Export variety and economic growth in East European transition economies*

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Abstract

Utilizing panel data for 14 East European transition economies, we discuss the link between product variety and growth. The empirical work relies upon some direct measures of product variety calculated from 5-digit OECD trade data. On balance, the results suggest that the growth patterns of the East European transition countries may be best represented by Ventura’s (1997) model of outward orientation and integration with the world economy.

JEL classifications: C33, F43, O31, O33, O52.
Keywords: Product variety, transition economies, Eastern Europe, economic growth, panel data.

* The paper was written while the first author was a Visiting Scholar at the Research Department of the Bank of Finland (see www.bof.fi/bofit/eng/6dp/03abs/pdf/dp0803.pdf). The opinions expressed in the paper are strictly personal and do not necessarily reflect those of the Bank of Finland. We are most grateful to Philippe Aghion and an anonymous referee whose comments and suggestions have greatly contributed to enhancing the quality of the paper.
1. Introduction

In the past two decades, the study of the forces that shape per capita income levels and growth rates over the long run has become one of the most attractive areas of economic research. On the theoretical front, various endogenous and semi-endogenous growth models were developed to explain why sustained growth occurs in the absence of exogenous growth in total factor productivity. On the empirical front, the Summers and Heston (1988, 1991) dataset initiated an impressive empirical growth literature.

In this paper, we present new empirical evidence on the determinants of economic growth across East European transition economies with a focus on the impact of product variety. Given the importance of product variety in the recent economic growth literature, one might expect that the widespread acceptance of the new approach arose from its empirical success. In fact, this is not the case. Although the theoretical papers make much of the plausibility of anecdotal evidence and its consistency with stylized facts, none test the new theory formally.\(^1\) A possible reason for this discrepancy between economic theory and empirical and econometric evidence is that none of the best 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 or across countries. Therefore, the impact of product variety has often been taken for granted although the evidence is scattered and the link between product variety and growth has not been subject to a great amount of empirical scrutiny. This paper attempts to fill this gap.\(^2\) A key ingredient of the transition process is the structural change consisting in the reallocation of resources on the basis of market incentives. Product variety is therefore a potentially useful concept in analysing the structural changes that have actually occurred in Eastern European transition economies.\(^3\)

The remainder of the paper is organized as follows: In Section 2, we review theoretical models of the relationship between product variety and economic growth. Section 3 describes the methodology for estimating the degree of product variety in transition countries, explains the data and presents the indices. Estimation

\(^1\) Moreover, Temple’s (1999) survey of new growth evidence and the up-to-date growth resources available at www.bris.ac.uk/Depts/Economics/Growth contains no work on product variety. The reason is probably that direct measures of product variety are notoriously difficult to obtain even within a single country, and international comparisons yet more difficult.

\(^2\) Compare Addison (2002), Feenstra et al. (1999a) and (1999b), Funke and Ruhwedel (2001) and Feenstra and Kee (2004). Based on similar methodologies to measure product variety, the UN Economic Analysis Division has recently published a comprehensive analysis of the benefits from product differentiation in modern economies (Economic Survey of Europe, 2004, No. 1, Chapter 6; to be found at www.unece.org/ead/survey.htm).

\(^3\) For a survey of the literature on growth in transition, see Campos and Coricelli (2002) and the references cited therein.
results are given in Section 4. Section 5 concludes with a general summary of the paper and a discussion of some outstanding issues.

2. Expanding product variety and economic growth

To illustrate the interaction between product variety and economic growth and develop a testable hypothesis, we first provide an informal discussion of the simple semi-endogenous growth model put forward by Jones (1998) in which economies become more productive as a widening of the product spectrum available occurs. Semi-endogenous growth means that (i) technological change itself is endogenous, while (ii) long-run growth is pinned down by exogenous factors. The importance of hypothesis (ii) lies in the property that the steady-state growth rate is independent of public policy. There is a single final output, \( Y \), produced by labour \( L \) and differentiated capital goods \( x_j \), with \( j \in [0, n] \). Time is continuous. Without much loss of generality, we use the familiar constant-returns Cobb–Douglas production function and assume that labour supply is offered inelastically, so that

\[
y = L^{1-\alpha} \int_0^n x_j^\alpha \, dj, \tag{1}
\]

where \( 0 < \alpha < 1 \), and

\[
\int_0^n x_j \, dj = K. \tag{2}
\]

Thus, the total amount of differentiated intermediate goods used in production equals the total supply of capital.

We can also interpret \( n \) as a measure of the complexity of production. The basic idea is that a larger variety of intermediate goods allows producers to increase productivity through selection of intermediate inputs that more closely match their production requirements. 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

\[
x = \frac{K}{n}. \tag{3}
\]

Let us now look at the output dynamics in this economy. By rearranging, we get

\[
y = K^\alpha (nL)^{1-\alpha}. \tag{4}
\]
Thus, aggregate production for the economy takes the familiar Cobb–Douglas form. The degree of product variety, $n$, enters the production function just like labour-augmenting technology. Capital evolves according to

$$\dot{K} = s_K Y - \delta K,$$

where $s_K$ is the investment share of output and $\delta$ is the rate of depreciation. The product variety dynamics obey

$$\dot{n} = \phi A^{\gamma} n^{1-\gamma},$$

where $\phi$ is a reduced-form coefficient that reflects, among other things, the share of labour devoted to R&D. We assume $\phi > 0$ and $0 < \gamma < 1$. Equation (6) reflects the fact that every act of innovation builds on previous ideas, i.e., $\dot{n}$ is a function of $n$. The last two terms in Equation (6) 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, $n$. In the following empirical part of the paper we think of the US as the technological frontier. Equation (6) can be rewritten by dividing both sides by $n$, whereby

$$\frac{\dot{n}}{n} = \phi \left( \frac{A}{n} \right)^{\gamma}.$$

Equation (7) makes clear that the growth rate of product variety in the economy positively relates 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 the growth rate of $n$. In other words, the scope for differential variety growth between the ‘technology leader’ and the catching-up economy (the ‘laggard’) is higher where the initial gap is larger since lower amounts of R&D investment are required to introduce new products already available elsewhere than to develop them ex novo.

It can be argued that this Schumpeterian growth model is only relevant for a handful of advanced economies that undertake significant investments in R&D and are located on the frontier of technological development. On the other hand, the issue for most East European economies is not whether to devote resources to innovation and perform leading-edge R&D but whether to adopt and assimilate technologies that have been developed by others. The term $(A/n)$ in Equation (7) is a technical trick that captures this important aspect of technology adoption along the lines of Easterly et al. (1994) and Parente (1994).

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4 What prevents East European countries from relying solely on intermediate goods invented elsewhere? A reasonable assumption is that introducing a new good into the production process requires teaching workers how to use a new technology. Parente (1994) has presented an endogenous growth model in which firms adopt more advanced technologies, and subsequent to these adoptions accumulate expertise in those technologies.
Finally, we assume that the world frontier expands at the constant leading-edge technology growth rate

\[
\frac{\dot{A}}{A} = g,
\]

and that the labour force of the economy grows at a constant rate \( m \). 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, \( A \). The empirically implementable steady state output per capita \( y^* \) along the balanced growth path is given by

\[
y^*(t) = \left( \frac{sK}{m + g + \delta} \right)^{\alpha/\lambda - \alpha} n^*(t),
\]

or

\[
y^*(t) = \left( \frac{sK}{m + g + \delta} \right)^{\alpha/\lambda - \alpha} \left( \frac{\phi}{g} \right)^{1/\gamma} A^*(t).
\]

The economic interpretation of Equation (10) is straightforward. The model proposes two answers to the question of why different economies have different steady state income levels. First, the model emphasizes the importance of product variety, 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 (9) and (10) is similar to the basic Solow model. The term implies that countries investing more in physical capital will be richer.\(^5\) To understand the mechanics of the model, consider a country that opens up its economy to the rest of the world. We can model this as an increase in \( \phi \). According to (10), a higher value of \( \phi \) raises \( y^* \). Starting from steady state, the higher \( \phi \) causes the growth rate of \( n \) to be higher than \( g \) along the transition to the new steady state. Over time, however, the ratio \( A/n \) decreases, and therefore the growth rate of \( n \) returns to \( g \).

In other words, a policy change such as opening up the economy (interpreted as an increase in \( \phi \)) has a long-run level effect, but (just as in the original Solow model) no counterfactual long-run growth effect.\(^6\) This mechanism seems particularly

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\( ^5 \) In extensive sensitivity analyses of cross-country growth regressions, Levine and Renelt (1992) and Sala-i-Martin (1997) have shown that investment in physical capital is the most robust variable explaining cross-country growth differences.

\( ^6 \) An increase in \( \phi \) can also be thought of as a reduction of the ‘barriers to adoption’ stressed by Parente and Prescott (1994).
relevant for an assessment of the growth prospects of the economies in Eastern Europe, where pre-transition production volume was rather high, while the variety of products was rather limited. In order to embark on a path of transition and growth, a refined division of labour and an expanding variety are crucial for the future growth process.\footnote{The steady-state relationships (9) and (10) refer to the long run, i.e., to a time horizon when income per capita is not influenced by transitory shocks and transition dynamics. Naturally, all real world data incorporate deviations from their steady-state levels. The usual way to deal with this problem is to capture these transitory influences by adding fixed effects and appropriate trends.}

It is, however, not necessarily the case that time-varying overall variety measures are inextricably related to the innovation process inherent in semi-endogenous or endogenous growth models. In particular their correlations may also be consistent with an export-led growth model. To see this, imagine a Ventura (1997) view of the world in which growth is driven by physical capital accumulation and countries specialize according to comparative advantage. Ventura (1997) has conjectured that exporting manufacturers specializing in capital-intensive sectors may have played a key role in the rapid East Asian growth process. In a nutshell, the model maps out the following scenario: the East Asian countries fight the tendency of the rate of return to capital to fall by continuously shifting to a productive activity with a greater capital to labour ratio. Such an option would not be available in a closed economy. In Ventura’s (1997, p. 60) words: ‘As the capital stock grows, resources are moved from labour-intensive to capital-intensive industries, raising the demand for capital and sustaining the value of its marginal product. International trade converts an excess production of capital intensive goods into exports, instead of falling prices.’ In such a world, countries that accumulate more physical capital will grow relatively faster and will also move resources towards capital-intensive sectors at a higher pace. In other words, despite the diminishing returns to capital, the rate of return to capital was prevented from dwindling to very low values as a result of this specialization.

In order to understand the empirical relevance of these two alternative modelling approaches, a logical next step might be to take a look at the transition process in Eastern Europe. Was it the innovation process inherent in endogenous growth models or export-led growth that created growth in incomes? This will be the subject of the next sections.

3. Operational proxies for product variety across East European transition economies

The mechanism we are discussing does not lend itself easily to empirical testing, because the variety of goods in an economy is not directly measured by statistical agencies. The construction of a consistent international dataset for product variety therefore is a necessary first step in exploring the endogenous growth mechanism.
To explore this question, we have to pick a value of $n$ in actual economies, i.e., we have to measure the supply-side factor product variety. To get a measure of the variety of products across countries, the following two questions have to be addressed:

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

In the following empirical work, we adapt the most celebrated method for measuring product variety mapped out by Feenstra (1994), Feenstra and Markusen (1994) and Feenstra (2003, pp. 363–66) to cut through the statistical fog surrounding the product variety subject. These studies show how an exact measure of product variety can be constructed from a CES production function when the inputs enter non-symmetrically. The procedure considers two units of observations denoted by $s$ and $t$ representing two countries. Suppose that output $y_t$ in country $t$ is given by

$$y_t = f(x_t, I_t) = \sum_{i \in I_t} a_i x_t^{\sigma_i/(\sigma-1)} - \sum_{i \in I_t} \sigma_i b_i p_{it}^{\sigma_i},$$

(11)

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

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

(12)

where $p_{it}$ are the prices of the inputs and $b_i = a_i^\sigma$. Following Feenstra (1994) and Feenstra et al. (1992, 1999a, 1999b), we choose the set of intermediate products common to both countries as $I = I_s \cap I_t$. Relative product variety $\Delta PV_{st}$ is then defined as follows:

$$\Delta PV_{st} = \ln \left[ \frac{\sum_{i \in I_t} p_{it} x_{it}}{\sum_{i \in I_t} p_{it} x_{it}} \right].$$

(13)

What is the economic sense behind this result? To get a feeling for (13), consider, for example, the case where the set of inputs in country $t$ is larger than in country $s$, i.e., we have two sets $I_s = \{1, \ldots, N_s\}$ and $I_t = \{1, \ldots, N_t\}$ with $N_t > N_s$. Here, the
common set of products is $I = I$, and the denominator is 1. The numerator exceeds unity, indicating that product variety in country $t$ is larger than in country $s$. In the special case where all inputs enter symmetrically, the numerator in (13) simplifies to $\ln(N_t/N_s)$.

Our goal in computing $\Delta PV_{st}$ is to gauge how well different economies are performing in terms of product variety. Since this is crucial to our argument, let us explain the logic by way of a textbook example depicted in Table 1. The table shows the quantities of the two goods (1 and 2) produced in two hypothetical economies ($A$ and $B$) in a given year.

From these data and Equation (13), we derive product variety in country $A$ relative to country $B$ ($\Delta PV_{Country A/Country B}$) as

$$\Delta PV_{Country A/Country B} = \ln\left(\frac{150/100}{150/150}\right) = \ln(1.5) \approx 0.405. \quad (14)$$

In this simple example, $\Delta PV_{Country A/Country B}$ is positive and therefore (relative) product variety in country $A$ is greater than in country $B$. The reason is that country $A$ spreads its outputs thinly over both product categories. On the contrary, product variety in country $B$ relative to country $A$ ($\Delta PV_{Country B/Country A}$) is given by

$$\Delta PV_{Country B/Country A} = \ln\left(\frac{150/150}{150/100}\right) = -\ln(1.5) \approx -0.405. \quad (15)$$

The negative sign for $\Delta PV$ indicates that product variety in country $B$ is less than in country $A$. Admittedly, this is an extremely simple situation. In reality, product variety spans thousands of products, so the calculations inevitably become messier and more cumbersome. Nevertheless, the essential idea of the procedure is identical.

The procedure above is implemented using highly disaggregated trade data at the five-digit SITC level for 14 East European transition economies for the years 1993 to 2000, i.e., after the transformational recession. The most important advantage of

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8 We have not included the more peripheral countries of the former Soviet Union in which the preconditions were much less favourable to reform in terms of tradition, history and institutions, and where the commitment to reform has been half-hearted.
these data is that the classification of goods is consistent across countries. The classification distinguishes 1,473 commodities according to the Standard International Trade Classification (SITC Revision 2). The use of pre-established product categories makes it impossible to measure gains in product variety within any specific category and beyond the number of pre-established product categories. This puts a premium on the level of disaggregation in the data. One should expect a greater ability to differentiate between nations as the data become more detailed. All data were collected from the OECD database International Trade by Commodities Statistics (ITCS) Classification, Paris 2002 and are expressed in current US dollars. Given the data source, all proxies we derive pertain to exports (imports) to (from) OECD countries alone. As a caveat, the data have two distinct problems. First, we can only count pre-defined categories up to the maximum number of categories in the ITCS coding system. In other words, we cannot see new products arriving within an existing category and one cannot see the invention of truly new categories because there is a de facto fixed frontier in observable product variety. Second, the time series dimension of the data (eight years) is rather short. Nevertheless, we believe the topic to be of such economic and social significance, that a willingness to experiment with trade data is justified, especially since the most important goods are probably exported and/or imported. The consensus view is that trade liberalization in the sample countries has been considerable since the fall of the Berlin Wall. Once trade with the EU was liberalized, a huge re-orientation took place. Today the EU is, by far, the most important trading partner for most sample countries.

It is helpful to begin with a brief sketch of the data. Our first measure of product variety is export variety in country \( i \) \((i = 1, \ldots, 14)\) relative to the US \( \Delta PV_{EX} \). A country-by-country overview appears in Figure 1, which conveniently provides an idea of magnitudes. The first impression is that export variety in all countries under investigation (Belarus, Bulgaria, Czech Republic, Estonia, Georgia, Hungary, Latvia, Lithuania, Poland, Romania, Russia, Slovakia, Slovenia, Ukraine) is lower than in the US. Negative (positive) values for the index indicate lower (higher) product variety than in the US. Furthermore, the export variety measures yield a rather

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9 It would be preferable in principle to use national production data, but they are neither available at a sufficiently disaggregated level nor are the available data consistent across countries. In extensive discussions of quality and variety, Grossman and Helpman (1991) and Coe and Helpman (1995) and Bayoumi et al. (1999) focus on levels of investment in R&D at home and abroad. An obvious problem with this indirect approach is that the lag between R&D expenditures and the production of new varieties is potentially very long. Furthermore, it is also the case that many improvements in quality and variety can be realised without any R&D costs. In particular, increases in variety can occur through imitation, which involves little or no R&D spending.

10 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 (13). Cross-country trade data arranged by highly disaggregated product categories suffer from the fact that some countries drop category codes, redefine codes and/or add codes. These changes may sometimes induce statistical artefacts in the data that look like variety changes. At this level of detail, the trade statistics also contain errors that may distort the overall results. We have not minimized potential outliers (‘punch-in-mistakes’) as removing outliers would be completely arbitrary and data mining.
Figure 1. Relative export variety ($\Delta PV_{EX}$)
The proliferation of varieties is highest in the front-runners such as the Czech Republic, Hungary, Poland and Russia, and other EU-accession candidates. In contrast to this group, the degree of export variety is much lower in Romania and Slovakia, and Georgia posts the lowest ratio. An ‘eyeballing’ of the dataset immediately suggests that the variety gap is smaller in the more advanced transition economies with more Western-oriented industries. Nevertheless, keeping in mind the collapse of the Soviet Union and its implications for trade, the speed of re-orientation of trade in the transition economies is remarkable.\footnote{Boeri and Oliveira-Martins (2001), Brenton and Gros (1997), Havrylyshyn and Al-Atrash (1998) and Kaminski et al. (1996) discuss trade re-orientation during transition.}

One problem with indicators of product variety focusing solely upon export data is that, even when differentiated, inputs not produced at home are, in principle, available in other countries through trade. In other words, product variety in any country depends not only on exports but potentially also upon imports. We thus also calculate import variety relative to the US ($\Delta PV_{IM}$) and import and export variety ($\Delta PV_{EXIM}$). The results are given in Figures 2 and 3 below. It is apparent that import variety increased much faster than export variety for all countries in the sample. One of the first steps of the transition towards a market economy was the opening-up of trade, and therefore the lifting of existing restrictions to purchase of differentiated consumer and intermediate goods by domestic consumers and producers. Accordingly, we see an increasing spectrum of imported products.

Simple product counts are an alternative measure of changes in product variety over time. To obtain the alternative measure, we simply count all product categories that show recorded exports or imports.\footnote{A number of challenges may be levelled against the count proxy. For example, it is obvious that this simple measure of $PV$ is biased because each product category receives an equal weight.} The results are shown in Tables 2 and 3 below.

The results in Table 2 and 3 lead to some unsurprising conclusions. First, import variety is generally higher than export variety. Second, Belarus and Georgia displayed the least export variety while the Czech Republic, Hungary and Poland showed the most. Third, among the three Baltic states, Estonia turned out to be more diversified than Latvia or Lithuania. This picture of diversity will probably have a bearing on how the EU accession countries will cope with the additional adjustments required by the accession process itself and on what footing they will be able to participate in the enlarged European Union.

### 4. Strategies for empirical testing

In this section, we give directions along which the above hypotheses could be tested. Transition generates difficulties for the estimation for at least two reasons. First, transition marks a fundamental economic reorganization. Second, transition in Eastern Europe covers only a rather short time period. We thus opt for an
Figure 2. Relative import variety ($\Delta PV_{IM}$)
Figure 3. Relative export and import variety ($\Delta PV_{EXIM}$)
### Table 2. Export variety using the simple count-based measure

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*Note*: The maximum number of product categories is 1,473.

### Table 3. Import variety using the simple count-based measure

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<td>973</td>
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<td>1196</td>
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<td>Ukraine</td>
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<td>1049</td>
<td>1021</td>
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</tr>
</tbody>
</table>

*Note*: The maximum number of product categories is 1,473.
estimation method that exploits the full time dimension of the data by using all the information from a panel rather than just the time-averaged information from a cross-section.\textsuperscript{13} We are aware of the difficulties involved in this choice and subject our core results accordingly to a number of tests to gauge their robustness \textit{vis-à-vis} endogeneity, potential reverse causality and omitted variables. Given the limited transition time period, any conclusion is necessarily tentative.

Equation (9) provides a natural route to explore the semi-endogenous growth model. Consider writing (9) for one country, and then taking the ratio of that equation with the analogous equation for the US. We next obtain the relative per capita GDP of the two countries on the left, and obtain the relative savings/investments rates on the right, along with the relative number of product varieties. All variables are expressed relative to the US, since we designate the US as the technological leader. The relative investment share, $I_Y$, is added to the regressions to capture different per capita income levels arising from different levels of investment in physical capital. All data except $PV$ are drawn from the \textit{World Development Indicators 2002} database. The basic model for country $i$ and time $t$ presented above yields the following basic specification

\begin{equation}
Y_{it} = \alpha + \beta I_{Yit} + \gamma \Delta PV_{it} + \varepsilon_{it} \quad i = 1, \ldots, 14; \quad t = 1994, \ldots, 2000, \tag{16}
\end{equation}

where $Y_{it}$ is per capita GDP in country $i$ relative to the US in percent (PPP, international dollars), $I_{Yit}$ is defined as the share of fixed gross investment in GDP in country $i$ relative to the US, and $\Delta PV_{it}$ is product variety relative to the US. The sample period starts in 1994, i.e., after the initial GDP declines at the beginning of the 1990s. The initial period of transition from a planned to a market economy is too specific to be captured within a traditional growth theory framework.

The fundamental problems in estimating (16) involve unobservable factors that affect per capita GDP across countries, the potential endogeneity of the right-hand side regressors, spatial autocorrelation across countries and spurious regression.\textsuperscript{14} The first concern is that the unconditional estimates may be spurious, merely reflecting the joint impact of unmeasured variables on growth. To account for unobservable country characteristics (somewhat heroically assumed to be constant

\textsuperscript{13} We have used panel data estimation techniques because cross-section estimation has important limitations. The first, and most obvious, is that the number of observations it too small. On a more technical ground, the cross-section framework only permits a very limited treatment of the problem of estimation bias resulting from parameter heterogeneity and omitted variables. Islam (1995) and Caselli \textit{et al.} (1996) have demonstrated that panel data estimation techniques can be used to overcome some of the limitations of the cross-section approach, although panel data models are not immune to methodological issues.

\textsuperscript{14} The so-called spurious regression bias could be important, because the regressions in (relative) levels include fixed effects, which implies that the parameter estimates make use of the time series variation. To deal with this potential bias, we have performed panel cointegration tests using the procedure suggested by Pedroni (1999). The estimation results indicate that the series in (16) are indeed cointegrated (see the appendix).
within the sample period), we remove fixed effects by taking deviations from mean as we are unsure of what these factors are or because we lack the necessary data. The second concern is that the estimates may be plagued by bias arising from the endogeneity of the regressors. Endogeneity creates contemporaneous correlation between the regressors and the residual, thus creating biased and inconsistent coefficient estimates. We follow the usual tactic in such cases and rely on instrumental variables.\textsuperscript{15} The third concern is that the estimates are polluted by spatial correlation. Spatial correlation can be expected when growth is driven by stochastic shocks common to all countries in the sample. While one does not usually worry about cross-sectional correlation for randomly drawn samples at country level, these aggregate units are likely to exhibit cross-sectional correlation that has to be dealt with in a non-random sample of countries.\textsuperscript{16} Spatial dependence will generally not interfere with consistent parameter estimation, but standard techniques that fail to account for the presence of spatial correlation will yield inconsistent standard errors of these parameters.\textsuperscript{17} The spatial error component model is given by

\[
y_{it} = X_{it}'\beta + u_{it} \quad i = 1, \ldots, N; \quad t = 1, \ldots, T, (17)
\]

here \(y_{it}\) is the observation on the \(i^{th}\) country for the \(t^{th}\) time period, \(X_{it}\) denotes the \(k \times 1\) vector of observations on the regressors and \(u_{it}\) is the disturbance. In vector form, the disturbance in (17) is assumed to have random country effects, as well as spatially auto-correlated remainder disturbances, i.e.,

\[
u_t = \alpha + \varepsilon_t, (18)
\]

with

\[
\varepsilon_t = \rho W \varepsilon_t + v_t, (19)
\]

\textsuperscript{15} The method of instrumental variables is a powerful tool of econometrics because it allows consistent parameter estimation in the presence of correlation between explanatory variables and disturbances. Obviously, the performance of IV estimators depends upon the choice of instruments. Following the panel methodology of Arellano and Bond (1991), we use instruments based upon the lagged values of the explanatory variables in the regression. The assumption is that the explanatory variables are ‘weakly’ as opposed to ‘strongly’ exogenous; that is, they are assumed to be uncorrelated only with future errors. Please note that we have instrumented our measures of product variety, although these measures are constructed from highly disaggregated trade data and therefore unlikely to suffer from the endogeneity problems that plague traditional openness measures as discussed in Frankel and Romer (1999).

\textsuperscript{16} Even without common stochastic shocks, rising vertical specialization tends to accelerate the global propagation of shocks as industry-specific shocks are immediately transmitted to countries along the production chain. An extensive literature deals with this type of correlation. Space constraints do not permit investigation of econometric techniques in much detail. A textbook treatment of several of these issues can be found in Baltagi (2001), pp. 195–97.

\textsuperscript{17} Monte-Carlo simulations in Driscoll and Kraay (1998) indicate that the presence of even modest spatial dependence can impart large bias to OLS standard errors.
where \( \alpha' = (\alpha_1, \ldots, \alpha_N) \) denotes the vector of random country effects. \( \rho \) is the scalar spatial autoregressive coefficient with \( |\rho| < 1 \). \( W \) is a known \( N \times N \) spatial weight matrix; its diagonal elements are zero. \( W \) also satisfies the condition that \( (IN - \lambda W) \) is non-singular. Finally, \( \upsilon_q \) is assumed to be \( \text{INN}(0, \sigma^2_v) \). Parametric corrections for spatial correlation are possible only if \( W \) is known.\(^{18}\) Driscoll and Kraay (1998) conveniently present a simple modification of the standard non-parametric time series covariance matrix estimator that is robust to general forms of spatial and temporal dependence and heteroscedasticity. The model is identified by an \( l \times 1 \) vector of orthogonality conditions \( \mathbb{E}[Z_{it}'u_t] = 0 \), where \( Z_{it} \) is an \( l \times 1 \) vector of instrumental variables with \( l \geq k \). Thus, the Spatial Correlation Consistent (SCC) IV estimator simultaneously handles the issue of endogeneity of right-hand side variables. A further advantage of this flexible spatial dependence estimator is that calculation of the estimator is quite straightforward.

Table 4 gives the point estimates and the \( z \)-values allowing for dependence using the SCC estimator. What do these results reveal? For the investment share, highly significant coefficients indicate that investment in physical capital had a pronounced positive impact on growth. The results also suggest a statistically significant (albeit small) direct effect from export variety (\( \Delta PV_{EX} \)) on economic growth, implying that investment in physical capital does not carry all the information relevant for economic growth. While far from conclusive, the results show that

\[ \text{Table 4. The baseline regression model} \]

<table>
<thead>
<tr>
<th></th>
<th>(1)</th>
<th>(2)</th>
</tr>
</thead>
<tbody>
<tr>
<td>Constant</td>
<td>−0.08 (0.3)</td>
<td>−0.12 (0.7)</td>
</tr>
<tr>
<td>( \Delta PV_{EX} )</td>
<td>0.02 (0.2)</td>
<td>−</td>
</tr>
<tr>
<td>( \Delta PV_{EX} )</td>
<td>−</td>
<td>0.03 (2.3)</td>
</tr>
<tr>
<td>( IY )</td>
<td>3.72 (2.2)</td>
<td>2.85 (2.3)</td>
</tr>
</tbody>
</table>

\[ \text{Note: We have used a Bartlett kernel and select the bandwidth } m = 2. \text{ The qualitative results obtained are robust to changes in the window specification. The } z \text{-values in parentheses are based on the heteroscedasticity and spatial correlation consistent standard errors as discussed in the text. See text for data definitions and sources. The sample period is 1994–2000.} \]

\(^{18}\) The \textit{ad hoc} first-order contiguity matrix \( W \) that embodies spatial relationships of the adjacent-neighbour variety has been used most frequently in the literature. Unfortunately, economic behaviour is often more consistent with spatial weight matrices that allow neighbours to have differential impacts. Estimating \( W \), however, is impossible without imposing restrictions since the number of spatial correlations increases at the rate \( N^2 \), while the number of observations grows at rate \( N \).
Table 5. The augmented growth regressions

<table>
<thead>
<tr>
<th></th>
<th>(1)</th>
<th>(2)</th>
</tr>
</thead>
<tbody>
<tr>
<td>Constant</td>
<td>−0.09 (0.4)</td>
<td>0.10 (0.4)</td>
</tr>
<tr>
<td>(\Delta PV_{EX&amp;IM})</td>
<td>0.04 (0.5)</td>
<td>−</td>
</tr>
<tr>
<td>(\Delta PV_{EX})</td>
<td>−</td>
<td>0.14 (3.5)</td>
</tr>
<tr>
<td>(IY)</td>
<td>3.78 (2.3)</td>
<td>4.04 (1.7)</td>
</tr>
<tr>
<td>(EBRD)</td>
<td>0.41 (0.6)</td>
<td>3.47 (2.4)</td>
</tr>
</tbody>
</table>

Note: We have used a Bartlett kernel and select the bandwidth \(m = 2\). The qualitative results obtained are robust to changes in the window specification. The \(z\)-values in parentheses are based on the heteroscedasticity and spatial correlation consistent standard errors as discussed in the text. See text for data definitions and sources. The sample period is 1994–2000.

Export variety matters for growth.\(^{19}\) On the other hand, trade variety \((\Delta PV_{EX&IM})\) is not found to show a significant statistical association with growth.\(^{20}\)

In Table 4, we showed that export variety is beneficial for growth in transition economies. How robust is this result \(\text{vis-à-vis}\) omitted variables? If such conditioning variables have independent effects on growth, their omission from the regression specification will clearly lead to bias in the estimated coefficients. A related question is why then would exporting countries not switch to variety-intensive exports as a matter of course. Why do we see some countries that remain primary exporters? And this is where active government effort in being a midwife to industrialization may come in: privatization, infrastructure, property rights have been mentioned as some of the policy instruments. The countries that have successfully made the transition have had growth-augmenting ‘social capabilities’, i.e., a wide range of institutional and political requirements which are necessary for transition and growth. Introducing a wider range of explanatory variables beyond our fixed effects is one way to deal with this issue. Table 5 summarizes our attempts to deal with the effects of omitted variable bias.\(^{21}\) We have re-estimated our equations adding the aggregate EBRD transition indicator \((EBRD)\) to the list of regressors

\(^{19}\) This result is consistent with recent contributions looking at the impact of exports as conveyors of productivity (see, for example, Bernard and Jensen (1999) and Clerides \textit{et al.} (1998)).

\(^{20}\) These results may explain why the explanatory power of overall openness measures often turned out insignificant in previous growth regressions (see, for example, Havrylyshyn \textit{et al.} (1999), pp. 35–38).

\(^{21}\) We have also tried to add human capital indicators. The best measures would be in terms of the output of education, but due to the difficulty of obtaining such measures, input measures have to be used. Unfortunately, available proxies such as school enrollment rates, average years of education and the proportion of the labour force that has received primary, secondary or tertiary education do not vary over time (see, for example, the \textit{World Development Indicators} World Bank). The use of fixed effects, however, prevents the analysis of such variables that do not change much over time.
which provides a deliberate policy measure which differs across countries and time. Another loose justification comes from the fact that the aggregate EBRD transition indicator is the most frequently used transition indicator. The EBRD’s indicator ranges from 1 to 4+, with 1 representing conditions unchanged from those prevailing in a centrally planned economy with dominant state ownership of means of production and 4+ for conditions in an advanced market economy. In the empirical work below we assign a value of 1/3 to a ‘+’ sign and −1/3 to a ‘−’ sign. The aggregate indicator covers key areas of structural reform including privatization and restructuring, foreign trade liberalization and competition and financial markets reform that have an independent impact upon growth.

The main feature is that the overall pattern of results remains. Table 5 invites two major conclusions. First, the signs of the coefficients are in conformity with the theoretical model. The coefficients of the relative investment share and the product variety measures are still significant. Second, economic policy as reflected in the aggregate EBRD transition indicator investment has significantly contributed to the productivity recovery in Eastern Europe, all else being equal. In other words, government policies in a wide range of areas is important in explaining both the time and cross-sectional dimension of output paths during the transition phase.

With 1,473 commodity groups, apart from estimating the growth equation for the entire list of products, the data can also be broken down into subgroups. As a further step, we therefore employ the added bonus of the disaggregate dataset to test whether Ventura’s (1997) model of outward orientation and integration with the world economy captures the essence of the recent growth experience in Eastern Europe. To begin with, we have divided the trade sectors into the following two subgroups or sub-sectors: Labour-intensive goods (n = 528; SITC 26, 61, 63, 64, 65, 66, 69, 81, 82, 83, 84, 85, 89) and capital-intensive goods (n = 666; SITC 1, 51, 52, 53, 54, 55, 57, 58, 59, 62, 67, 68, 7, 87, 88). This is the route that we take in the empirical implementation of the hypothesis that sustained growth of open economies is accompanied by the gradual shift in production towards more capital-intensive sectors without incurring diminishing returns. The pattern of the disaggregated variety measures and the catching-up for both classes of goods can be seen in Figures 4 and 5. The product variety measures presented below again suggest that the dynamics of product variety across countries is uneven and diverse.

The results of the disaggregated statistical horse race are summarized in Table 6 below. What do we learn? The key message of Table 6 is remarkable: the poor fit found here for the labour-intensive products merely highlights an insight – all

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22 We have not added an indicator of initial conditions because Krueger and Ciolko (1998) have shown that government reform is to some extent determined by initial conditions leading to multicollinearity problems. Favourable starting conditions might generate better results with respect to growth, making it easier to accept the negative effects of reform, resulting in faster and more encompassing reform.

23 This result is consistent with the empirical evidence for Korea, Taiwan and Japan and the OECD countries in Feenstra et al. (1999a, 1999b) and Funke and Ruhwedel (2001).
Figure 4. Relative export variety of capital-intensive goods ($\Delta PV_{EX-CAP}$)
Figure 5. Relative export variety of labour-intensive goods ($\Delta PV_{EX-LAB}$)
product variety is not equal and the results are at odds with the spirit of the semi-endogenous growth model presented above. On the contrary, Ventura’s (1997) model of outward orientation and integration with the world economy ‘trumps’ the semi-endogenous growth model and therefore seems to capture the essence of the recent growth experience in Eastern Europe.

This evidence is robust to a number of sensitivity analyses. To determine the extent to which lagged variables influence the trends in this paper, we have re-run the estimations above introducing dynamics into the specification. It is somewhat reassuring that the coefficients estimated using lagged PV-measures are generally similar to those estimated with contemporaneous variety proxies.

5. Summary and implications

Product-level trade data provide a completely new dimension for testing and thinking about alternative growth theories. Economists have long speculated that increasing product variety may generate higher per capita incomes and economic growth. Yet, despite significant policy implications, systematic empirical evidence on the actual magnitude of product variety is just beginning to emerge. Using contemporary econometric methods and data which provide a new dimension for testing the implications of alternative growth models, our empirical results give us no cause to question the role of export variety fostering economic growth of the East European transition economies. The result that export product variety in capital-intensive industries and investment are spearheading the growth process is consistent with Ventura’s (1997) neoclassical export-led growth model as sketched above. If this interpretation is correct, conventional macroeconomic approaches

<table>
<thead>
<tr>
<th>Table 6. Estimation results for capital-intensive versus labour-intensive goods</th>
</tr>
</thead>
<tbody>
<tr>
<td>(1)</td>
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<tr>
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</tr>
<tr>
<td>$\Delta PV_{EX-CAP}$</td>
</tr>
<tr>
<td>$\Delta PV_{EX-CAP}(-1)$</td>
</tr>
<tr>
<td>$\Delta PV_{EX-LAB}$</td>
</tr>
<tr>
<td>$\Delta PV_{EX-LAB}(-1)$</td>
</tr>
<tr>
<td>$IY$</td>
</tr>
<tr>
<td>$EBRD$</td>
</tr>
</tbody>
</table>

*Note:* We have used a Bartlett kernel and select the bandwidth $m = 2$. The qualitative results obtained are robust to changes in the window specification. The z-values in parentheses are based on the heteroscedasticity and spatial correlation consistent standard errors as discussed in the text. See text for data definitions and sources. The sample period is 1994–2000.
are misspecified by the exclusion of a varieties effect. An overarching issue is that although we now have more knowledge of product variety differences across countries, our understanding of the mechanism that creates these differences is still poor. We hope that our findings, which are based on a highly disaggregated empirical approach, encourage researchers to pursue these issues further.

References


Appendix

Panel cointegration tests

There are different methods for testing cointegration in panels. Initially developed panel cointegration tests applied panel unit root tests directly to the residuals from an Engle Granger type two-step methodology. But the recent opinion in the literature suggests that the test statistics using this approach would be biased towards accepting stationarity due to lack of exogeneity of the regressors and the dependency of the residuals on the distribution of the estimated coefficients. For these reasons it is important to have a test procedure for cointegration which is robust to the presence of heterogeneity in the alternative. We have therefore applied the Fully Modified Ordinary Least Squares (FMOLS) estimator developed by Pedroni. Pedroni (1999) has suggested seven different test statistics for testing the null hypothesis that for each member of the panel the variables are not cointegrated against the alternative that for each member of the panel the variables are cointegrated. Pedroni (1999) discusses the construction of seven panel cointegration statistics, four based on pooling along the within-dimension and three based on pooling along the between-dimension. Within the first category, three of the four tests involve the use of non-parametric corrections. The fourth is a parametric ADF-based test. In the second category, two of the three tests use non-parametric corrections while the third is again an ADF-based test. The test statistics given in the first category are based on estimators that effectively pool the autoregressive coefficient across different countries for the unit root tests on the estimated residuals, while the test statistics given in the second category are based on estimators that simply average

Table A1. Panel cointegration tests

<table>
<thead>
<tr>
<th>Panel Cointegration Test</th>
<th>Test statistic</th>
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<tbody>
<tr>
<td>Panel $\nu$-statistic</td>
<td>0.86</td>
</tr>
<tr>
<td>Panel $\rho$-statistic (non-parametric)</td>
<td>−2.48</td>
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<tr>
<td>Panel $t$-statistic (parametric)</td>
<td>−2.04</td>
</tr>
<tr>
<td>Panel $t$-statistic (parametric)</td>
<td>−0.85</td>
</tr>
<tr>
<td>Group $\rho$-statistic</td>
<td>−3.85</td>
</tr>
<tr>
<td>Group $t$-statistic (non-parametric)</td>
<td>−2.63</td>
</tr>
<tr>
<td>Group $t$-statistic (parametric)</td>
<td>−5.79</td>
</tr>
</tbody>
</table>

Note: Export variety has been used as a measure for PV. All test statistics are standard normally distributed $N(0,1)$ and therefore the critical value is 1.69. The tests include a constant and heterogeneous time trends in the data.
the individually estimated coefficient for each country. The estimation results for Equation (16) are given in Table A1.

The table shows that the null hypothesis is rejected in five out of the seven cases. Although we do not get a definite answer, we conclude that there is a cointegration relationship.