging conclusion is that the index of relative product variety across countries is significantly correlated with relative per capita income levels.

Keywords: Product variety, economic growth, panel data, OECD countries

JEL Classification: F41, F43, C23

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1. Introduction

One of the biggest challenges in economics today is to explain what causes different levels of per capita income in various countries. The first family of economic growth models trying to answer that question is organised around the basic neoclassical growth model devised by Robert Solow and Trevor Swan more than 40 years ago. Their models describe an economy of perfect competition, whose output grows in response to larger inputs of physical and human capital and raw labour. This economy obeys the law of diminishing returns: each new bit of physical and human capital yields a slightly lower return than the one before. Together, these assumptions give the traditional neoclassical growth model two crucial implications. First, the long-run growth rate depends upon technical progress, a variable that the model itself makes no attempt to explain. The second implication is that poorer countries should grow faster than rich ones. The reason is diminishing returns: since poor countries start with less physical and human capital, they should reap higher returns from investment. The other main approach, which has also spawned an extended family of models, goes by the name of endogenous growth theory. The idea is to endogenise technical progress, i.e. to bring improvements in productivity, notably due to innovation, imitation, product variety, human capital and public infrastructure, fully inside the model - so that ongoing economic growth emerges as a natural consequence. One important implication of these models is that economic growth is in the sphere of policy. Recently, so-called semi-endogenous growth models incorporating features of both families have been presented.

The empirical literature on economic growth has recently been enriched by various papers that provide direct tests of endogenous growth models using time series and panel data sets. These studies have produced conflicting evidence in that two of them, specifically Jones (1995) and Evans (1997), find evidence against the endogenous growth hypothesis while two other papers by Kocherlakota and Yi (1996, 1997) provide supportive evidence. The conclusions of the first two papers are largely based on the finding that shocks to the investment rate tend to be permanent while shocks to the growth rate are transitory. This finding is interpreted as running contrary to standard endogenous growth models, while it is consistent with the recent semi-endogenous growth literature. On the other hand, Kocherlakota and Yi (1996, 1997) present evidence from the UK and the US that suggests that permanent changes in policy variables influence the growth rate permanently, even though the growth rate appears stable over time.

In this paper we present new empirical evidence on the determinants of economic growth across countries focussing upon the impact of product variety. Given the importance of product variety in the recent economic growth literature, one would think that there are several well-known empirical papers studying whether greater product variety in fact does increase a country´s per capita income level in practice. Sadly, this is not the case.[1] To the best of our knowledge, there is only one disaggregated study analysing the link between product variety and growth in Korea and Taiwan.[2] In other words, the empirical evidence for growth model emphasising product variety as a potential channel for economic growth is much less persuasive than is commonly believed and therefore the profession´s faith in the merits of product variety may be hasty.

The remainder of the paper is organised as follows. In section 2 we formulate a simple theoretical model of the relationship between product variety and economic growth. Section 3 describes the methodology to estimate the degree of product variety in the OECD countries, describes the data, and presents the indices. Estimation results are given in section 4 and the last chapter summarizes our basic findings.

2. Product Variety and Economic Growth

In order to illustrate the interaction between product variety and economic growth we adapt a simple semi-endogenous growth model put forward by Jones (1998).[3] We suppose that countries produce a homogenous output good, Y, using labour, L, and a range of differentiated capital goods, x_j. Production in the final goods sector is given by

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[1] One indicator is that none of the most-well known international datasets used in the empirical growth literature (the Heston-Summers dataset, the Barro-Lee dataset and the World Bank World Development Indicators database) contain any information on product variety over time and/or across countries. Another indicator is that the up-to-date survey of the new growth evidence by Temple (1999) does not contain any work on product variety. The reason for this state of affairs is probably that direct measures of product variety are difficult to obtain and therefore empirical work in this area seems to be a risky business.

[2] Compare Feenstra et al. (1997) and (1999). In a related paper, Weinhold and Rauch (1997) have constructed a Herfindahl specialization index for 28 different manufacturing industries to analyse the link between openess, specialization and productivity growth. Owen and Wren-Lewis (1993) and Driver and Wren-Lewis (1999) have analysed the impact of variety and quality upon foreign trade using rough proxies such as cumulated investment and R&D flows.

[3] The model draws on the original theoretical analysis concerning the production of and the demand for 'variety' and 'quality' by Grossman and Helpman (1991).

(1) $$Y = L^{1-a}\int_0^n x_j^a \, dj$$

where $0 < \alpha < 1$ and

(2) $$\int_0^{n(t)} x_j(t)\,dj = K(t) \quad ,$$

i.e. the total number of differentiated intermediate goods used in production is equal to the total supply of capital.[4] Intermediate products are treated symmetrically throughout the model, so that $x_j = x$ for all j. Therefore intermediate goods are used the same amount, x, and we can determine x as

(3) $$x = \frac{K}{n}$$

The final-goods production function can then be rewritten as:

(4) $$Y = nL^{1-a}x^a \qquad \Rightarrow$$

(5) $$Y = nL^{1-a}n^{-a}K^a \qquad \Rightarrow$$

(6) $$Y = K^a(nL)^{1-a}$$

Thus, aggregate production for the economy takes the familiar Cobb-Douglas form and the degree of product variety, n, enters the production function just like labour-augmenting

[4] A similar production technology has been considered by Easterley et al. (1994).

technology and therefore is the ultimate engine of growth. The standard capital accumulation constraint is given by

(7) $$\dot{K} = s_K Y - dK$$

where s_K is the investment share of output and d is the rate of depreciation. The development of product variety over time is modelled as

(8) $$\dot{n} = f \; e^{y u} A^{g} n^{1-g}$$

where u is the proportion of the labour force dedicated to invent new differentiated intermediate goods and A is the world technology frontier. We assume $\phi > 0$, $\psi > 0$ and $0 < \gamma < 1$. Equation (8) has two features that merit discussion. First, the increase in product variety in an economy is proportional to the time inventing new products. Second, the last two terms in equation (8) suggest that the change in product variety is a weighted average of the world frontier level of product variety, A, and the individual country´s degree of product variety, n. Equation (8) can be rewritten by dividing both sides by n:

(9) $$\frac{\dot{n}}{n} = f \; e^{yu} \left(\frac{A}{n} \right)^{g}$$

Equation (9) makes clear that the growth rate of product variety in the economy is positively related to the ratio (A/n). The closer an individual country´s degree of product variety, n, is to the world frontier of variety, A, the smaller the ratio A/n, and the smaller is the growth rate of n. Finally, we assume that the world frontier expands at a constant rate g, i.e.

(10) $$\frac{\dot{A}}{A} = g$$

and that the labour force of the economy grows at the constant rate m.

In order to solve for the steady state growth path, we proceed in the usual fashion. Along the balanced growth path we have $g = g_y = g_n = g_A$, i.e. the long-run growth rate is given by the (exogenous) growth rate of the technological frontier, A.[5] Steady state output per capita y^* along the balanced growth path is given by

$$(11) \qquad y^*(t) = \left(\frac{s_K}{m + g + d} \right)^{a/1-a} n^*(t)$$

or

$$(12) \qquad y^*(t) = \left(\frac{s_K}{m + g + d} \right)^{a/1-a} \left(\frac{f}{g} e^{yu} \right)^{1/g} A^*(t)$$

The model proposes two answers to the question of why different economies have different steady state income levels. First, the model emphasises the importance of product variety, providing a "new growth theory" interpretation of the basic neoclassical growth model since the steady state income level, y^*, depends upon the degree of product variety, n. In the model increased product variety accelerates per capita income levels by more fully realizing dynamic economies of scale. Second, the initial term in brackets in (11) and (12) is similar to the basic Solow model. This term implies that countries investing more in physical capital will be richer.[6] In order to understand the mechanics of the model, let us consider a country which decides to open up its economy to the rest of the world. We can model this as an increase in ϕ. According to

[5] Even with no differences across countries in the long-run growth rate, one can explain a large variation in rates of growth with transition dynamics.

[6] In extensive sensitivity analyses of cross-country growth regressions, Levine and Renelt (1992) and Sala-i-Martin (1997) have shown that investment in physical investment is the most robust variable explaining cross-country growth differences. Explaining differences in the level of income across countries by appealing to differences in n and s_K, however, obviously begs new questions. Why is it that some countries invest more in physical capital than others and why do individuals in some countries spend more time u to develop new intermediate goods? This model cannot address this question. A more complete model answering this question has to assume utility-maximising individuals to choose to work in either the final-goods sector or in

continued

(12), a higher value of ϕ rises y^*. Starting from steady state, the higher ϕ causes the growth rate of n to be higher than g along the transition to the new steady state. Over time, however, the ratio A/n is decreasing, and therefore the growth rate of n returns to g. In other words, policy changes like opening up the economy (interpreted as an increase in ϕ) have a long-run *level* effect but no long-run *growth* effect just like in the original Solow model. It is this link between product variety and per capita income that we shall test below.

3. Measuring Product Variety Across OECD Countries

The questions we would like to address in this empirical section is how to pick a value of n, i.e. how to measure the 'supply-side' factor product variety. In order to get a direct measure of the variety of products across countries, the following two questions have to be addressed:

1. Which methodology can be used to estimate the degree of product variety across countries?
2. What highly disaggregated data do we have on differentiated products which are consistent across countries?

In the following empirical work we adapt the methodology developed by Feenstra (1994) and Feenstra and Markusen (1994). They have shown how an exact measure of product variety can be constructed from a CES production function when the inputs enter non-symmetrically.[7] The procedure considers two units of observations denoted by s and t representing either two time periods or two countries. Suppose that output y_t in period t (country t) is given by the production function

$$(13) \qquad y_t = f(x_t, I_t) = \left[\sum_{i \in I_t} a_i x_{it}^{(s-1)/s} \right]^{s/(s-1)}$$

the intermediate goods sector expanding product variety. In order to simplify the analysis, we will not develop this more complete here.

[7] In the theoretical model above the inputs entered a Cobb-Douglas aggregator function in a symmetric way. For empirical purposes, however, it seems more reasonable to allow the inputs to enter the production function non-symmetrically and the elasticity of substitution to differ from one.

where σ denotes the elasticity of substitution, x_{it} is the quantity of input i in period t (country t), and the total set of inputs in period t (country t) is denoted by I_t. For example, when the inputs available in period t (country t) are numbered 1 through N_t, then $I_t = \{1,...,N_t\}$. The corresponding cost function is

$$(14) \qquad c\,(p_t, I_t) = \left[\sum_{i \in I_t} b_i\, p_{it}^{(s-1)/s}\right]^{s/(s-1)}$$

where p_{it} are the prices of the inputs and $b_i = a_i^{\,s}$. A standard definition of an input index is the change in nominal expenditure (E_t/E_s) deflated by an input price index, where $E_t = \sum p_{it} x_{it}$. Following Feenstra (1994), Feenstra et al. (1997) and Feenstra et al. (1999), we chose the set of intermediate products common to both periods (countries) as $I = I_s \cap I_t$. The quantity index for intermediate inputs is then measured by

$$(15) \qquad Q\,(p_s, x_s, p_t, x_t) = \frac{\dfrac{E_t}{E_s}}{P\,(p_s, x_s, p_t, x_t)}$$

where $P(\cdot)$ is the input price index given by the Sato (1976)-Vartia (1976) formula. Total factor productivity, TFP, is defined as the difference between the growth of output and this input index, i.e.

$$(16) \qquad TFP_{st} = \ln\left(\frac{y_t}{y_s}\right) - \ln\left(\frac{\dfrac{E_t}{E_s}}{P\,(p_s, x_s, p_t, x_t)}\right)$$

$$(17) \qquad TFP_{st} = \frac{1}{(s-1)}\,\Delta PV_{st}$$

where the change in product variety ΔPV_{st} is defined as follows:[8]

$$\text{(18)} \qquad \Delta PV_{st} = \ln \left(\frac{\dfrac{\sum_{i \in I_t} p_{it} x_{it}}{\sum_{i \in I} p_{it} x_{it}}}{\dfrac{\sum_{i \in I_s} p_{is} x_{is}}{\sum_{i \in I} p_{is} x_{is}}} \right).$$

To interpret equation (18), consider the case where the set of inputs is growing over time, i.e. we have two sets $I_s = \{1, ..., N_s\}$ and $I_t = \{1, ..., N_t\}$ with $N_t > N_s$. In this case the common set of products is $I = I_s$ and the denominator is 1. The numerator will exceed unity, indicating that product variety has increased over time. In the special case where all inputs enter (13) and (14) symmetrically, the numerator in (18) simplifies to N_t/N_s.

The two observations s and t in (13) - (18) can either be interpreted as two successive observations in time or as two countries. The first interpretation allows to construct an index of product variety in a specific country over time, while the second interpretation allows to construct the level of product variety across countries.

The procedure above is implemented using highly disaggregated annual world export and import data at the 6-digit industry level for the years 1989 to 1996 for 19 OECD countries.[9] The most important advantage of these data is that the classification of goods is consistent across countries. On the other hand, these data have obviously their problems. First, the time series dimension of the data (8 years) is rather short. Second, some intermediate goods produced at home are not traded internationally. Nevertheless, we believe the topic to be of such economic and social significance, that a willingness to experiment with trade data is justified, specially since the most important goods are probably exported and/or imported.[10] Our first measure of product variety is

[8] Compare Feenstra (1994), Proposition # 1.

[9] The classification distinguishes about 6400 commodities. Data were collected from the OECD database INTERNATIONAL TRADE BY COMMODITIES STATISTICS - ITCS Classification, Paris 1997. All data are espressed in current US $. In principle it would be preferable to use national production data but they are neither available at a sufficiently disaggregated level nor are the available data consistent across countries.

[10] In their extensive discussion of quality and variety, Grossman and Helpman (1991) and Coe and Helpman (1995) and Bayoumi et al. (1999) have focused on levels of investment in R&D at home and abroad. A clear

continued

export variety in country i (i = 1, ..., 18) relative to the US (ΔPV_{EX}). The results for 1989 - 1996 are summarized in Figure 1, which we invite the reader to review. The first impression is that export variety in all countries under investigation is lower than in the US.[11] The degree of relative export differentiation is highest in Canada, followed by France, the UK, Germany, Denmark and Italy. In contrast to this group of countries, the degree of export variety is much lower in Greece, Portugal, Norway and Turkey. The lowest ratio turns out for Iceland. One problem with indicators of product variety focussing solely upon export data is that even when differentiated inputs are not produced at home they are in principle available in other countries through trade. In other words, product variety in any country does not only depend on exports but potentially also upon imports. As a second measure of relative product variety we have therefore calculated product variety relative to the US based upon exports and imports ΔPV_{EXIM}. The results are given in Figure 2 below. With the distinguishable exception of Iceland, relative product variety in the various countries now looks much more like in the US. Finally, we have also calculated both product variety measures for secondary industries which rely much more on differentiated products and therefore fit the idea of endogenous growth much clearer than primary industries that rely more heavily on natural resources ($\Delta PV_{EX\text{-}SEC}$ and $\Delta PV_{EXIM\text{-}SEC}$).[12] The results for this subsample of the whole dataset are given in Figure 3 and 4.

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problem here is that the lag between R&D expenditures and the production of new varieties could be very long. Furthermore, it is also the case that many improvements in quality and variety can be realised without any R&D expenditure being incurred. In particular, increases in variety can occur through imitation, which involves little or no R&D expenditure.

[11] Negative (positive) values for the index indicate lower (higher) product variety than in the US. The negative numbers are a result of the log transformation in (18).

[12] Industrial coverage of the ITCS classification groups industries into 'Primary Products' (textile products, wood products, paper & printing, rubber products, primary metal, leather products and stone, clay & glass) and 'Secondary Products' (food products, beverages & tobacco, apparel, chemicals & plastics, fabricated metal products, machinery, electrical products, transportation equipment and instruments).

Figure 1: Relative Product Variety Based Upon Exports

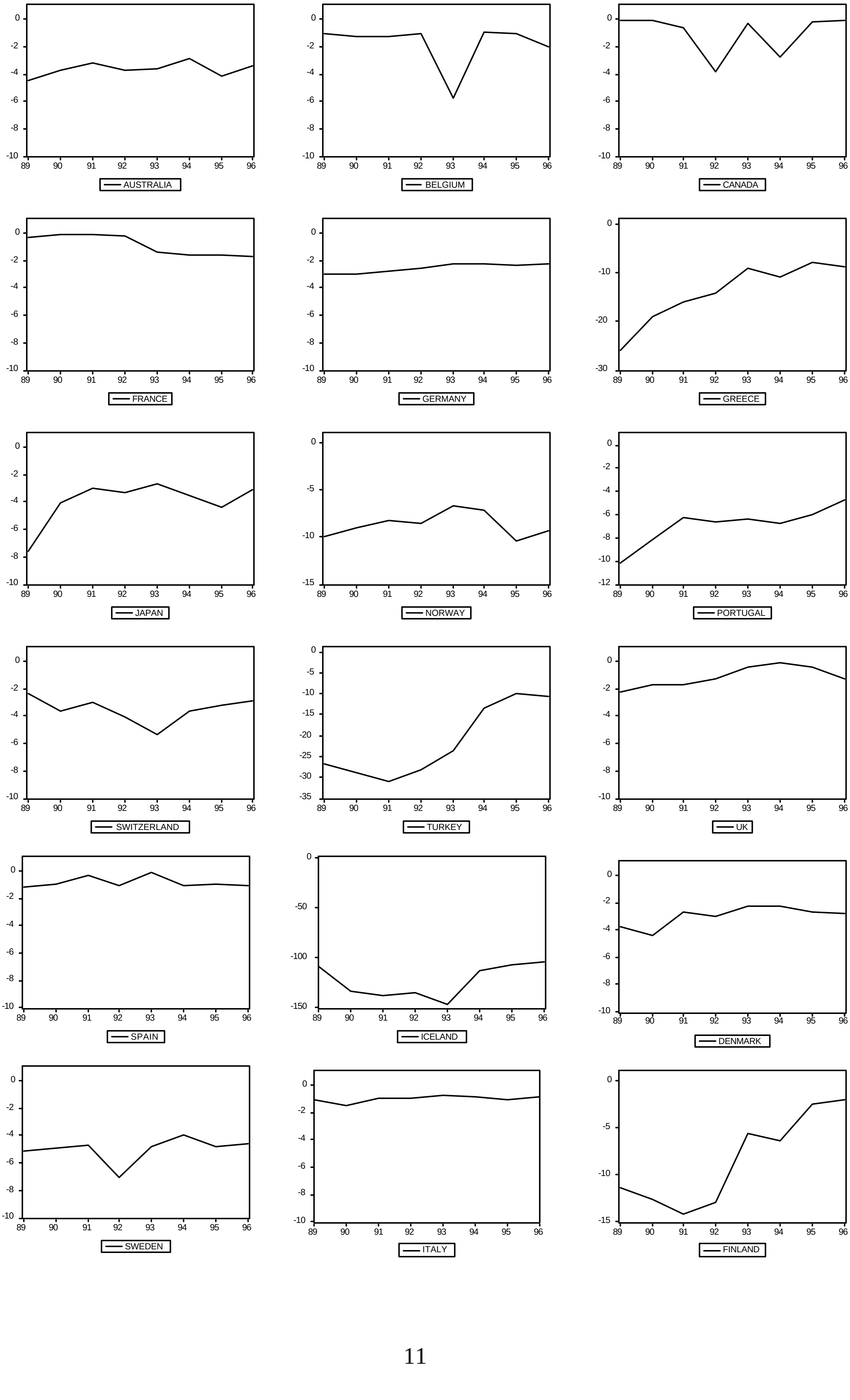


Figure 2: Relative Product Variety Based Upon Exports and Imports

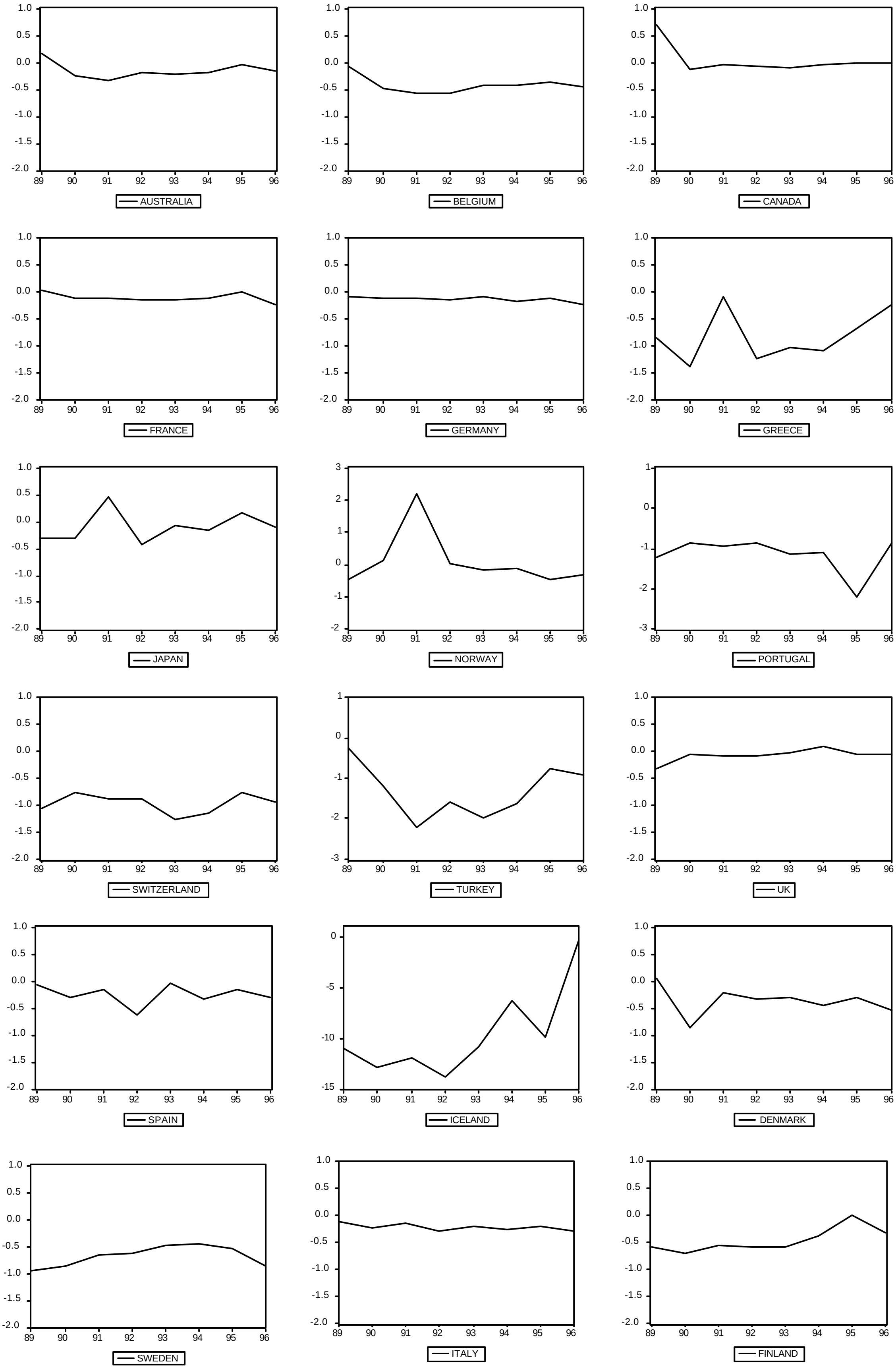


Figure 3: Relative Product Variety Based Upon Exports of Secondary Products

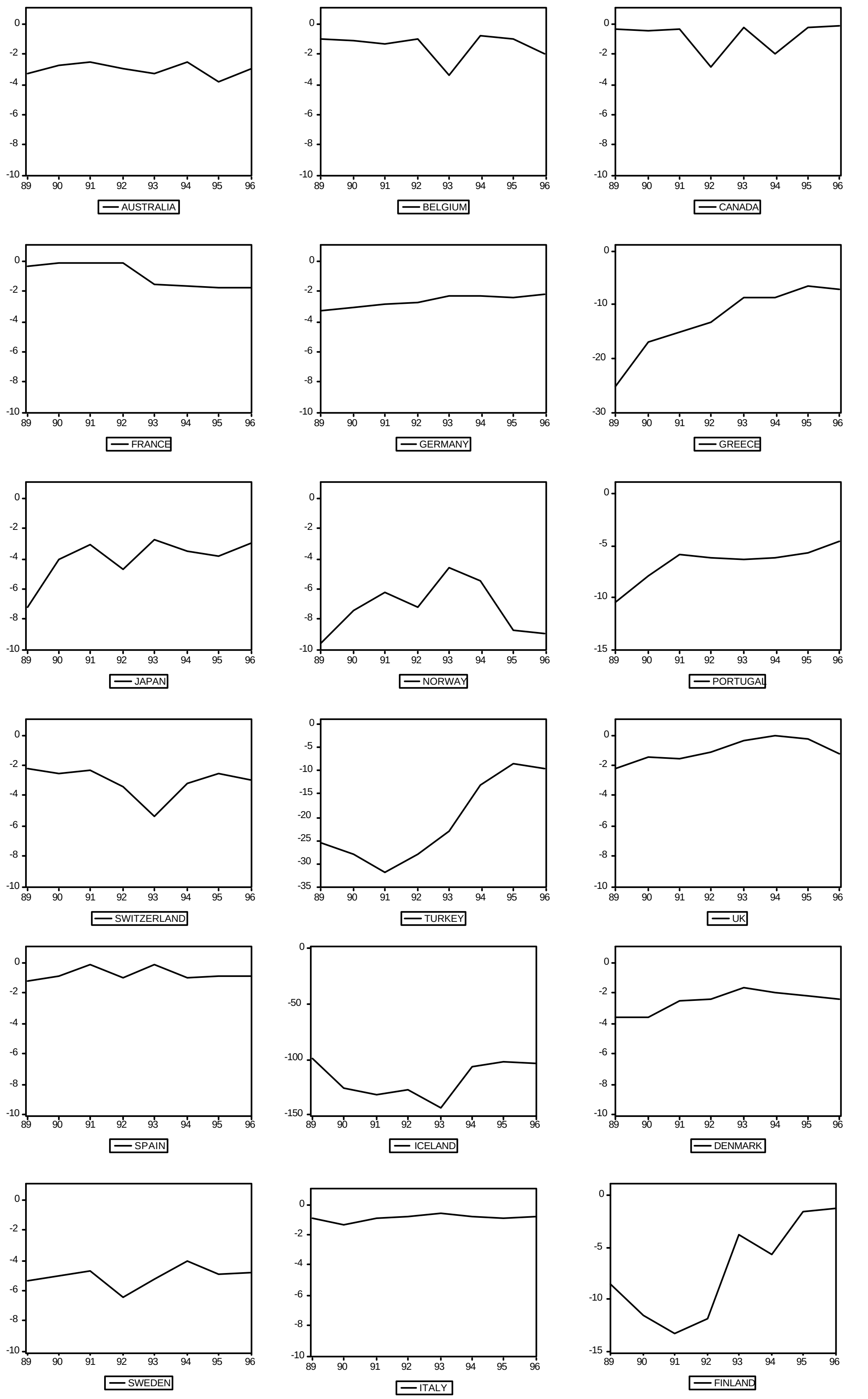
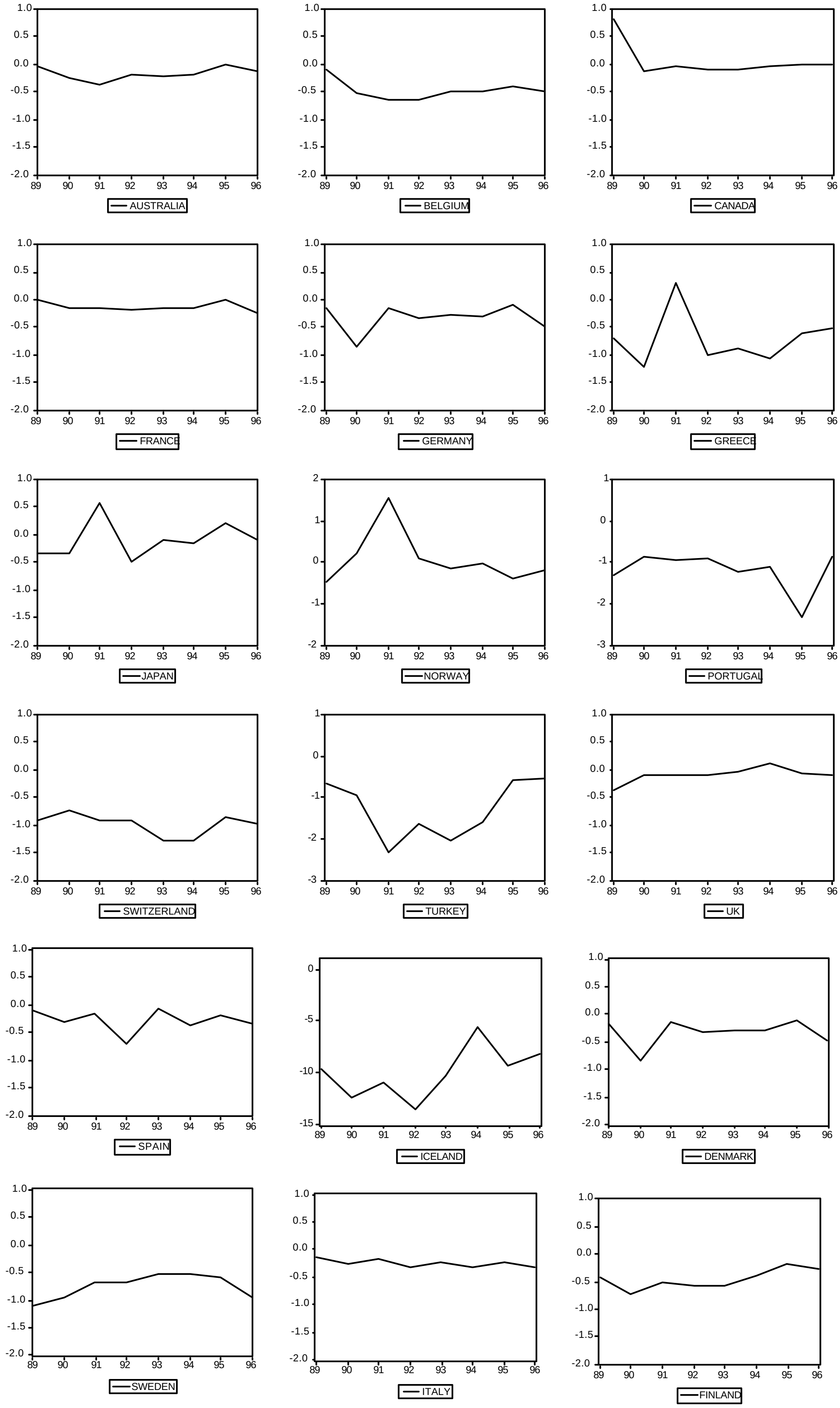


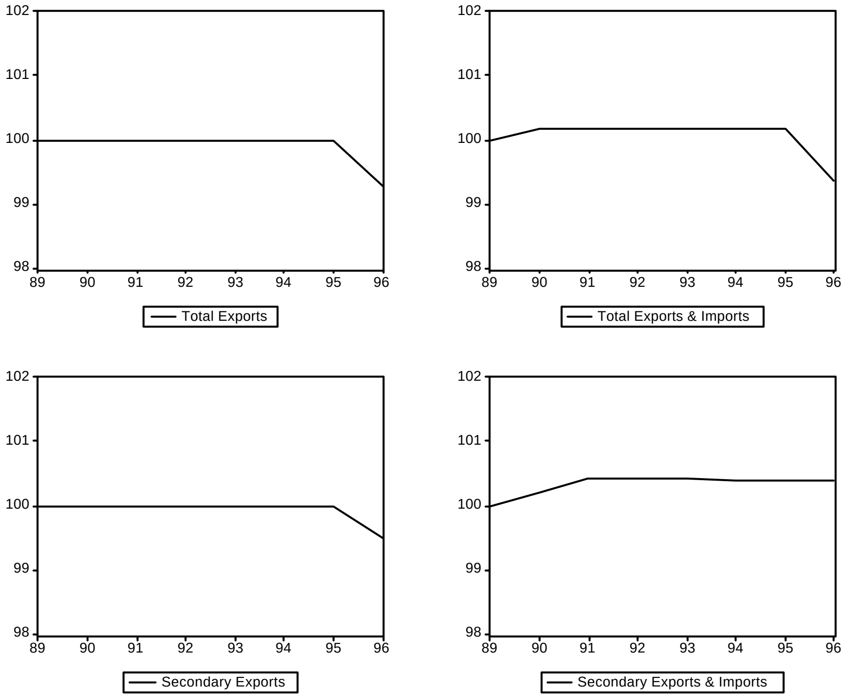
Figure 4: Relative Product Variety Based Upon Exports and Imports of Secondary Products



When the two observations *s* and *t* in (13) - (18) are interpreted as two successive observations in time, than time series for product variety in individual countries can be calculated. Figure 5 below gives annual indices for product variety in the United States for the years 1989 to 1996.[13]

Figure 5: Index of Product Variety in the U.S., 1989 - 1996

(All Series 1989 = 100)



The four time series indicate that product variety in the United States has hardly changed at all over the sample period.

4. Econometric Results

This paper utilizes an estimation method that exploits the full time dimension of the data by using all the information from a full panel rather than just the time-averaged information from a cross-section. The choice of the list of regressors depends upon equation (11). All variables are

[13] To save space, the time series for the other OECD countries in the dataset are not reported but are available from the authors upon request.

expressed relative to the US since the product variety measures are given in ratio form.[14] The relative investment share, *IY*, is added to the regressions to capture different per capital income levels arising from different levels of investment in physical capital.[15] In addition to the product variety indices which are the variables of immediate interest, additional control variables were included in the regressions. The 'fixed effects' are controlled for directly through country-specific dummies. The advantage of including 'fixed effects' is that we are explicitly holding constant a bunch of factors which are very difficult to control for in cross-country comparisons (either because we are not sure of what these factors are or because we do not have the necessary data). This is important because it allows us to get a clearer picture of the interactions among the variables highlighted by the theoretical model presented above. A final ingredient are country-specific time trends.[16] The basic model for country *i* and time *t* thus becomes

$$(19) \quad Y_{it} = a_i + d_i T + b \; IY_{it} + g \; \Delta PV_{it} + e_{it} \qquad i = 1,...,18,\ t = 1989,...,1996$$

where Y_{it} is per capita GDP in country *i* relative to the US in percent (PPP, constant 1987 international \$), IY_{it} is defined as the share of investment in GDP in country i relative to the US in percent, and ΔPV_{it} is product variety relative to the U.S. The α_i (δ_i) parameters represent the fixed effects (country-specific time trends).[17] The country effects control for any persistent differences across countries, such as initial conditions, higher level of technical know-how, cultural differences, higher government investment expenditures, or freer access to knowledge.[18] One potential problem with equation (19) is the variable IY_{it} needs to be treated as endogenous. There is less concern about the endogeneity of the product variety measures since they have been derived from completely different datasets and there is therefore no possibility of correlation due

[14] Canova and Marcet (1995) argue that such a normalization should eliminate a significant part of the cyclical noise in the data.

[15] Y_{it} and IY_{it} were calculated using data from the *World Development Indicators* 1998 database.

[16] We have included a set of year dummies which should largely take care of the contemporaneous correlation across countries.

[17] The methodology followed by the vast majority of researchers up to 1995, i.e. cross-sectional regression, was based on the hypothesis of a growth process characterized by a smooth path towards a steady state. As Islam (1995) and Caselli et al. (1996) have demonstrated empirically that this underlying hypothesis is invalid, we have used panel data estimates which are not subject to this restrictive hypothesis on the growth process and allow for heterogeneity in steady state output levels.

[18] Since we include country dummiey variables, we cannot include initial per capita GDP which also varies across countries but not over time.

to common trends arising among macroeconomic variables. In order to deal with this causality and simultaneity bias, we have produced *IV* estimates of equation (19) that allow for heteroscedasticity of general form.

Prior to estimating equation (19) we have, however, analysed the univariate time series properties of the variables under consideration. Im et al. (1995) have presented an ADF-type unit root test which increased the power of univariate unit root tests by exploiting the panel structure of the data.[19] The null hypothesis being tested across units is H_0: $\beta_i = 0 \ \forall i$ against H_1: $\exists \ i$ s.t. $\beta_i < 0$ where β_i are the coefficients of the lagged levels terms in the individual ADF equations. Due to the heterogeneity each equation is estimated separately by OLS and the heterogeneous panel test statistic is then obtained as the (studentized) average of the test statistics for each equation. The *t*-bar statistic is given by:

$$(20) \qquad \Gamma_{\bar{t}} = \frac{\sqrt{N}\left(\bar{t} - E\left(t_i \mid b_i = 0\right)\right)}{\sqrt{Var(t_i \mid b_i = 0)}}$$

where *t*-bar is the average of the individual ADF *t*-statistics and the other two terms in (20), $E(\cdot)$ and $Var(\cdot)$, are the associated means and variances, respectively, for which Im et al. (1995) provide values on the basis of simulations for various lag lengths and sample sizes. Under the null hypothesis of a unit root this statistic has a standard normal distribution and is valid in the presence of heterogeneity across units as well as residual serial correlation across time periods. Under the alternative of stationarity, this statistic diverges to negative infinity. Table 1 presents the *t*-bar statistic.

[19] Based upon Monte Carlo evidence, the Im et al. (1995) unit root test for heterogeneous panels has substantially more power (and better small sample properties) than individual ADF tests.

Table 1: Panel Unit Root Tests for the Sample of 18 OECD Countries

	Without Trend	With Trend
Y	-43.66	-15.23
IY	-23.79	-19.04
ΔPV_{EX}	-19.57	-15.18
ΔPV_{EXIM}	-41.31	-13.43
$\Delta PV_{EX\text{-}SEC}$	-31.76	-12.99
$\Delta PV_{EXIM\text{-}SEC}$	-51.38	-16.54

Notes: The statistics have been calculated using demeaned data.

All t-bar statistics easily reject the null of nonstationarity. This implies that we don´t have to consider cointegration tests for panel data.[20] The IV estimation results are available in Table 2 below.

Table 2: *IV* Growth Regressions for the OECD Countries

IY	0.18 (6.8) [8.3]	0.14 (6.4) [7.3]	0.18 (6.6) [8.0]	0.16 (6.6) [7.6]
ΔPV_{EX}	0.11 (3.5) [2.6]	-	-	-
ΔPV_{EXIM}	-	0.39 (3.7) [3.4]	-	-
$\Delta PV_{EX\text{-}SEC}$	-	-	0.09 (3.0) [2.6]	-
$\Delta PV_{EXIM\text{-}SEC}$	-	-	-	0.37 (2.4) [2.3]
N×T	126	126	126	126
R^2 of the first stage	0.99	0.99	0.99	0.99

Notes: Standard t-values are given in parentheses while White´s heteroscedasticity-consistent t-values are given in square brackets below the coefficients. The sample period is 1990 - 1996. The country-specific fixed effects and the country-specific time trends are included but not reported. F-statistics indicate that both sets of dummy variables are significant at the 1 percent level. See text for data definitions and sources. Alternatively, we have also estimated (19) with AR1-IV methods but the ρ-coefficient turned out insignificant.

[20] McCoskey and Kao (1998) and Pedroni (1999) have recently proposed various residual-based cointegration tests for heterogeneous panels.

Table 2 reveals two main conclusion. First, all coefficients are signed in a manner consistent with the theoretical model. The coefficients of the relative investment share and the product variety measures are significant in all four equations.[21] This implies that investment in physical capital does not carry all the information relevant for economic growth. Second, an interesting result is that the coefficients are similar for all products and for secondary products. Overall, the generally positive association between product variety and per capita income provides some degree of confirmation for the semi-endogenous growth model presented above.[22]

One potential problem and bias of the cross-country estimates in Table 2 may be that all variables are normalized relative to the U.S. Supposing that the model works fine for the U.S. with eight time series observations and that the observations for the variables in the other countries are white noise, the model test should be insignificant.[23] But by measuring each country relative to the U.S., the American model is replicated in all other countries as they are measured relative to the U.S. This reduces the standard error of the estimated coefficients and at the same time, a bias emerges in the estimated coefficients in the direction of the American model. In other words, the empirical test of the product variety theory presented above may not be invariant to the empirical specification. To test for the robustness of the results, we have therefore estimated the equation

(21) $$y_{it} = a_i + d_i T + b \; iy_{it} + g \; \Delta pv_{it} + e_{it} \qquad i = 1,...,19, \; t = 1989,...,1996$$

where y_{it} is the logged level of per capita income in country i, iy_{it} is the investment share in country i in percent, and Δpv_{it} is the index (1989=100) of product variety in country i relative to the product variety in 1989 for the country in question (compare Figure 5 for the U.S. measures). In other words, equation (21) does not contain any information about the level of product variety

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[21] This result is consistent with the empirical evidence for Korea, Taiwan and Japan in Feenstra et al. (1997) and (1999).

[22] As the paper measures product variety in traded goods and not only capital goods, an alternative interpretation of this result, however, is the demand theory formulated by Linder (1991), where high income countries have a more advanced and differentiated consumption structure. According to Linder´s (1991) theory, the causal link runs from real income per capita to the degree of product variety. Barker (1977) acknowledges the contribution of Linder (1961) and develops a similar variety hypothesis according to which consumers love variety and therefore exports and imports tend to increase more than proportionally with real income per capita.

[23] This objection raised against the empirical evidence in Table 2 does not seem to be rather realistic given the visual impression from Figure 5.

across countries. The α_i (δ_i) parameters again represent the fixed effects (country-specific time trends) which have been included to capture cross-country differences in steady states arising from a rich set of conditioning variables. A final feature of equation (21) is that the number of observations increases with 1×8. The *IV* estimation results are given in Table 3 below.

Table 3: *IV* Growth Regressions for the OECD Countries

iy	0.01 (3.7) [3.6]	0.01 (2.9) [2.6]	0.01 (3.8) [3.7]	0.01 (4.0) [4.0]
Δpv_{EX}	0.001 (0.4) [0.4]	-	-	-
Δpv_{EXIM}	-	0.001 (1.6) [1.3]	-	-
$\Delta pv_{EX\text{-}SEC}$	-	-	0.001 (0.2) [0.3]	-
$\Delta pv_{EXIM\text{-}SEC}$	-	-	-	0.007 (0.6) [0.6]
N×T	133	133	133	133
R^2 of the first stage	0.99	0.99	0.99	0.99

Notes: Standard *t*-values are given in parentheses while White´s heteroscedasticity-consistent *t*-values are given in square brackets below the coefficients. The sample period is 1990 - 1996. The country-specific fixed effects and the country-specific time trends are included but not reported. *F*-statistics indicate that both sets of dummy variables are significant at the 1 percent level. See text for data definitions and sources. Alternatively, we have also estimated (21) with AR1-*IV* methods but the ρ-coefficient turned out insignificant.

Overall, the results for the investment share variable are qualitatively similar to the results in Table 2. On the other hand, the four product variety measures lose their significance. In other words, the panel results point to the strong rejection of Δpv_{it} in equation (21). To put two and two together, a "salomonic" interpretation of the results is that the cross-country level of product variety in (19) is more informative when trying to determine relative steady state income levels than the pure time series dimension within countries. Another fair conclusion is that the significant results in Table 2 are mainly determined by the cross-sectional variation in the data. One reason for this might be that the short sample period does not allow to separate short-term fluctuations from long-term growth dynamics.

5. Conclusions

Understanding growth is surely one of the most important tasks in economics. This paper is part of an upsurge of empirical work on growth and tries to shed some light on the relative merit of models emphazising the importance of product variety. The distinctive feature of this paper consists in calculating direct measures of product variety relative to the US for 18 OECD countries from 1989 to 1996. Although the issue is still far from being settled, the merging conclusion is that the index of relative product variety is significantly correlated with relative per capita income levels. The empirical results have some implications for the debate between openess and growth which normally is rather silent on the issue of the mechanism through which this robust empirical relationship occurs. In this paper we have identified one channel through which increased trade may lead to growth, namely a strongly outward-oriented trade regime makes a greater variety of products and technologies available. However, the assessment of the performance of the various product variety indices for a larger set of countries and longer time spans is important and has to be investigated when additional data are available. We leave this for future research. In the meantime, maintaining outward-oriented pro-trade policies can have significant benefits.

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