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Multivariate Cointegration Analysis of Aggregate Exports: Empirical Evidence for the United States, Canada, and Germany

by

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Abstract:

A shortcoming of most empirical studies on aggregate exports is their exclusive focus on the demand side. Moreover, the effect of globalization is often neglected leading to implausibly high income elasticities. This paper models export demand and supply simultaneously and incorporates a new proxy for globalization. Owing to the non-stationarity of the data, the vector error correction model is the appropriate econometric framework. Using the Johansen procedure, two cointegration relationships are found and identified as export supply and demand. Overidentifying restrictions derived from economic theory are tested. Finally, after checking for weak exogeneity, a parsimonious partial model is presented and the adjustment paths of the endogenous variables are discussed.

Keywords: cointegration analysis; Johansen procedure; export demand and supply; trade elasticities

JEL Classification: F31; F41

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Contents

1. The Estimation of Aggregate Exports: Question Marks in a Seemingly Explored Field

The econometric analysis of aggregate exports has a long and well established tradition in applied international economics. The availability of long and relatively reliable foreign trade data also makes it one of the more rewarding fields of time series econometrics. Trade equations are of considerable interest for macroeconomic policy as they give an indication on whether trade deficits of a single country can be addressed by devaluation (elasticity optimism) or not (elasticity pessimism). Knowledge on the determinants of aggregate exports is interesting in its own right because small open economies earn a major part of GDP in exports (e.g. Canada: about 40 percent). In the short run empirical estimations of exports can substantially increase the quality of business cycle forecasts which are of interest both for the private sector and for macroeconomic policy. In the long run, the foreign trade potential of a region, a country or larger economic areas may also be gauged with the help of econometric analysis and can be a support tool in the planning process for localization strategies of firms and public policy on infrastructure, trade, taxes, regulations and other structural issues.

The obvious interest in aggregate exports and the „success of the standard trade model" (Bayoumi 1998: 1), the mainstream tool of analysis, may explain the existence of a very large empirical literature. Influent surveys summarizing empirical results of studies published during the respective 10 to 20 years before comprise Stern et al. (1976), Goldstein and Khan (1985), and Sawyer and Sprinkle (1999).[1] Recent contributions either focus on comparisons of long-run and short-run trade elasticities between countries (e.g. Senhadji and Montenegro 1999, Murata et al. 2000, Hooper et al. 2000, Lapp et al. 1995), try to exploit the

[1] A more in-depth discussion of earlier results is provided in Strauß (2002a).

information of bilateral trade data using panel data techniques in order to still increase the number of observations and thus the statistical reliability of results (Bayoumi 1998), discuss the evolution of trade elasticities over longer time horizons (Marquez 1995) or deal with the question whether aggregate price-elasticities in foreign trade correctly represent those observed in markets for single goods or sectors (e.g. Meier 1998).[2] The focus in this study is on exports on the highest level of aggregation. International comparison plays a role (although not the dominant one) as three countries (the Unites States, Canada, and Germany) are analyzed during the same time period (1975–2000) and using the same empirical framework.

Yet the main motivation for this paper stems from methodological and specification issues. Actually, the existence of a standard framework of analysis does not mean that no controversies remain. There is such a huge variance in estimated elasticities that this cannot only be due to differences between the countries, periods and data types analyzed but must also have to do with the methods at hand. A shortcoming of most empirical studies on aggregate exports is their exclusive focus on the demand side. It is assumed that the price-elasticity of supply is infinite at least in the long run, which seems to exempt the researcher from dealing with export prices and quantities simultaneously. Based on the assumption of a horizontal long-run export supply curve, a strange dichotomy between empirical studies of export quantities on the one hand (supposed to primarily reflect the behavior of demand) and empirical studies of export prices (supposed to reflect the behavior of suppliers) on the other hand has developed. There may be theoretical reasons in favor of a horizontal long-run supply curve but it is a challenging task to find out whether this assumption is backed by the data or not. This was already done by Goldstein and Khan (1978) who came out with strong evidence of upward-sloping supply curves for

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2 The suspicion of a „downward bias“ in aggregate price elasticities due to aggregation was first formulated in an influent article by Oructt (1950).

eight OECD countries. Yet since that time not only the observation period has changed. The taking into account of the non-stationarity of most macroeconomic time series and the emergence of cointegration analysis are likely to call the earlier results into question.

Moreover, Goldstein and Khan impose the nature of the adjustment mechanism *assuming* that demand imbalances provoke quantity reactions whereas supply imbalances lead to price reactions. Browne (1982) argues that this might be an implausible assumption for small open economies (SME) and suggests exactly the opposite reaction pattern. As countries are never pure SMEs in reality, it is conceivable to observe both price and quantity changes in response to each type of imbalance (demand and supply). It would be desirable to have an empirical framework allowing for various types of adjustment in order to test which one fits the data best.

Goldstein and Khan's promising approach of distinguishing export demand from export supply is pursued in this paper within a framework of multivariate cointegration analysis that allows for a simultaneous determination of all system variables found to be endogenous by appropriate statistical tests. It seems straightforward to me to model export quantities and export prices jointly as they belong to the same bundle of goods and result from the interaction between the two sides of the same market (demanders and suppliers of "Home's" aggregate exports). Doing so not only promises to yield richer information on the interactions between export volumes and prices as well as potential feedbacks on foreign economic activity and domestic prices. It also paves the way to a better economic understanding of the long run. As Johansen and Juselius (1992) point out, there may be more than one cointegration relationship in systems of three and more variables and neglecting this issue by proceding to single-equation estimations involves the danger of finding (and wrongly interpreting) linear combinations of economic long-run relationships rather than the relationship of interest itself. Moreover, ignoring the potential endogeneity of explanatory variables means throwing away information and leads to inefficient estimates (Johannsen 1992a). The empirical strategy of this paper therefore consists of taking into account all possible feedbacks, finding the number of cointegrating vectors and of identifying them in a way to find the long-run relationships of export demand and export supply. Once this is done, the number of feedbacks is reduced and the equation system simplified by statistical tests for weak exogeneity and tests for long-run elasticities taking on specific values predicted by economic theory. These tests take the form of over-identifying restrictions on cointegration vectors and loading coefficients.

Another shortcoming in the literature this study tries to address is the missing distinction between economic growth and globalization as distinct sources of export growth. On the microeconomic level it makes a big difference whether a French car retailer orders more German cars because disposable household income in France rises or whether German motors and other vehicle parts are shipped to France, assembled, sent back to Germany where the bodywork is sprayed, and ultimately sold all over the world. The first example points to the link between exports and foreign aggregate production or income, the second one to the link between exports and the ongoing international division of labor. If this second feature is not taken into account explicitly by an additional variable, one ends up with foreign-production elasticities of exports much higher than one. It is argued in the next section that this does not make much sense from a theoretical viewpoint. All in all, the innovation of this paper consists of the combination of taking into account both demand and supply of exports, the analytical distinction between globalization and foreign production as well as modern cointegration analysis.

The study is organized as follows: Section 2 recapitulates the theoretical determinants of export supply and demand, section 3 describes the data sources and methods of calculation of the regressors, section 4 provides an extensive analysis of the nature of the time series used. Section 5, the core part, presents a full I(1) analysis in the spirit of Johansen and Juselius (1992), section 6 presents the partial model, and section 7 gives some summary remarks.

2. The Theoretical Determinants of Aggregate Export Demand and Supply[3]

Foreign economic activity, e.g. GDP or another production measure, is a major determinant in all standard empirical export models. This variable is a proxy for either real disposable income of foreign households or the volume of production of foreign firms depending on whether the demand for exports of the country considered („Home“) is theoretically derived from utility maximization by foreign households or cost minimization by foreign production firms using the country's exports as an input in their final goods production.[4] Given the very high share of intermediate and investment goods in German exports, the latter framework is chosen in this study following Sandermann (1975) and Clostermann (1996 and 1998).[5] The demand for aggregate exports derived in this way depends on the level of aggregate production in the trading partner countries and of the relative price of Home's exports compared to all other inputs (foreign capital, labor, intermediate goods), i.e. the real effective exchange rate of Home's currency using Home's export price level (rather than producer prices) as the correct price index in adjusting the nominal exchange rate for inflation differentials. As shown in Strauß (2001), in this framework estimated coefficients of foreign industrial production of 1 are consistent with constant returns to scale in the foreign production function. Foreign production

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3 An extensive discussion of the theoretical determinants is given in Strauß (2002a).

4 This framework also covers exports of final goods if the latter are considered as inputs to the service production by foreign wholesalers or retail salers (Clostermann 1998: 204).

5 Following these authors, Strauß (2001) derives export demand from a foreign production function with constant elasticity of substitution (CES) and applies this framework to a cointegration analysis of Euroland's aggregate exports of goods and services.

coefficients above 1 are consistent with increasing returns to scale if and only if export demand is price-elastic. Yet if the price elasticity of exports (as measured by the long-run increase in export volumes following a 1 percent real depreciation of Home's currency) is comprised between 0 and 1 – which is frequently observed for many countries – a foreign production coefficient above one points to diminishing returns to scale. However it is difficult to imagine how in growing economies with technological progress doubling all inputs should yield less than twice the initial output in the long run. For theoretical reasons one would therefore expect a long-run elasticity of exports with respect to foreign production of close to one (or even slightly below 1 in case of price-inelastic exports allowing for increasing returns to scale). Estimates for parameters of foreign economic activity are very often found to be higher than 1 in standard demand models which is at odds with these theoretical considerations.

The picture becomes clearer if one realizes that the simple export demand derived from the CES production function cannot explain the totality of growth in aggregate exports as it does not contain all the other conditions that favorably influenced export quantities in the past decades. The successive abolishment of tariff barriers under the GATT and then the WTO boosted exports far beyond what can be captured by production figures. Falling transportation costs were another factor whose beneficial effects on the volumes of international trade and investment are well explained in the theoretical literature (e.g. Krugman 1980, Kleinert 2001). Furthermore, the derived export demand function stresses the role of *relative* prices of European exports neglecting that exporters all over the world saved huge amounts of costs during the last twenty years by "slicing up" the production chain ("outsourcing"), by buying inputs internationally rather than locally ("global sourcing"), and by creating networks of multinational firms, which usually leads to growing trade in intermediate inputs (Kleinert 2000). The integration process can also be understood as the opening-up of a country or the further intensification of trade between two economies producing heterogenous products with increasing returns to scale (at the firm level, not necessarily at the macroeconomic level) for consumers with a love for variety: this always leads to a greater variety for each consumer, lower product prices because of higher output per variety and higher intra-industry exports relative to GDP, i.e. a higher degree of openness of the economy (Helpman and Krugman 1985: 141f.).

As it is not my purpose to explain the globalization process *per se* but to focus on its implications on the demand for (and supply of) German, U.S. and Canadian exports, I will present an empirical proxy for globalization in the next section. This additional variable will be able to fill the gap between growth in international production and growth in national exports thus making the assumptions of the CES model compatible with real world data.

As to the supply side, the perhaps most prominent empirical contribution allowing for a non horizontal aggregate export supply curve is Goldstein and Khan (1978). In their model, the driving force is the relative profitability of the exporting activity. As exporting firms are generally free to supply their goods and services either on the domestic market or abroad, they allocate their output on both markets according to the price signals received. The relative profitability is therefore defined as the ratio between the average price received on export markets and the one received on the domestic market. At the difference to the demand for exports, *domestic* prices enter the analysis as soon as the supply side is discussed. A rise in the relative profitability of exporting will lead to an increase in exports. However, this increase is not infinitely large as suggested by the widespread assumption of a horizontal export supply. As exporters compete with producers of non-tradables for a finite stock of capital, labor and other resources and as non-tradables may well be produced with different factor intensities than exportables, there is a cost in increasing export supply even in the long run. This gives rise to an upward-sloping export supply curve and the claim that the increase in exports depends on the strength of the export price increase.

When absolute levels of exports (rather than export shares in GDP) are estimated, as is the case here, the trade-off between producing more exports and producing more non-tradables is of course weakened by an increase in the productive capacity of the economy. This is why according to Goldstein and Khan (1978) trend (or potential) output should act as a positively-signed shift variable for the export supply curve. They find empirical evidence in favor of the relevance of potential output in export supply but there is at least one reason to be skeptic about this finding: As the analysis by Goldstein and Khan (1978) was published before the groundbreaking article by Nelson and Plosser (1982) on the non-stationary nature of most macroeconomic time series and the development of modern cointegration analysis, it specifies all variables in levels rather than growth rates. The existence of a long-run economic relationship (cointegration) is therefore assumed rather than tested for. A violation of this assumption would imply the danger of spurious regression between trend output and exports as both have a positive deterministic trend component. This is why the validity of their empirical results is questionable.

In later approaches taking into account the influence of imperfect competition, the role of strategic interaction and pricing to markets is stressed giving prices set by foreign competitors also a role to play. The theoretical foundations are laid by Dornbusch (1987) and Krugman (1987), empirical applications are given e.g. by Knetter (1993) and Bayoumi (1998). Moreover, the forces of globalization described above are not confined to the demand side of exports as optimizing firms have a strong incentive to delocalize production to low-cost places when transportation and other distance costs fall over time.[6] Just as

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6 Although it is not straightforward to show theoretically that the „trade-saving" effect of foreign direct investment (resulting from delocalization of production of a good to the country to which the same good was formerly exported) is smaller than the „trade-

potential growth, the ongoing international division of labor shifts the supply curve outward thus allowing higher export supply at given prices. Alternatively, at given export supply levels, prices are falling.[7]

To recapitulate, the volume of exports demanded (X) is determined by the level of aggregate foreign production (Y^*), the real effective exchange rate (E) and a proxy for globalization (F),

[1] $X^d = g(Y^*, E, F)$, where $E = W \cdot Q / P^*$,

i.e. the real effective exchange rate of the domestic currency is defined as the nominal effective exchange rate, W (in units of the foreign currency per national currency unit), adjusted for inflation differentials by the ratio of the domestic export price level and the average foreign producer price level.

In turn, export supply depends on relative profitability of exporting (RP), the price level of foreign goods and services in national currency units (P^* / W), potential output ($\tilde{Y}$) and globalization, i.e.

[2] $X^s = h\left(RP, P^*/W, \tilde{Y}, F\right)$, where $RP = Q/P$.

To yield a parsimonious equation system, it is necessary to sort out the components of the diverse price and exchange rate measures mentioned in [1] and [2]. The next section presents the data used in the empirical analysis and the way in which the regressors are constructed. It will show that three time series for prices are enough as soon as one chooses a unique currency framework (all variables in national currency units or all variables in foreign currency units).

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creating" effect of foreign direct investment (resulting, *inter* alia, from exports of specialized inputs), this happens to be the case empirically.

7 Helpman and Krugman compare the integration process between two countries with the opening-up of trade between formerly autarkical economies when stressing the role of increasing returns to scale at the firm level. They point out: „Moreover the [...] factors that lead to an autarky relatively lower price of manufactured products in the large country [...] make the post-trade relative price of manufactured products lower than [...] prior to trade." (Helpman and Krugman 1985: 156).

3. Time Series Used for Estimation

In this section, 12 time series are introduced and discussed. Their symbols and definitions are summarized in Table 1. In the end, only the six variables $\{x, q_{NCU}, y^*, p^*_{NCU}, p_{NCU}, f\}$ will be used in the econometric estimations. Yet the price variables, when transferred into another currency, are composed measures as are the nominal (w) and the real effective exchange rates (e). In this case, it is important to look at the time series properties, especially on the degree of integration, of the underlying series in order to detect hidden cointegration relationships.[8]

Table 1: Acronyms for Variables

Acronym	Meaning
x	log of index of real exports of goods and services
y^*	log of index of export-weighted foreign industrial production
$\tilde{y}$	log of index of potential output
f	log of index of world trade intensity (ratio of real world merchandise exports to real world GDP, both in U.S. dollars)
w	log of index of nominal effective exchange rate (in units of the representative foreign currency basket per unit of domestic currency)
e	log of index of real effective exchange rate (w adjusted for differential between domestic export and foreign producer price index)
p_{NCU}	log of domestic producer price index (in domestic currency units)
q_{NCU}	log of domestic export price index (in domestic currency units)
p^*_{NCU}	log of trade-weighted foreign producer price index (in domestic currency units)
p	log of domestic producer price index (in units of the representative foreign currency basket)
q	log of domestic export price index (in units of the representative foreign currency basket)
p^*	log of trade-weighted foreign producer price index (in units of the representative foreign currency basket)

8 For example, if the nominal exchange rate and the level of foreign prices in foreign currency units were found to be I(1) but the level of foreign prices in *domestic* currency units to be I(0), this would probably stem from a cointegration relationship between foreign prices and the exchange rate with coefficients (1,1).

Beginning with aggregate exports, I now give a closer look on the time series used for estimation. As the focus of this study is on the structural long-run and cyclical behavior of aggregate exports, I am looking for the broadest possible measure of a country's transboundary sales on goods markets. Therefore, real exports of goods and services from the System of National Accounts (SNA) of the respective country are taken as the best proxy for the overall export volume of the United States, Canada, and Germany, respectively.[9] For the sake of consistency, the deflator of exports of goods and services (taken from the same SNA source) represents the suitable proxy for export prices. Nominal and real exports are generally denominated in billions of national currency units and are growing over time with a deterministic (and possibly even a stochastic) trend. To avoid the appearance of intercepts only due to differences in scales of the variables used and to make a meaningful use of linear estimation techniques, the series are first brought to a value equal to 100 in the base year, then the natural logarithm (indicated by small letters) is taken. These operations apply to all time series discussed. According to different choices of the base year in the three countries considered, all variables used in estimations equal 100 in 1996 (average of the four quarters) for the United States, in 1992 for Canada, and in 1995 for Germany. In the following, x stands for real exports of goods and services and q_{NCU} for the deflator of real exports in national currency units. Whenever prices are in national currency units (NCU), the corresponding variable gets a subscript "NCU". The source for both export volumes and the

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9 With national accounts data being themselves only an estimation based on the aggregated firms- and household-specific „hard" data, one should normally prefer statistical data of a more primary nature. Customs trade data would be a suitable choice if one were interested in goods trade only, as for most OECD countries both values and volumes of merchandise trade are available in a satisfactory quality. However, as soon as services are part of the analysis, the aspect of data consistency between goods trade, services trade and total trade becomes important. In some countries (such as Germany), balance-of-payments data on services exports are more reliable a measure than the corresponding SNA sources but they only exist in nominal terms and therefore cannot replace SNA data either.

deflator is OECD (2001a). The symbol $y*$ stands for the index of industrial production in the most important trading partner countries. This index is a weighted arithmetic average with the average share in total exports of each trading partner during 1996–98 serving as a fixed weight. Two comments should be given to answer the questions "Why *industrial* production?" and "Why only the *most important* partner countries?" As to the first question, industrial production may be questioned as a representative proxy for total economic activity as a considerable part of total exports, e.g. consumer goods, is directed to households. However, industrial production does a good job in empirical applications; due to its stronger ups and downs, it shapes the macroeconomic cycle and thereby approximates overall economic activity remarkably well (Filardo 1997 and Döpke 1995) even in the wake of tertiarization.[10] The higher variance in industrial production provides it with a higher explanatory power compared to GDP data. Moreover, the variable is more welcome for forecasting purposes as data on industrial production are published with a smaller delay than their SNA counterparts and are available on a monthly (rather than quarterly) frequency. As to the second question, the choice of partner countries is heavily constrained by the availability of quarterly data from 1975 to 2000. This rules out the taking into account of, say, Central and Eastern European countries which have become important for Germany's exports during the second half of the nineties. Furthermore, the trading partner should have a certain weight in Germany's or America's foreign trade because otherwise the marginal utility of a more representative country sample could be outweighted by the potential marginal cost of aggregating over countries with different statistical standards (e.g. OECD versus non-OECD countries). Given these constraints, the number

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[10] Despite the declining share of goods producing industries in total employment, the share of these industries in total real economic activity was roughly constant during the second half of the twentieth century. The slight decline in the more recent past is partly due to outsourcing activities which only make visible in the SNA what had been industry-dependent services since ever (Filardo 1997: 76).

of partner countries is smaller for Germany (19) than for the United States (33).[11] Yet the degree of representation is high in each case: it amounts to 76 percent for Germany and 87 percent to both the United States and Canada.[12] The primary sources for y^* are OECD (2002a) for industrialized countries and IMF (2001) for emerging economies. Potential output of a country, expressed by the symbol $\tilde{y}$, is considered to be an important determinant of the volume of exports supplied by this country according to Goldstein and Khan (1978). I take the potential GDP series for the United States, Canada, and Germany from the OECD Economic Outlook Database (OECD 2001b) as a proxy.[13] As the series is only available at semiannual frequency, the missing values for the first and third quarters of each year are filled in by linear interpolation.

The next variable on the list in Table 1, f, is the "globalization" variable. As it is the ratio between real world merchandise exports and real world GDP, I call it the trade intensity of world production. The rise in f during the past twenty years is just a proxy for all the trade-enhancing phenomena outside an increase in production and an improvement of a country's price competitiveness and may stand for falling transportation costs, a lowering in tariffs and other trade barriers, the increasing importance of networks within multinational firms (MNF), higher cost-awareness by firms in industrialized countries which switch from regional or national input purchasing systems to outright global sourcing strategies, the integration of new trading partners into the world trading system (such as Central and Eastern European Countries or China), the slicing-up of the value-added chain within production firms leading to a substantial increase in

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[11] Germany's exports are more concentrated on other OECD countries than US exports.

[12] Canada is treated somewhat differently than the other two countries (see below).

[13] For Germany there is a break in the series as West Germany's potential GDP is taken for the period 1975 to 1990 and potential GDP of „whole" Germany from 1991 to 2000. Chaining the series together is impossible as the OECD does not provide an overlapping year with potential GDP for both territories. Although this break should be dealt with explicitly, I do not discuss it any further because potential GDP finally turns out to be of no long-run interest in the cointegration analyses below.

vertical trade (transboundary flows of specialized inputs and semi-finished goods) etc. Clearly this variable only makes sense in an empirical model in which national exports rather than exports of a whole continent or even the world are determined because with world trade at the right hand side of the export equation, there would be an endogeneity problem. For Germany and Canada, there is no such problem as these countries are sufficiently small in a global context. As the exogeneity argument is less convincing for the United States, f is computed *ex* U.S. exports (in the numerator) and *ex* U.S. GDP in the denominator in order to obtain the trade intensity of production in the world outside the United States. In this way, the left hand side and the right hand side of equations mutually exclude each other and there should remain no endogeneity problem. However, as the volume of world trade is a highly procyclical variable and y^* is in the set of variables, one should make sure to avoid a multicollinearity problem when estimating the system. In the long run, such a problem does not exist as world trade is normalized by a measure of world economic activity such that only the "excess" increase in trade remains. Yet in the short run, the turning points, accelerations and decelerations of foreign industrial production and the trade intensity of world production might well coincide. As f is intended not to reflect business-cycle movements but the long-run structural phenomena mentioned above, it is thought to replace the linear deterministic trend sometimes used in empirical studies (e.g. Döpke and Fischer 1994 and Strauß 1998) and suggested by econometric textbooks on the practical modeling of foreign trade flows (e.g. Whitley 1994) rather than to "compete" with y^*. To avoid multicollinearity in the short run, world trade intensity is smoothed over several years. Anyway only annual data are available for world GDP. Starting at the annual frequency, first a centered three-year moving average of the ratio of real world trade to world GDP is computed. This ratio is then converted into quarterly frequency with the "quadratic match average" option implemented in EViews 4 to make f operational for estimations based on quarterly data.

As w and e, the next on the list of Table 1, are combined variables, I first turn to p_{NCU}, the index of domestic producer prices (PPI) in national currency units which is taken from OECD (2002a). The PPI is a proxy for either sales prices of tradeables on the domestic market or of marginal costs of domestic production, the theoretical variables which matter for export supply in the model by Goldstein and Khan (1978) and the one by Dornbusch (1987), respectively. The index of foreign producer prices in units of the representative foreign currency, p^* (last line in Table 1) is a series combined from national PPIs without currency conversion. The source is OECD (2002a) for industrialized countries and IMF (2001) for all other countries. All indices are brought to 100 in the base year (e.g. 1995 for Germany), then compound as a weighted geometric average. The weights are the shares of the respective partner countries in the total goods supply competing with Home's (here: Germany's) goods supply around the world.[14] Geometric rather than arithmetic weighting is used in order to make the index less sensitive to (small) trading partners with high inflation. The weights of the 17 and 30 trading partners (for Germany and the U.S., respectively) are taken from Deutsche Bundesbank (1998) and Board of Governors (1998).

The variables p^*_{NCU}, p and q involve both price indices and effective exchange rates as each of these series is converted from the original currency into another currency. The domestic PPI and the domestic export deflator, both expressed in units of the representative foreign currency unit, are calculated as ($p_{NCU} + w$) and ($q_{NCU} + w$). Thereby w, the nominal effective exchange rate of the national currency (as a value concept, i.e. in units of the representative

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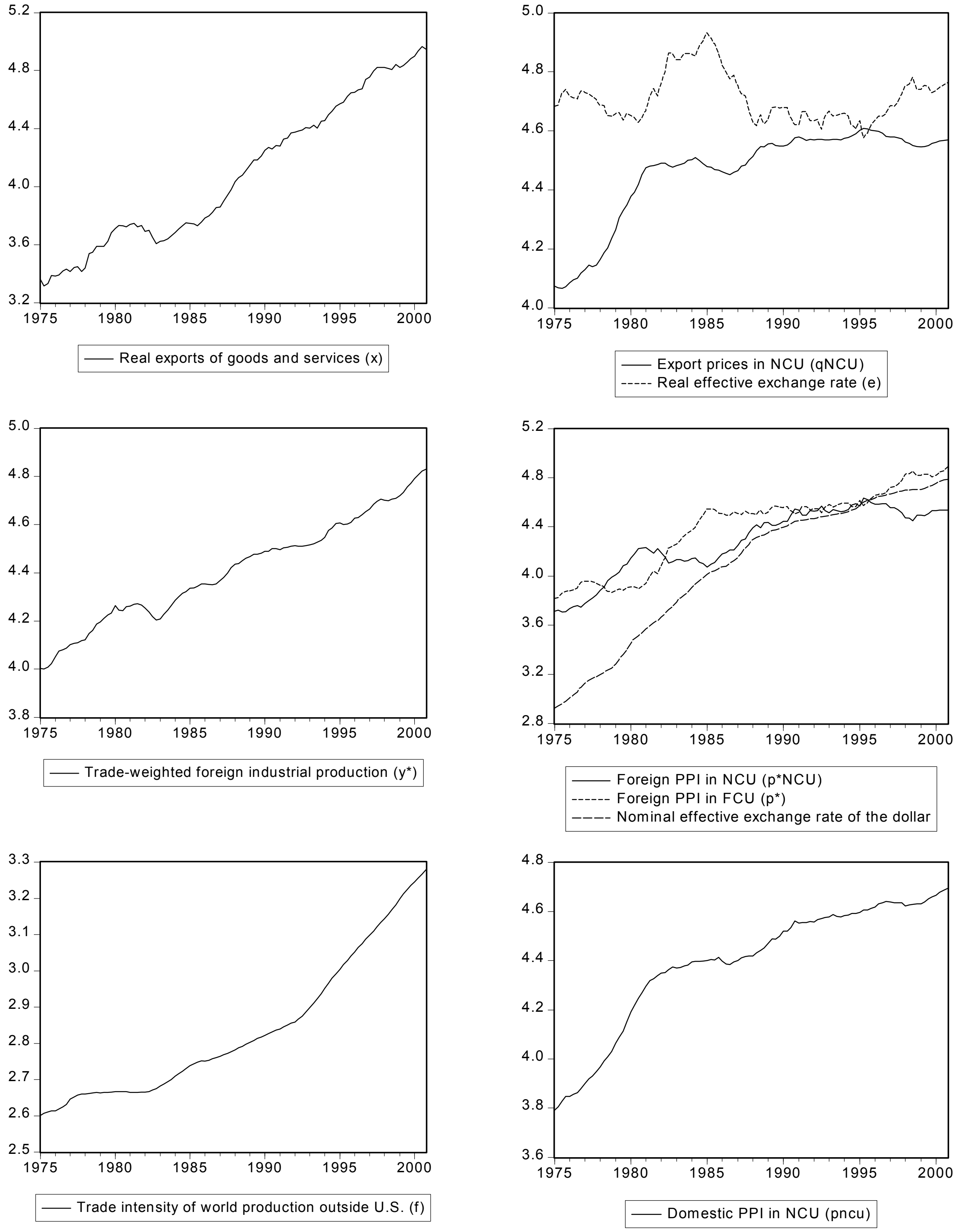
[14] For instance, German exports compete with Japanese goods not only in Japan but also in other foreign markets. In the weight of Japan, this "third-market effect" is taken into account.

Figure 1: Levels of Variables for the Empirical Export Model: United States[a]

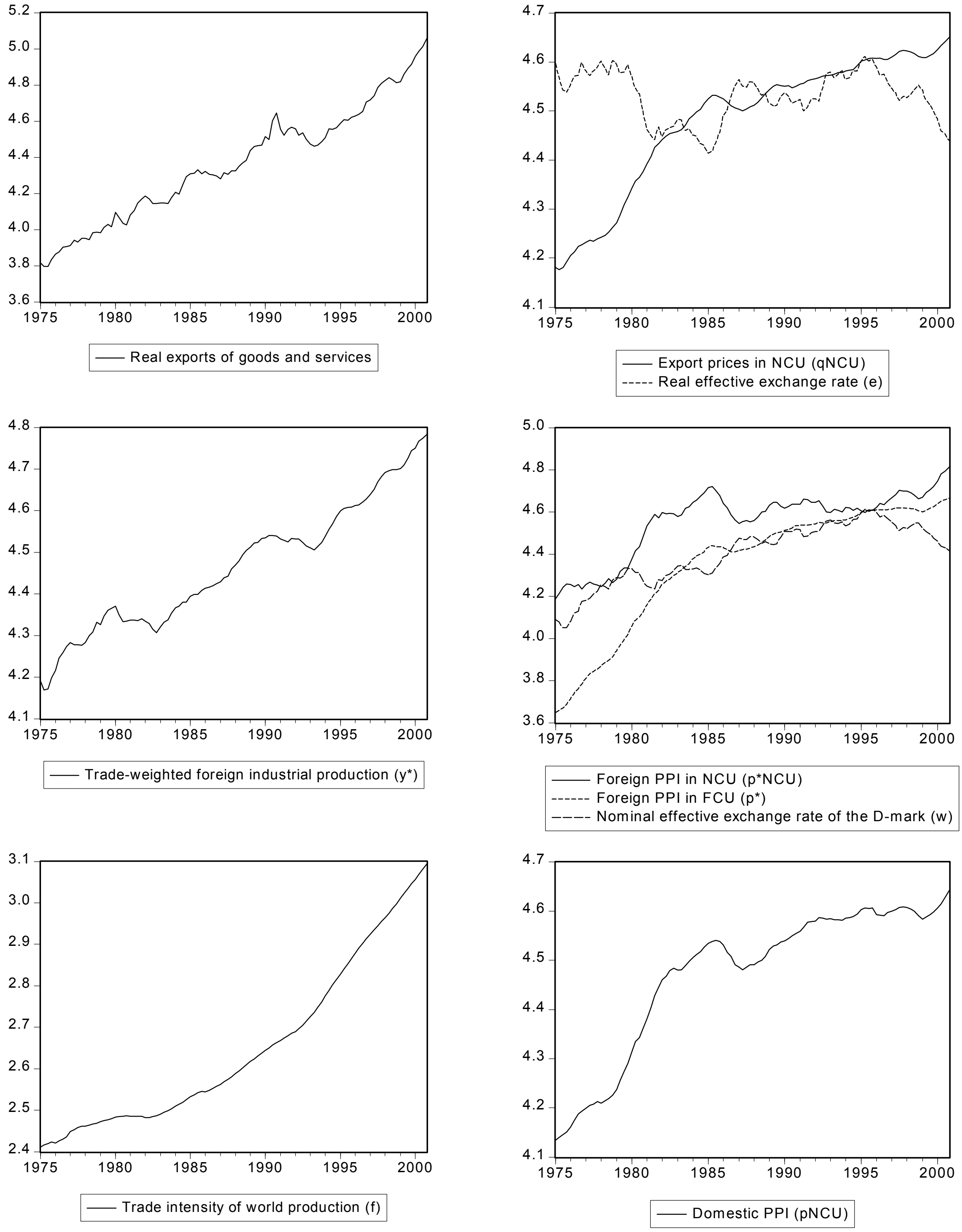


[a]Natural logarithms of the time series used explicitly or implicitly in the empirical model of aggregate export supply and demand. All original time series (except *f*) are indices (1996=100). NCU denotes "national currency units" (i.e. U.S. dollars), FCU "foreign currency units", PPI denotes producer price index.

foreign currency per national currency unit) is calculated as a geometric average from the bilateral nominal exchange rate indices (equalling 100 in the base year) using the same weights as for p^*. As to p^*_{NCU}, it is obtained by a two-step procedure: first the PPI for each foreign country i is converted into national currency (e.g. the Deutsche mark) according to $(p_i^* - w_i)$, then all these converted PPIs are aggregated as a geometric weighted average using the weights just mentioned. Accordingly, e, the index of the real effective exchange rate of the national currency (again as a value concept), is a weighted geometric average of the $(w_i - p_i^* + q_{NCU,i})$ of each partner country.

These computations make sure that whatever set of domestic and foreign price variables is used (($p_{NCU}, q_{NCU}, p^*_{NCU}$) or ($p, q, p^*$)), all prices in the empirical system are expressed in units of the same currency. In principle, one may choose either national currency units or the representative basket of foreign currency units. From the viewpoint of export demand, the latter choice would be intuitively more interesting as a foreign buyer gauges the attractiveness of, say, a German export good by looking not on the Euro price but on the price in units of his local currency. From the viewpoint of a German supplier, in turn, an analysis in national currency units (i.e. all prices in Deutsche mark or Euro) would be more appealing because the bulk of exports are invoiced in Euro, labor and capital costs are paid in Euro and because domestic wholesale prices, which determine the relative profitability of the exporting activity, are in Euro, as well. In both sets of prices, nominal exchange rates enter the analysis only indirectly. *Where* this happens to be the case constitutes the difference between the two specifications. When all price indices are expressed in national currency units (subscript NCU), w enters the analysis via the foreign price variable as the price level in each trading partner country, p_i^*, has to be "translated" into national currency units. When the price variables are expressed in foreign currency units (price variables without subscripts in Table 1), both domestic producer prices

Figure 2: Levels of Variables for the Empirical Export Model: Germany[a]



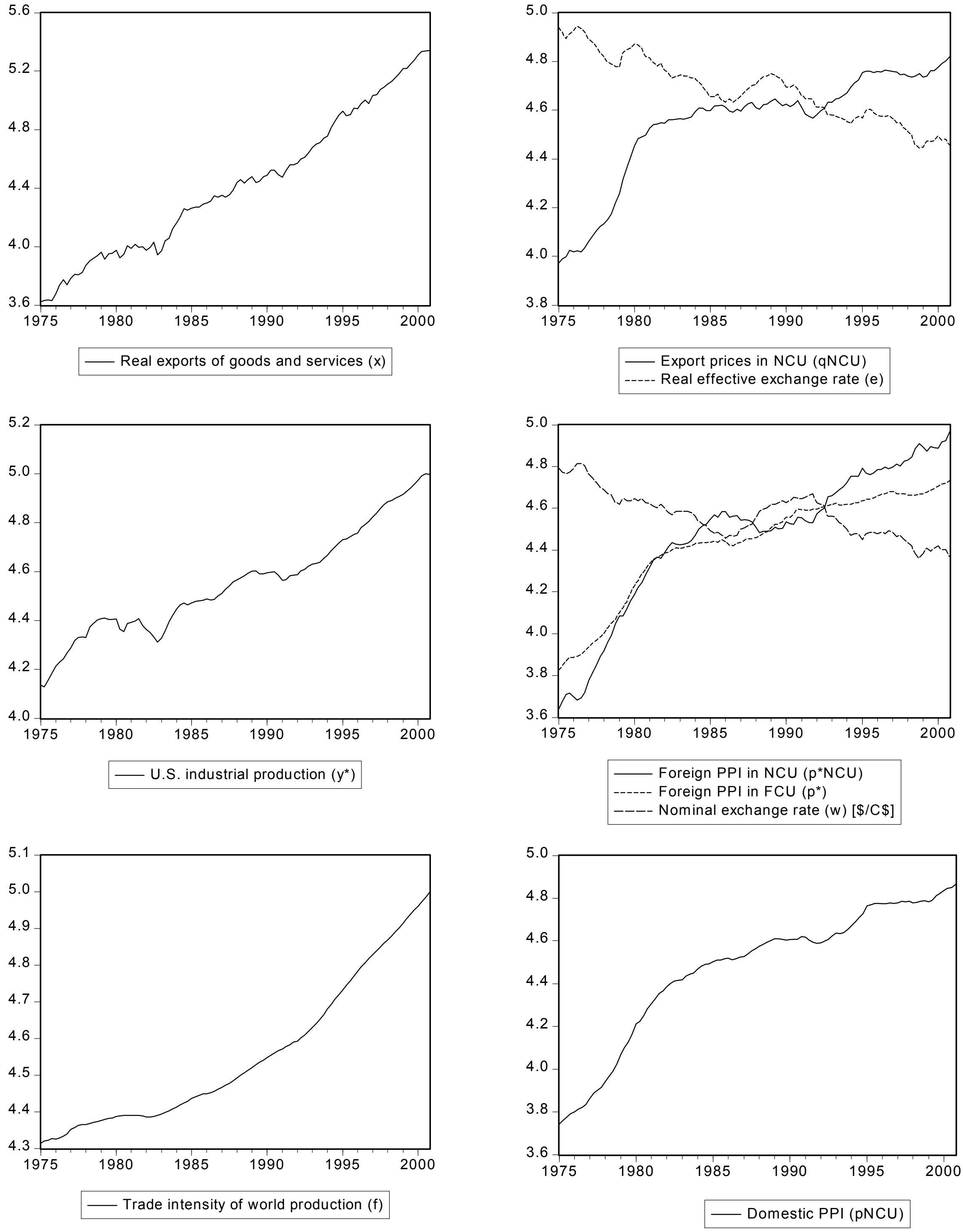
[a]Natural logarithms of the time series used explicitly or implicitly in the empirical model of aggregate export supply and demand. All original time series (except *f*) are indices (1995=100). NCU denotes "national currency units" (i.e. Deutsche marks or Euros), FCU "foreign currency units", PPI denotes producer price index.

and the deflator of exports are to be converted into units of the representative foreign currency. Hence there would be two combined variables. In turn, by choosing prices in national currency units, the number of combined variables is limited to one. In addition to this technical argument, the set of prices in NCUs turn out to yield somewhat more stable and reliable results in the econometric estimations below and are thus preferred for all three countries under investigation.

The operation $(p_i^* - w_i)$ which obtains the foreign price level in national currency units imposes a preliminary restriction on the data the validity of which is not checked any more. For instance, in the context of export demand, the use of p_{NCU}^* implies the assumption that a 1 percent devaluation of the domestic currency has the same (stimulating) effect on the demand for exports as a 1 percent increase in the foreign price level. The main motivation for unifying the currency in the estimation system is to decrease the number of variables (from 7 to 6 in my case) and thus to reduce complexity of the system.

Canada represents a special case within the group of three countries under investigation. Canada's exports are heavily concentrated on the U.S. market which absorbed about 87 percent of them in 1999 (OECD 2002b). This is why quite a lot of regressors in the Canadian export system only refer to the United States and are therefore less aggregated in nature than the regressors constructed for the other two countries. Specifically, the variables $x, q_{NCU}, \tilde{y}, f$ are constructed in the same way as for Germany, but all the other variables are different in the Canadian case: y^* is the index of U.S. industrial production (1992=100), $w\,(e)$ is the nominal (real) exchange rate of the Canadian dollar with respect to the U.S. dollar, p and q are expressions in U.S. dollar terms of the Canadian PPI and of the deflator of Canada's SNA exports, respectively. Finally, p^* (p_{NCU}^*) is the U.S. producer price index expressed in terms of U.S. dollars (Canadian dollars). Taking more trading partners of Canada into account would probably not change the Canadian results in a significant manner. Yet

Figure 3: Levels of Variables for the Empirical Export Model: Canada[a]



[a]Natural logarithms of the time series used explicitly or implicitly in the empirical model of aggregate export supply and demand. All original time series are indices (1992=100). NCU denotes "national currency units" (i.e. Canadian dollars), FCU "foreign currency units" (i.e. U.S. dollars), PPI denotes producer price index.

pragmatism is not the only reason to restrict the regressors to the dominant trading partner. The comparison between Canada and the other two countries could also be useful if it turned out that time series properties or cointegration results were different due to aggregation over a number of different trading partners.

4. Unit Root and Stationarity Tests for Detecting the Order of Integration

The conventional definition of cointegration is given by Engle and Granger (1987: 253): "The components of the vector x_t are said to be co-integrated of order d, b, denoted x_t ~ CI(d,b), b>0 if (i) all components of x_t are I(d); (ii) there exists a vector α (≠0) so that $z_t = \alpha \cdot x_t \sim I(d-b), b > 0$. The vector α is called the co-integrating vector." A more flexible concept is suggested by Campbell and Perron (1991: 25) who do not exclude stationary variables from the vector x_t. However, such an allowance only makes sense if x_t contains at least two integrated variables which can cointegrate (i.e. form a linear combination at a lower degree of integration. When applying economic equilibrium models empirically, stationary linear combinations are of special interest because they tell us to which level an otherwise random-walk like variable should converge given certain levels of the other variables in the cointegration space. As many economic time series are often found to be integrated of order one (Nelson and Plosser 1982),[15] others to be stationary (e.g. interest rate

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[15] Perron challenges Nelson and Plosser's findings by arguing that many macroeconomic time series would be stationary were it not for single breakpoints (e.g. the oil price shock in 1973) easily seen by visual inspection. He allows for a permanent change in the intercept (Perron 1990) or the slope of a deterministic trend (Perron 1989) at a known date, re-specifies the Augmented Dickey-Fuller test equation accordingly and corrects the Augmented Dickey-Fuller distribution for the additional parameters to be estimated. As a result, the hypothesis of non-stationarity is rejected in many cases. In the wake of Perron's papers the assumption that the date of the breakpoint is known has been questioned. Zivot

spreads) and still others even tend to be integrated of order two (e.g. price indices in some countries, cf. Juselius 1993: 5), checking for the order of integration of a time series is a natural start to cointegration analysis because it allows to know whether the levels (in case of I(1) variables) or the first differences of a variable (when it is I(2)) should be used to compose the long-run relationships. The objective of this section consists in demonstrating by appropriate tests that the entity or at least a sufficient number of the time series enumerated in Table 1 are integrated of order 1 and can therefore be used in a vector error correction framework involving their levels and first differences. The order of integration is tested for by two different methods: the augmented Dickey-Fuller unit root test and the KPSS test. The former asserts the variable is I(1) in the null hypothesis while the latter formulates the stationarity assumption as the null.

4.1. The Augmented Dickey-Fuller Test

4.1.1. Idea and specification issues

In the simplest version of the Dickey-Fuller test (DF test) a unit root in time series y_t is searched for by estimating

[3a] $y_t = \mu_a + \rho_a y_{t-1} + u_t \quad \Leftrightarrow \quad$ [3b] $\Delta y_t = \mu_a + \gamma_a y_{t-1} + u_t$,

where $\gamma_a = \rho_a - 1$ and $u_t \sim iid\left(0, \sigma^2\right)$

and by then testing the one-sided Hypothesis $H_0 : \gamma_a = 0$ against the alternative $H_1 : \gamma_a < 1$. Thus the null is non-stationarity of y_t. Under the null, the *t*-values of γ_a do not follow a standard *t*-distribution but a Dickey-Fuller distribution (Dickey and Fuller 1981: 1062). Actually, applying standard critical *t*-values would result in over-rejection of the null. The reason for this deviation from the

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and Andrews (1992) present a procedure to detect the most probable date of one break-point in a given sample.

standard t-distribution is that the tests based on the DF-distribution are similar tests, e.g. the test equation for Δy_t contains more deterministic elements than the data-generating process (d.g.p.) for Δy_t in order to ensure that the null and the alternative hypotheses be nested by the test (Harris 1995: 30–31). Thus equation [3b] above can serve as a valid test equation only if the initial value y_0 of the d.g.p. and thus μ_a equal zero.[16] Therefore a deterministic trend has to be incorporated into [3b] if the economic time series at hand contain a deterministic trend in levels. The question whether this is the case or not can be addressed by sequential testing, with the help of economic theory or by visual inspection. The former two of these ways are briefly sketched in the following.

Perron (1988) suggests a test sequence which aims at determining both the order of integration and the most probable deterministic structure of a time series. Starting with a DF test equation including trend and intercept,

[4] $$\Delta y_t = \mu_b + \delta_b t + \gamma_b y_{t-1} + u_t \ ,$$

first the null of a unit root (against the alternative of trend stationarity) and then the joint hypothesis of a unit root and no trend are tested. If the latter is rejected whereas the former is not, the deterministic trend is accepted as an element in the d.g.p. of y_t. The d.g.p. and the DF test equation thus have the same design ("exact test" as it is called by Harris (1995: 31)). As a consequence, the critical values of the standard normal distribution can now be taken to evaluate $\hat{\gamma}_b$ because the deterministic trend asymptotically dominates the stochastic trend (West 1988). The null of non-stationarity might now be rejected more easily. However, if the joint F-test on the deterministic and the unit root is *not* rejected, it is concluded that there is no deterministic trend in y_t, so δ_b can be dropped

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[16] Imagine the d.g.p. for Δy_t actually has an intercept and thus the level of the series a deterministic trend. Then the estimated coefficient for $\hat{\rho}_a$ could be one even if y_t *were* in fact stationary because setting $\hat{\rho}_a = 1$ in [3a] would be the only way to fit the deterministic trend.

from the model. Then an analogous testing sequence is applied to the model described in [3b] (unit root against the alternative of mean stationarity) with the joint hypothesis being formulated with respect to the intercept instead of the trend. It might turn out in the *F*-test that even the intercept is zero, then Fuller's (1976) critical values $\hat{\tau}$ for a test equation without deterministic elements (unit root against the alternative of stationarity around zero) should be used. The motivation for dropping unnecessary nuisance parameters is that the critical values of the DF distribution become smaller with each additional deterministic element so that overloading the model with unjustifiable deterministic parameters leads to under-rejection of the null. The procedure stops whenever the null of a unit root is rejected.[17]

There are several objections against the use of the sequential testing procedure just described. First, a general size problem can result from sequential testing by the potential accumulation of wrong test decisions, e.g. when the *F*-tests rejects although the nulls are true. Second, the use of standard normal distribution is misplaced as soon as the deterministics of the d.g.p. do not match the one of the test equation. To rely on the result of the *F*-test involves a risk: If the true d.g.p. does not contain a deterministic trend, its presence in the test equation introduces a downside bias in the estimate of γ_b thereby making a rejection of the null of non-stationarity more probable (Harris (1995: 50)). Thus it is not clear whether the rejection of the null in the *F*-test really stems from the presence of a deterministic trend or whether the process is in fact stationary. Clearly, the *F*-test cannot solve the DF test's problem of low power against a trend-stationary alternative. Third, if the true value of ρ_b is close to but smaller than one, convergence to the asymptotic test distribution is slow. Thus in small samples the DF distribution may be a better approximation (Banerjee et al.

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[17] A graphical illustration of Perron's sequence is given by Enders (1995: 257).

1993). To sum up, sequential testing is not appropriate to solve the huge uncertainty surrounding the true nature of the d.g.p. in economic time series.[18]

This is why a pragmatic mixture of economic judgement and visual inspection is adopted in the following to find out the appropriate deterministic structure of the DF test equations, whereas Perron's procedure is only referred to in a few cases as a control device. For all three countries in the sample, the logs of variables of real economic activity $(x, y^*, \tilde{y})$ are supposed to contain a deterministic trend due to wide acceptance of economic growth as a normal phenomenon (Kaldor 1961; Solow 1999). A linear trend is allowed for in f, the world trade intensity, because of the historical observation of world exports growing much faster than world GDP since the end of World War II. This trend may not be stable in the very long run as historically phases of accelerating and decelerating world trade succeed to one another. During the sample period at hand (1974–2000), conditions have been favorable to a persistent increase in f, even at a faster pace after 1985 than before. Visual inspection strongly supports this view. As to the level of export and producer prices (p_{NCU}, p^*, q_{NCU}), their DF equations contain a deterministic trend because since the deflationary experience during the Great Recession, virtually all central banks all over the world have tried to avoid deflation by allowing for some positive annual inflation deviations from which are generally not tolerated in the long run.[19] Finally, exchange rates are the only variables for which a d.g.p. without intercept is conceivable but a distinction should be made between the real exchange rate (e) and the nominal exchange rate (w). Whereas a persistent

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[18] A recent example of the debate on the appropriate procedure to decide on the degree of integration is the controversy between Assenmacher (1998) and Meier (2001) on the behavior of German GDP. On the basis of annual data from 1953 to 1995 both find a deterministic trend, but while the former uses DF critical values and finds German GDP to be I(1), the latter rejects the null using standard normal distribution after rejecting the insignificance of the trend parameter, as advised in Perron (1988).

[19] From a theoretical standpoint an explicit or implicit inflation target that is enforced vigorously by the so-called "Taylor principle" represents a stability condition in macroeconomic models of the new IS-LM type (McCallum 2001: 156–157).

inflation differential between the domestic economy and the economies of major trading partners (e.g. due to a catching-up process in economic development described by the Balassa-Samuelson effect or to different monetary policy strategies) might lead to a deterministic trend in the nominal exchange rate, the validity of the purchasing power parity (PPP) should have e being stationary around mean zero (in the strong version of PPP) or around a non zero mean (in the weaker version).[20] The PPP doctrine presents a case for specifying the real effective exchange rate without trend in the DF equation.

A more difficult decision regards the combined price variables p, p^*_{NCU} and q which are linear combinations between the log of the original price index and the log of the nominal exchange rate (w). Results are reported for $p = p_{NCU} + w; p^*_{NCU} = p^* - w; q = q_{NCU} + w$, where NCU stands for national currency unit of the country for which exports are analyzed. As potential trends could cancel each other out, the initial presumption of price levels containing a positive deterministic trend is supplemented by visual inspection and a look at the F-test results from Perron's (1988) test sequence (Table 2). A lower degree of integration in $p; p^*_{NCU}; q$ than in the original series could point to a stationary linear combination between the respective series composing each of the variables.

Besides the deterministic components, the choice of the number of lagged first differences is of crucial importance. This is because the DF distributions have been drawn from an ARIMA (1,1,0) process and are highly sensitive to a change in the MA properties (Schwert 1989). However, the DF critical values still hold for higher-order AR-processes as is illustrated by Harris (1992) who generates data by an integrated AR (3) process and then studies in how many cases the

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20 The strong version of PPP postulates $E=WP^*/P=1$ and therefore $e=logE=0$. In the weak version, the ratio of domestic and foreign prices converges to some constant different from one (Obstfeld and Rogoff 1996).

Table 2: Deterministic Trends in U.S., Canadian and German Exchange Rates and Price Indices Denominated in Other than the Original Currencies

Variable	Definition	Presence of a trend according to			
		economic judgement	visual inspection	Perron's F-test[a]	t-value of $\hat{\delta}$
United States					
w	w	yes or no	no[b]	1.21	1.26
p	$p_{NCU}+w$	yes	yes	3.02	0.81
q	$q_{NCU}+w$	yes	yes	3.91	0.15
p^*_{NCU}	p^*-w	yes	yes[c]	2.34	0.53
e	$w-p^*+q_{NCU}$	no	no	1.22	–1.16
Canada					
w	w	yes or no	yes[d]	2.75	–2.03
p	$p_{NCU}+w$	yes	yes[e]	3.67	1.28
q	$q_{NCU}+w$	yes	no[f]	3.02	0.29
p^*_{NCU}	p^*-w	yes	yes	5.36[g]	2.17
e	$w-p^*+q_{NCU}$	no	yes	3.68	–1.78
Germany					
w	w	yes or no	yes[h]	2.23	–0.03[h]
p	$p_{NCU}+w$	yes	yes[h]	12.69***	–1.94[h]
q	$q_{NCU}+w$	yes	yes[h]	11.24***	–2.28[h]
p^*_{NCU}	p^*-w	yes	yes[i]	2.35	1.71
e	$w-p^*+q_{NCU}$	no	no	2.06	–0.51

* (**, ***) means rejection at the 10% (5%, 1%) significance level. – [a]Rejection means the trend should be in the equation. – [b]Periods of strong appreciation (1980–1985 and 1995–2000) are preceded and followed by periods of rough constancy in the index. – [c]Clear upward trend in the original series in foreign currency units is interrupted during periods of strong dollar appreciation (1980–1985 and 1995–2000). – [d]Downward trend in the nominal exchange rate against the U.S. dollar only interrupted between 1987 and 1991. – [e]Clear upward trend only until 1989. – [f]Strong rise before 1981. – [g]10% critical value: 5.47. – [h]Strong rise from 1973–1996, shrinking levels in foreign currency units thereafter due to devaluation of the D-Mark. – [i]Only slight upward trend from 1986–2000 but strong rise from 1973–1985.

Source: Own calculations.

null of a unit root (which is true) is rejected based on the Dickey and Fuller (1981) 5 percent critical values. As each MA(1) process can be approximated by an AR(∞)-process, the problem of autocorrelation in the residuals of the simple DF test equation (indicating the presence of MA terms) can be addressed satisfactorily in practice by approximating the actual ARMA(*p,q*) d.g.p. by an AR(*k*) process with *k* being large enough to ensure that residuals be "white noise" (Said and Dickey 1984). Among the solutions of determining *k* suggested in the literature are the minimization of the Akaike information criterion and some mechanical rule of increasing *k* as the sample size grows (Schwert 1989). While the former tends to yield under-parameterized test equations and to fail in eliminating significant MA parameters leading to over-rejection of the null and, more generally, to the inapplicability of the DF distribution, the latter obtains size close to its nominal value at the cost of efficiency losses due to the inclusion of insignificant lags (Harris 1992).[21] However, as in small samples the unavoidable trade-off between power and size is especially strong[22], one may not want to overload the test equation with unnecessary parameters because this implies an unnecessary loss in power. I therefore specify the Augmented Dickey-Fuller (ADF) test equation such as to minimize the number of lagged first differences subject to freedom of residual autocorrelation:[23]

$$[5] \qquad \Delta y_t = \mu + \delta t + \gamma y_{t-1} + \sum_{j=1}^{l-1} \rho_j \Delta y_{t-j} + u_t \,,$$

where *l* is the lag length of the model.

[21] For instance, for *T=100,* a d.g.p. containing a non zero initial value, a drift, and $\gamma \in [0.80;0.98]$, ADF test equations with trend, intercept and lagged first differences up to order $l_{12} = \text{int}\{12(T/100)^{1/4}\}$ achieve power between 0.344 and 0.501 (Harris 1992: 385). Power generally increases in *T*, but interestingly, power does not decrease monotonously as γ approaches unity but reaches highest values for $\gamma = 0.94;0.96;0.98$ before collapsing for $\gamma = 1$.

[22] This trade-off is described in Blough (1992).

[23] This pragmatic solution to the trade-off is taken from Hassler (2001: 8). As quarterly data are used, tests for residual autocorrelation up to the fourth order are undertaken.

4.1.2. Results

The results of the ADF test for a unit root in the time series for Germany, the United States, and Canada are documented in Tables 3, 4, and 5, respectively. In the left half of each table, tests for the null of a unit root in the (logs of) levels of the time series are undertaken, so Δy_t is the dependent variable the test equation. As the null is not rejected for most of the variables at hand (33 out of 35),[24] the ADF test is also applied to the first difference of the variables to see whether two unit roots are present in their level. Then $\Delta^2 y_t$ is the dependent variable. The results of these tests can be seen in the right half of each table. The null of non-stationary first differences is rejected in 30 out of 35 cases leaving 5 variables found to be integrated of order 2. These are the index of Canadian producer prices in Canadian dollars, the export-weighted index of foreign producer prices in units of foreign currency for Germany, and the index of world trade intensity for all countries. However, if a deterministic trend is allowed for in the *growth* in world trade intensity, f turns out to be a limit case between I(1) and I(2) as the null of a unit root in Δf is then rejected for the United States and is at least close to rejection at the 10 percent significance level for Germany and Canada. A deterministic trend in Δf corresponds to a quadratic trend in the level of world trade intensity. Although this is not quite intuitive on economic grounds (at least in the long run), it fits well to the evolution of this time series during the last quarter of the twentieth century. The variables found to be trend-stationary are the producer price index in domestic currency units as well as the trade-weighted index of foreign industrial production for the United States (Table 3).

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24 In case of Canada (Table 4), p^* is identical with the U.S. p_{NCU} (Table 3). This reduces the number of stationary variables from three to two.

Table 3: Results of the Augmented Dickey Fuller (ADF) Unit-Root Tests – United States

Variable	Test for I (0)			Test for I (1)			Result
	Specification[a]	ADF test statistics[b]	LM (4)[c]	Specification[a]	ADF test statistics[b]	LM (4)[c]	
x	T, 0	–1.71	0.12	C, 0	–8.85***	0.39	I (1)
y^*	T, 2	–3.65**	0.23	C, 4	–4.53***	0.24	I (0)
$\tilde{y}$	T, 5	–2.84	0.18	C, 3	–3.00**	0.33	I (1)
f	T, 9	–0.81	0.29	C, 8	–1.29	0.10	I (2)[d]
w	T, 1	–1.49	0.14	C, 0	–8.00***	0.30	I (1)
	C, 1	–0.91	0.32	N, 2	–3.61***	0.38	I (1)
p_{NCU}	T, 6	–3.58**	0.13	C, 2	–2.37	0.14	I (0)[e]
q_{NCU}	T, 1	–1.36	0.12	C, 0	–5.27***	0.13	I (1)
p^*	T, 1	–0.45	0.16	C, 0	–3.39**	0.35	I (1)
p	T, 1	–1.47	0.34	C, 0	–7.77***	0.22	I (1)
q	T, 0	–1.07	0.10	C, 0	–8.22***	0.20	I (1)
p^*_{NCU}	T, 1	–1.31	0.13	C, 0	–7.76***	0.21	I (1)
e	C, 0	–1.04	0.35	N, 0	–9.33***	0.44	I (1)

* (**, ***) means rejection at the 10% (5%, 1%) significance level. – [a]T: model with drift and trend; C: model with drift; N: model without drift and trend. The figure indicates the number of lagged first differences in the test equation. – [b]ADF *t*-test. Critical values are τ_τ (model T), τ_μ (model C), and τ (model N), respectively. They are taken from Hamilton (1994:…. who derives them from the response-surface simulation proposed by MacKinnon (1991). – [c]Probability of the Lagrange Multiplier test for residual autocorrelation of first to fourth order. – [d]When allowing for a quadratic trend in f (accounting for the acceleration in world trade intensity since the mid-eighties), the ADF statistic is –5.15*** (model T,9; LM(4) [0.28]). Δf is then found to be stationary. – [e]If the "true" d.g.p. for Δp_{NCU} did not contain a constant (as suggested by the Perron (1988) testing sequence), the null of non-stationarity of p_{NCU} could be accepted at the 5% level, while the null of non-stationarity could be rejected for Δp_{NCU} at the 10% (not at the 5%) level.

Source: Own calculations.

The German reunification is likely to influence the results for German exports and potential output quite substantially. As West German exports peak in the second half of 1990 following German monetary union and a strong increase of intra-German deliveries of goods and services, which are not counted any more after the switch to pan-German statistic in 1991, x is likely to be biased towards

Table 4: Results of the Augmented Dickey Fuller (ADF) Unit-Root Tests – Canada

Variable	Test for I (0)			Test for I (1)			Result
	Specifica-tion[a]	ADF test statistics[b]	LM (4)[c]	Specifica-tion[a]	ADF test statistics[b]	LM (4)[c]	
x	T, 0	–2.16	0.61	C, 0	–10.25***	0.51	I (1)
y^*	T, 1	–2.39	0.15	C, 0	–6.12***	0.26	I (1)
$\tilde{y}$	T, 1	–1.98	0.71	C, 0	–2.83*	0.70	I (1)
f	T, 13	–0.31	0.21	C, 12	–0.42	0.14	I (2)[d]
w	T, 3	–2.33	0.31	C, 2	–3.60***	0.36	I (1)
	C, 4	–1.16	0.30	N, 2	–3.38***	0.46	I (1)
p_{NCU}	T, 3	–2.33	0.30	C, 2	–2.37	0.78	I (2)[e]
q_{NCU}	T, 1	–2.14	0.33	C, 0	–5.46***	0.50	I (1)
p^*	T, 6	–3.58**	0.13	C, 2	–2.37	0.14	I (0)[f]
p	T, 3	–2.22	0.15	C, 2	–2.87*	0.13	I (1)
q	T, 1	–1.84	0.13	C, 0	–6.12***	0.12	I (1)
p^*_{NCU}	T, 4	–2.01	0.25	C, 2	–3.49***	0.21	I (1)
e	T, 3[g]	–2.66	0.20	C, 2	–3.93***	0.11	I (1)
	C, 3	–2.02	0.19	N, 2	–3.88***	0.14	I (1)

* (**, ***) means rejection at the 10% (5%, 1%) significance level. – [a]T: model with drift and trend; C: model with drift; N: model without drift and trend. The figure indicates the number of lagged first differences in the test equation. – [b]ADF *t*-test. Critical values are Dickey and Fuller's τ_τ (model T), τ_μ (model C), and τ (model N), respectively. They are taken from Hamilton (1994: …) who derives them from the response-surface simulation proposed by MacKinnon (1991). – [c]Probability of the Lagrange Multiplier test for residual autocorrelation of first to fourth order. – [d]When allowing for a quadratic trend in f (accounting for the acceleration in world trade intensity since the mid-eighties), the ADF statistic is –2.87; although the decision is then tighter, Δf is still found to be non-stationary. – [e]If the "true" d.g.p. for Δp_{NCU} did not contain a constant (as suggested by the Perron (1988) testing sequence), the null of non-stationarity could be rejected for Δp_{NCU} at the 10% level. – [f]If the "true" d.g.p. for Δp^* did not contain a constant (as suggested by the Perron (1988) testing sequence), the null of non-stationarity of p^* could be accepted, while the null of non-stationarity could be rejected for Δp^* at the 10% level. – [g]As visual inspection suggests, there could be a deterministic trend in the real exchange rate of the Canadian dollar.

Source: Own calculations.

stationarity. Indeed, the ADF statistic is the lowest for all German variables but the null of a unit root is not rejected (Table 5).[25] Unlike exports, German potential output, $\tilde{y}$, undergoes a permanent positive shock due to the enlargement of territory. As a consequence, the test result could be biased towards non-stationarity if it is not corrected for the shift in the intercept occurring in 1991.[26] I therefore subtract an estimate of East German output to eliminate this shift as described in Table 5. However, the result $\tilde{y} \sim I(1)$ still holds.

The example of the index of foreign producer prices (p^*) in the German results underlines how sensitive the ADF results can be with respect to the deterministic components or to the lag length of the model (or to both). When a trend is in the test equation (corresponding to the result reported in the table), the null hypothesis is that p^* is I(1) with drift, the relevant alternative hypothesis is trend-stationarity of p^*. In turn, when there is no trend in the test equation, H_0 is $p^* \sim I(1)$, whereas the relevant H_1 is stationarity around a non-zero mean. It is curious that the former test favors I(1) with drift while the latter decides I(0) around a mean. Non-rejection of the nulls in the tests for I(0) and I(1) with a deterministic trend in p^* might be due to a loss in power as the number of lags is quite high.[27] The example illustrates that wrong assumptions on the deterministic structure of the d.g.p. can lead to wrong test decisions. Sequential testing of joint hypotheses of unit roots and deterministic components does not seem to be quite helpful as accumulation of wrong test decisions are likely to exacerbate power and size problems. The example of the German p^* is

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[25] See, however, the impossibility of rejecting the hypothesis of stationarity in the KPSS test (Table 8 below).

[26] There might also be a break in the deterministic trend as the German economy grew slowlier after 1991 than before. But this has not been taken into account because it is not clear whether the slowdown in growth rates is mainly due to reunification.

[27] A similar point could be made for *f*. However, there is no alternative to taking so many lags into account because the construction of the variable already involves autocorrelation up to order 12.

neither the only one (see, for instance, the ADF test for the U.S. p_{NCU} in Table 3), nor are the undesirable consequences of wrong assumptions on deterministic components limited to the ADF test.

Table 5: Results of the Augmented Dickey Fuller (ADF) Unit-Root Tests – Germany

Variable	Test for I (0)			Test for I (1)			Result
	Specification[a]	ADF test statistics[b]	LM (4)[c]	Specification[a]	ADF test statistics[b]	LM (4)[c]	
x	T, 0	–2.27	0.26	C, 0	–9.01***	0.46	I (1)
y^*	T, 4	–1.58	0.13	C, 3	–6.44***	0.21	I (1)
$\tilde{y}$	T, 0	–1.64[d]	0.58	C, 0	–9.25[e]***	0.94	I (1)
f	T, 13	–0.53	0.16	C, 9	–0.67	0.23	I (2)[f]
w	T, 1	–0.77	0.30	C, 0	–7.03***	0.56	I (1)
	C, 1	–2.12	0.30	N, 0	–6.88***	0.55	I (1)
p_{NCU}	T, 1	–1.84	0.52	C, 0	–3.79***	0.55	I (1)
q_{NCU}	T, 1	–2.15	0.43	C, 0	–4.27***	0.21	I (1)
p^*	T, 9	–2.17	0.12	C, 5	–2.10	0.15	I (2)[g]
p	T, 0	0.19	0.11	C, 0	–6.65***	0.89	I (1)
q	T, 6	0.49	0.89	C, 0	–6.87***	0.62	I (1)
p^*_{NCU}	T, 3	–2.17	0.17	C, 0	–6.14***	0.14	I (1)
e	C, 1	–1.97	0.13	N, 0	–7.19***	0.26	I (1)

* (**, ***) means rejection at the 10% (5%, 1%) significance level. – [a]T: model with drift and trend; C: model with drift; N: model without drift and trend. The figure indicates the number of lagged first differences in the test equation. – [b]ADF t-test. Critical values are τ_τ (model T), τ_μ (model C), and τ (model N), respectively. They are taken from Hamilton (1994: …) who derives them from the response-surface simulation proposed by MacKinnon (1991). – [c]Probability of the Lagrange Multiplier test for residual autocorrelation of first to fourth order. – [d]When eliminating the break in 1991 by subtracting the log of East Germany's GDP and applying Germany's potential growth rates to the West German 1990 level during the nineties, the results are: T, 1; –2.18; 0.21. Thus The finding of non-stationarity does not stem from the permanent output shift due to the enlargement of the German territory. – [e]When $\tilde{y}$ is corrected according to footnote d, the null of $\Delta\tilde{y} \sim I(1)$ can still be rejected at the 5% level (C, 1; –3.04**; 0.63). – [f]When allowing for a quadratic trend in f (accounting for the acceleration in world trade intensity since the mid-eighties), the ADF statistic is –2.77; although the decision is then tighter, Δf is still found to be non-stationary. – [g]Result surrounded by some uncertainty. If the d.g.p. for p^* contained no deterministic trend, the unit-root test for the level p^* would have led to rejection and thus mean-stationarity would have been found.

Source: Own calculations.

4.2. The KPSS test for stationarity

4.2.1. Idea and specification issues

The ADF test has been found to suffer from low power both against specific stationary alternatives (e.g. $\rho = 0.85$, cf. DeJong et al. 1989) and against fractionally integrated alternatives (Diebold and Rudebusch 1991). As the null is non-stationarity, low power implies non-rejection of the non-stationarity assumption when it is false in many cases. Therefore Kwiatkowski, Phillips, Schmidt and Shin (1992) propose a test design, the KPSS test, which has stationarity as the null and a unit root as the alternative hypothesis.[28] Testing both the null of a unit root and the null of stationarity allows for an empirically more satisfying sorting of time series into stationary ones, ones containing a unit root and ones not containing sufficient information to distinguish between these two cases (Kwiatkowski et al. 1992: 176). To achieve a test nesting a stationary null and an integrated alternative, Kwiatkowski et al. decompose the series under investigation, y_t, into a deterministic trend, ξt, a random walk, r_t, and a stationary error, ε_t, according to:

[6] $y_t = \xi t + \mu_t + \varepsilon_t$, where $\mu_t = \mu_{t-1} + u_t, u_t \sim iid\left(0, \sigma_u^2\right)$.

Under the null we have $\sigma_u^2 = 0$, i.e. $\mu_t = \mu_{t-1} = ... = \mu_0 = \mu$ reduces to a constant. By construction, one concludes $y_t \sim I(1)$ if $\sigma_u^2 = 0$ is rejected by a one-sided Lagrange-Multiplier (LM) test. [6] allows to test the null of trend stationarity against the alternative of a unit root with drift for $\xi \neq 0$ (case a) or to test the null of level stationarity against the alternative of a unit root and some

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28 In a related paper, Shin and Schmidt (1992) show that the KPSS statistic derived by Kwiatkowski et al. (1992) can also be used with a unit root being the null hypothesis and provide critical values. This unit root test turns out to be not very powerful, however. Specifically, it is less powerful than the ADF test of a unit root against the trend-stationary alternative relevant for most of the time series I use.

non-zero initial value for $\xi = 0$ (case b). Accordingly, the test statistic is constructed upon the de-trended series or upon $\bar{y}$, the deviations from mean:

[7a] $v_t = y_t - \mu - \xi t,$ [7b] $v_t = y_t - \bar{y}.$

In either case, as long as $\varepsilon_t \sim niid(0, \sigma_\varepsilon^2)$, the relevant test statistic is computed based on partial sums according to

[8] $LM = \sum_{t=1}^{T} S_t^2 / \hat{\sigma}_\varepsilon^2$, where $S_t = \sum_{i=1}^{t} v_i, t = 1,2,\ldots,T.$ [29]

However, the normality assumption for ε_t is not very realistic since many macroeconomic time series, stationary or not, are characterized by auto-correlated error terms. Therefore, Kwiatkowski et al. (1992: 164–167) derive the correct value and the asymptotic distribution of the test statistic under the much weaker assumption of ε_t just being stationary and allowing for

[9] $\varepsilon_t = \kappa_t + \theta \kappa_{t-1},\ |\theta| < 1,$

where $-1 < \theta < 0$ is the most relevant case in practice. As the off-diagonal elements of the variance-covariance matrix of ε_t then contain non-zero elements, auto-covariances $\sigma_{\varepsilon_t \varepsilon_{t+s}}$ matter and the denominator in [8], the variance σ_ε^2, has to replaced by the "long-run variance" ω_ε^2 defined as

[10] $$\omega_\varepsilon^2 = \sigma_{\varepsilon_t}^2 + 2 \sum_{s=1}^{\infty} \sigma_{\varepsilon_t \varepsilon_{t+s}}$$

and consistently estimated e.g. by

[11] $$\hat{\omega}_\varepsilon^2 = \frac{1}{T} \sum_{t=1}^{T} \hat{\varepsilon}_t^2 + 2 \sum_{s=1}^{k} \left[\left(1 - \frac{s}{k+1}\right) \frac{1}{T} \sum_{t=s+1}^{T} \hat{\varepsilon}_t \hat{\varepsilon}_{t-s} \right].$$

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[29] Under the null the residuals ε_t and v_t are identical.

The expression $[1-\frac{s}{k+1}]$ is the Bartlett window ensuring non-negativity of the estimated long-run variance. The truncation parameter has to become infinitely large with $T \to \infty$ but growing at rate $\sqrt{T}$ is usually satisfactory. Kwiatkowski et al. show that the thus modified test statistics, η_T (case a) and η_C (case b), follow the distributions

$$[12a]\ \ \eta_T \equiv \frac{\sum_{t=1}^{T} S_t^2}{T^2 \hat{\omega}_\varepsilon^2} \to KPSS_\tau(0) \ \text{ and } [12b]\ \ \eta_c \equiv \frac{\sum_{t=1}^{T} S_t^2}{T^2 \hat{\omega}_\varepsilon^2} \to KPSS_\mu(0),$$

with S_t^2 taken from [7a], [8] and from [7b], [8], respectively.

As the correct specification of *k*, the truncation parameter, is a difficult issue in practice, I allow for a high order of auto-covariances and report the results depending on four different truncation parameters ($k = 6; k = 12; k = 18; k = 24$). As to the interpretation of the test results, the decision is clear when all four truncation parameters lead to the same decision. However, when the critical value is crossed from above or from below with *k* increasing from 0 to 24, the null is accepted and rejected, respectively, if one relies exclusively on the result of the higher *k*.[30]

In order to achieve a sound decision both in cases of switching test results when *k* varies and in cases where the ADF and the KPSS tests yield different outcomes, one has to bear in mind the potential problems with size and power depending on the auto-correlation pattern of the d.g.p. Specifically, in presence of AR(1) error terms the KPSS test suffers from serious size problems that can be successfully addressed by increasing the truncation parameter. However, when increasing *k*, one faces a trade-off: higher *k* reduces the probability of over-rejection (size problem) but at the cost of non-rejection of the null when it

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[30] In some cases the KPSS statistic is not monotonous for $k \in [0;24]$, but this may stem from single significant autocorrelation coefficients of higher order which might be spurious.

is false in many cases (loss in power). As a general rule, power is the lower the smaller $\lambda = \sigma_u^2 / \sigma_\varepsilon^2$, e.g. due to a slowly decaying auto-correlation function (ACF) for the error ε_t.[31] In Kwiatkowski et al. (1992: 171–172), results from Monte-Carlo simulations are tabled for three different lag lengths. The highest losses in power occur when switching from an intermediate lag length (“*l4*”, corresponding to *k=7* for *T=100*) to a high lag length (“*l12*”, corresponding to *k=21* for *T=100*), whereas the most substantial gains in size occur when switching from *“l0”* (*k=0*) to *“l4”*. Thus if nothing is known about the residuals of the d.g.p., choosing *k* neither too low nor to high would be a risk-minimizing strategy.

However, specification and results of the ADF test reported above contain information which may serve as guidelines with respect to the choice of *k*. First, the ADF test result gives a prior on whether a series is difference stationary or at least how close it is to difference stationarity. If a series *y* is I(1), the auto-correlation function (ACF) of the de-trended *y* will not die out (Enders 1995: 181–182), and higher-order co-variance terms matter for a correct approximation of the long-run variance needed in the KPSS stationarity test of *y*. Quite large a lag length will then be necessary in the KPSS test on y_t to avoid over-rejection, with all the danger of losing power this entails. Likewise, finding $\Delta y_t \sim I(0)$ in the ADF test justifies the assumptions that the ACF for Δy_t decays quickly and that a smaller value of *k* will do and power will be saved in the KPSS test on Δy_t. Second, the number of lagged first differences required in the ADF equation to yield auto-correlation-free residuals gives an idea of how severe the trade-off between size and power could be for a time series integrated in levels ($\rho = 1$). A high number points to auto-correlation in ε_t,[32] a quite high value of σ_ε^2, and low power of the KPSS test. An intermediary value of *k* would

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31 Another general rule is that problems of excessive size and small power are more pronounced in KPSS tests for de-trended data than in those for de-meaned data.

32 Any AR(∞) process can be approximated by an MA(1) process.

therefore be advisable. Turning to unit root and stationarity tests for the first difference of a time series, Δy_t, there does not seem to exist any relevant problem of auto-correlated errors once the ADF test has found $\Delta y_t \sim I(0)$ using few lagged first differences. Given this "prior" in favor of difference stationarity and thus small values for σ_u^2 and λ, one can assume that size problems are less of a concern than power problems and prefer a rather small *k*. All in all, a high *k* makes sense to correct the size problem, but at, say, *k=24*, there is then a risk of accepting stationarity although it is wrong.

4.2.2. Results

The results of the KPSS tests for the United States, Canada and Germany which are documented in Tables 6, 7 and 8 are in contradiction to or at least much less clear-cut than the corresponding ADF findings for many time series. But unlike what could be expected from the problematic power properties of both types of tests, there is no pervasive tendency of the KPSS test finding lower orders of integration than the ADF test. The results rather deviate in both directions from the central I(1) finding of the ADF test. Five out of 35 variables turn out to be stationary, 16 are I(1) and 14 are found to be I(2). However, some qualifications to these results have to be made.

First, the choice of *k*, the truncation parameter, is crucial to the outcome and particularly so in the tests for I(1) represented in the right half of the tables. Here the 10 percent critical value is comprised within the interval spanned by the lowest and the highest of the 25 values of the KPSS statistic (for *k=0* to *24*) in no less than 12 out of 35 cases. An arbitrary decision-making procedure would pick only those lags which lead to non-rejection of the null and thus to the "desired" order 1 of integration. This would eliminate I(2)-ness in 5 cases bringing the number of I(2) variables down to 9. However, I commit myself to restrict arbitrariness along the lines described above because economic judgement (Table 2) and the results of the ADF tests (Tables 3, 4 and 5) provide some priors on the nature of the time series under investigation which in combination with knowledge on power and size properties of the KPSS test lets some *k* values appear more reasonable than others. In case of switching test outcomes, this range of reasonable *k* values is highlighted by bold figures in the lines of Tables 6, 7 and 8. Yet due to uncertainties, the range is chosen quite large.

Second, if a quadratic time trend is again allowed for *f*, the world trade intensity, Δf can be considered to be stationary around a linear trend thereby reducing the number of I(2) variables still further (from 9 to 6). Third, as discussed in the above sub-section on ADF test results, the finding of trend-stationary German exports of goods and services may be caused by the reunification boom and the redefinition of the German territory in 1991 and is probably a statistical artifact. In a similar vein, allowing for a linear trend in the KPSS test on *e* would lead to rejection of the null. Declaring German exports and its real effective exchange rate integrated of order one would decrease the number of stationary variables to three. Taking it all together, 26 out of 35 variables could be found to be I(1) if the pitfalls and uncertainties of the KPSS test procedure were applied less carefully. Then the differences in outcomes between ADF and KPSS tests would be much smaller.

As to the results country by country, they are still quite satisfying for the United States (Table 6): two variables are found to be trend-stationary due to non-rejection of the null in the Test for I(0). These are the index of foreign industrial production and potential output. The test for I(1) rejects in four cases (trade intensity, U.S. producer prices and U.S. export prices — both expressed in U.S. dollars — as well as foreign producer prices in foreign currency units). The remaining six time series are integrated of order one.

Table 6: Results of the KPSS Stationarity Tests – United States

Variable	Test for I (0)[a]					Test for I (1)[b]					Result
	Specification[c]	k=6	k=12	k=18	k=24	Specification[c]	k=6	k=12	k=18	k=24	
x	T	0.245***	0.155**	0.128*	0.121*	C	0.163	0.147	0.160	0.189	I (1)
y^*	T	0.080	0.066	0.073	0.096	C	0.088	0.110	0.142	0.186	I (0)
$\tilde{y}$	T	0.110	0.076	0.074	0.087	C	0.187	0.123	0.110	0.118	I (0)
f	T	0.378***	0.226***	0.176**	0.155**	T	0.151**	**0.116**	**0.118**	0.128*	I (1)[d]
						C	1.074***	0.656**	0.505**	0.427*	I (2)
w	T	0.226***	**0.140***	**0.115**	0.110						
	C	1.433***	0.827***	0.612**	0.508**	C	0.104	0.084	0.087	0.099	I (1)[e]
p_{NCU}	T	0.309***	0.191**	0.156**	0.146**	C	**0.671****	**0.439***	0.369*	0.344	I (2)
q_{NCU}	T	0.302***	0.190**	0.158**	0.150**	C	0.615**	0.432*	0.379*	0.365*	I (2)
p^*	T	0.398***	0.235***	0.179**	0.154**	C	1.065***	0.698**	0.524**	0.433*	I (2)
p	T	0.341***	0.202**	0.156**	0.137*	C	**0.415***	**0.307**	0.273	0.257	I (1)
q	T	0.350***	0.207**	0.159**	0.139*	C	**0.464****	**0.345**	0.300	0.275	I (1)
p^*_{NCU}	T	0.209**	0.141*	0.130*	0.139*	C	0.313	0.254	0.262	0.297	I (1)
e	C	1.300***	0.754***	0.560**	0.463*	C	0.089	0.076	0.082	0.097	I (1)[f]

* (**, ***) means rejection at the 10% (5%, 1%) significance level. Bold cells show truncation parameters k more sensible than others and within the range of these k values the result leading to I(1) of the level is chosen in ambiguous cases (see text for discussion). – [a]Test on level of time series. – [b]Test on first difference of time series. – [c]T: Null is trend-stationarity against the alternative of non-stationarity with drift. Use of the KPSS_τ test statistic as series is detrended before tested. C: Null is mean-stationarity against the alternative of non-stationarity without drift. Use of the KPSS_μ test statistic. The truncation parameter k indicates the order of residual covariances taken into account in the computation of the long-run variance, ω_ε^2. – [d]When allowing for a quadratic trend in f, the null of stationary Δf cannot be rejected for truncation parameters k from 11 to 19. – [e] KPSS_τ test on w is between rejection and non-rejection, but according to Table 1, KPSS_μ test is more relevant here. – [f] Even when allowing for a deterministic trend, the null of a trend-stationary e is rejected for $k<8$.

Source: Own calculations.

Turning to Canada (Table 7), only 11 variables remain to be analyzed because of the identity of p^* with the U.S. p_{NCU}. Stationarity in levels is rejected for all variables. The test for I(1) leads to rejection of the null of stationary first differences in three cases: potential growth, domestic producer prices in Canadian dollars, and world trade intensity. With the former two being relatively close to I(1) in levels and with the rate of growth in world trade intensity being stationary around a deterministic trend (when admitted), the Canadian data set is the most satisfying both with respect to the I(1)-ness in variables as found by the KPSS test and with respect to consistency of the KPSS test results with those from the ADF test.

Table 7: Results of the KPSS Stationarity Tests – Canada

Variable	Test for I (0)[a]					Test for I (1)[b]					Result
	Speci-fica-tion[c]	k=6	k=12	k=18	k=24	Speci-fica-tion[c]	k=6	k=12	k=18	k=24	
x	T	0.252***	0.168**	0.146*	0.140*	C	0.215	0.240	0.291	0.338	I (1)
y^*	T	0.188**	0.130*	0.119*	0.124*	C	0.137	0.147	0.182	0.219	I (1)
$\tilde{y}$	T	0.316***	0.195**	0.160**	0.142*	C	**0.610****	**0.403***	0.337	0.304	I (2)[d]
f	T	0.394***	0.233***	0.179**	0.155**	T	0.153**	**0.118**	**0.116**	0.122*	I (1)[e]
						C	1.244***	0.743***	0.554**	0.454*	I (2)
w	T	**0.150****	**0.100**	0.088	0.088						
	C	1.015***	0.637**	0.516**	0.465**	C	0.102	0.096	0.101	0.116	I (1)[f]
p_{NCU}	T	0.325***	0.198**	0.158**	0.144*	C	0.707**	0.463*	0.388*	0.349*	I (2)[g]
q_{NCU}	T	0.274***	0.171**	0.142*	0.133*	C	**0.452***	**0.328**	0.293	0.277	I (1)
p^*	T	0.309***	0.191**	0.156**	0.146**	C	**0.671****	**0.439***	0.369*	0.344	I (2)
p	T	0.313***	0.200**	0.168**	0.156**	C	**0.363***	**0.328**	0.356*	0.392*	I (1)[i]
q	T	0.270***	0.179**	0.156**	0.153**	C	0.269	0.260	0.287	0.326	I (1)
p^*_{NCU}	T	0.265***	0.164**	0.134*	0.125*	C	**0.477****	**0.334**	0.275	0.252	I (1)[j]
e	T	**0.133***	**0.096**	0.091	0.094	C					
	C	**0.474****	**0.333**	0.307	0.308	C	0.072	0.078	0.099	0.129	I (1)[k]

* (**, ***) means rejection at the 10% (5%, 1%) significance level. Bold cells show truncation parameters k more sensible than others and within the range of these k values the result leading to I(1) of the level is chosen in ambiguous cases (see text for discussion). – [a]Test on level of time series. – [b]Test on first difference of time series. – [c]T: Null is trend-stationarity against the alternative of non-stationarity with drift. Use of the KPSS_τ test statistic as series is detrended before tested. C: Null is mean-stationarity against the alternative of non-stationarity without drift. Use of the KPSS_μ test statistic. The truncation parameter k indicates the order of residual covariances taken into account in the computation of the long-run variance, ω_ε^2. – [d]Result close to I (1). – [e]When allowing for a quadratic trend in f, the null of stationary Δf cannot be rejected for truncation parameters from 12 to 21. – [f]If the d.g.p. for Δw contains an intercept (as suggested in Table 1), the KPSS_τ test only rejects the null up to truncation parameter 8 (result close to I (0)). – [g]Result close to I (1) (10% critical value: 0.347). – [h]The null of stationary Δq_{NCU} is not rejected up to truncation parameter 9. – [i]Result close to I (2). The null of stationary Δp^* is not rejected for truncation parameters 8 through 16. – [j]Result close to I (2). Null of stationary Δp^*_{NCU} is rejected for k from 0 to 11. – [k]Result close to I (0): The KPSS_τ test for e rejects null for k from 0 through 8, the KPSS_μ test for k from 0 to 10.

Source: Own calculations.

Finally, the results are least satisfying for Germany (Table 8). In the test for I(0) (left half of table) the null cannot be rejected for the levels of three variables (exports, foreign industrial production and the real effective exchange rate). In turn, the null of stationary first differences is rejected in seven cases by the test for I(1) (right half of table). This leaves us with only two I(1) variables (potential output and the index of foreign producer prices expressed in deutsche mark).

Table 8: Results of the KPSS Stationarity Tests – Germany

Variable	Test for I (0)[a]					Test for I (1)[b]					Result
	Specifica-tion[c]	k=6	k=12	k=18	k=24	Specifica-tion[c]	k=6	k=12	k=18	k=24	
x	T	0.081	0.067	0.076	0.096	C	0.124	0.156	0.195	0.227	I (0)
y^*	T	0.104	0.078	0.081	0.103	C	0.132	0.142	0.172	0.217	I (0)[d]
$\tilde{y}$	T	0.192**	**0.123***	**0.104**	0.102	C	0.107	0.099	0.101	0.110	I (1)[e]
f	T	0.386***	0.229***	0.176**	0.154*	T	0.155**	0.119*	**0.118**	**0.123***	I (1)[f]
						C	1.198***	0.719**	0.537**	0.442*	I (2)
w	T	0.232***	0.171**	0.157**	0.156**						
	C	1.395***	0.822***	0.610**	0.500**	C	0.437*	0.472**	0.423*	0.423*	I (2)
p_{NCU}	T	0.321***	0.196***	0.157**	0.144**	C	**0.527****	**0.393***	0.340	0.325	I (2)[g]
q_{NCU}	T	0.343***	0.208**	0.163**	0.146**	C	**0.525****	**0.412***	0.343	0.309	I (2)[g]
p^*	T	0.363***	0.219***	0.171**	0.151**	C	0.985***	0.629**	0.488**	0.414*	I (2)
p	T	0.350***	0.223***	0.181**	0.163**	C	0.960***	0.739***	0.574**	0.500*	I (2)
q	T	0.351***	0.226***	0.184**	0.166**	C	0.858***	0.707**	0.559**	0.504**	I (2)
p^*_{NCU}	T	0.274***	0.171**	0.140*	0.128*	C	0.158	0.161	0.155	0.159	I (1)
e	C	0.197	0.133	0.115	0.115	C	0.100	0.119	0.115	0.129	I (0)[h]

* (**, ***) means rejection at the 10% (5%, 1%) significance level. Bold cells show truncation parameters k more sensible than others and within the range of these k values the result leading to I(1) of the level is chosen in ambiguous cases (see text for discussion). – [a]Test on level of time series. – [b]Test on first difference of time series. – [c]T: Null is trend-stationarity against the alternative of non-stationarity with drift. Use of the KPSS_τ test statistic as series is detrended before tested. C: Null is mean-stationarity against the alternative of non-stationarity without drift. Use of the KPSS_μ test statistic. The truncation parameter k indicates the order of residual covariances taken into account in the computation of the long-run variance, ω_ε^2. – [d]Stationarity of y^* is rejected only for truncation parameters k=1 through 4. – [e]Result close to I (0). – [f]When allowing for a quadratic trend in f, the null of stationary Δf cannot be rejected for k from 13 to 20. – [g]Result close to I (1) (10% critical value: 0.347). – [h]Stationarity of e is rejected only for k=0, 1 and 2. However, allowing for a linear trend in the test equation would lead to rejection of the null.

Source: Own calculations.

4.3. Synopsis of Results from ADF and KPSS tests

The comparison between the ADF and the KPSS test results proves that for most of my time series determining the correct order of integration is difficult. What has been said about Nelson and Plosser's original macroeconomic data also applies to my case: "…the appropriate conclusion is that the data are not sufficiently informative to distinguish between these hypotheses [trend-stationarity versus unit root with drift]." (Kwiatkowski et al. 1992: 175). Putting all peaces of evidence together, one clearly stationary variable is found (foreign industrial production from the perspective of the United States), and two variables undoubtedly contain two unit roots: Foreign producer prices from the German perspective and Canada's producer price index in Canadian dollars. More generally, national producer price indices expressed in domestic currency tend to be I(2) according to the KPSS test results. Yet the ADF test in case of Germany provides clear evidence in favor of I(1) and is not clear-cut in case of the United States where the result is sensitive to the specification of the deterministic structure. Only seven variables are clearly integrated of order one according to both tests (Table 9). The bulk of the empirical variables used (25 of the 35) are either found not to have the same number of unit roots according to the ADF or the KPSS test, respectively, or are characterized by somewhat ambiguous results of the KPSS test.

The existence of a stationary variable in cointegrating vectors does not harm as long as there are still at least two variables containing the same number of unit roots (1 or more) which can co-integrate at a lower degree of integration. My attention therefore turns to the variables which can hardly be decided on to be I(1) but are rather I(2) according to the procedures outlined and the empirical results found in this section. Besides world trade intensities for all three countries, these are the index of foreign producer prices for Germany, the U.S.

index of producer prices in U.S. dollars (see p_{NCU} in Tables 3 and 6 as well as p^* in Tables 4 and 7) and the index of Canadian producer prices in Canadian dollars. The normal therapy would consist of entering the inflation rate and its first difference into the VECM to deal with I(1) variables in the cointegration space and with I(0) variables in the short-run relationships. However, this therapy is not applied for two reasons.

Table 9: Synopsis of Results from ADF and KPSS Tests

	Germany	United States	Canada
Variables clearly I (0)		y^*	
Variables clearly I (1)	p^*_{NCU}	x, w, p^*_{NCU}, e	x, y^*, q
Variables clearly I (2)	p^*		p_{NCU}
Variables with uncertain results[a]	$x(1)$, $y^*(1)$, $\tilde{y}(1)$, $f(2)$[b], $w(1)$, $p_{NCU}(1)$, q_{NCU}, $p^*(1)$, $p(1)$, $q(1)$, $e(1)$	$\tilde{y}(1)$, $f(2)$[b], $p_{NCU}(2)$, $q_{NCU}(1)$, $p^*(1)$, $p(1)$, $q(1)$	$\tilde{y}(1)$, $f(2)$[b], $w(1)$, $p^*(2)$, $q_{NCU}(1)$, $p(1)$, $p^*_{NCU}(1)$, $e(1)$

[a]The number in brackets is the degree of integration decided on (see text for discussion). – [b]If quadratic trend is allowed for in the levels, KPSS test does not reject the null of Δf being stationary.

Source: Own calculations.

First, the theoretical relationships generally imply combinations of price level variables to be the relevant determinants. Export demand is commonly modeled with the real effective exchange rate of the domestic currency as the proxy for price competitiveness of the domestic export sector, involving the linear combination $q_{NCU} - \beta_1 p^*_{NCU}$ or $q - \beta_2 p^*$ with β_1 and β_2 generally restricted to one. Likewise, the most important determinant of export supply is the relative profitability of exporting involving $q_{NCU} - \beta_3 p_{NCU}$ or $q - \beta_4 p$ in the cointegration space with β_3 and β_4 often restricted to one. From an econometric modeling perspective both price variables in each relationship should be entered either in levels *or* in first differences, intermediate solutions make no sense economically. Only for Canada could one consider to use inflation rather than price level differences as p_{NCU} is found to be I(2) and p^{*}_{NCU} seems to be close to I(2) according to the KPSS test such that a majority of price variables needed in a VECM of export demand and supply (two out of three) would not be difference-stationary. For both Germany and the United States, in turn, only one out of six price variables is I(2) such that modeling the whole system in national currency units for Germany and in foreign currency units in case of the United States would completely avoid the use of I(2) variables.

Second, international comparability of results is one core objective of this study. A unified framework for the initial system estimation is therefore recommendable. As the evidence for price indices used being I(1) dominates in the current country sample, I decide to treat these variables as if they were I(1) throughout. As to the log of world trade intensity, it is also treated as being I(1). This decision is motivated on economic grounds. It is plausible to expect that a simple increase in world exports per unit of world GDP (and not only an acceleration of this increase) positively affects the long-run level of exports of a country which participates in the international division of labor.

4.4. A Multivariate Test for Stationarity

As the ADF and KPSS tests do not lead to clear answers on the degree of integration of the time series used, the question of stationarity is additionally addressed within the multivariate frameworks explored in more detail in the next section. Since the number of cointegration relations increases for each stationary variable included in the cointegration space, the test outcome is useful for finding the minimal set of variables needed for cointegration. Conditional on the number of stationary relationships in the cointegration space, *r*, it is tested for each variable if the data support the $(p * r)$-matrix of cointegrating vectors to be partitioned into one stationary time series (the variable at hand) and a

$(p*(r-1))$–submatrix φ of the other stationary relationships. The cases $r=0$ and $r=p$ (p being the number of variables) automatically imply, respectively, that all variables are non-stationary and that all variables are stationary and are therefore not reported. First the variables of the system enter in levels, which corresponds to a test for I(0). Then the same analysis is done with the variables entering in first differences to test difference-stationarity against the hypothesis of two unit roots in variables (test for I(1)). To save space, only the specification which turns out to yield best results in the cointegration analysis in section 5 below is reported. It contains six variables for each country: real exports, foreign industrial production, world trade intensity and the three price indices (domestic export prices, domestic and foreign producer prices), all expressed in units of the export-weighted foreign currency basket.[33]

The results of the test for I(0) are shown in Table 10. In case of the U.S. variables, stationarity can be ruled out whatever the rank will be found to be. This somehow revokes the earlier finding of stationary foreign industrial production (Table 9) and the ADF test result of a stationary domestic producer price index (Table 3). For Germany the null hypothesis of stationarity can be rejected for each variable as long as the number of cointegration relationships does not exceed 2. Only if $r \geq 3$, some of the variables seem to be stationary. The results for Canada are similar to the ones for Germany insofar as stationarity can be rejected for all time series as long as the number of cointegration relationships does not exceed 2.

33 Including potential GDP in the VECM is going to yield unsatisfactory results and its exclusion from the model will be justified by over-identifying restrictions (see below).

Table 10: Multivariate Tests for Level-Stationarity of Time Series in the Set (x, y^*, f, p, q, p^*) for Germany, the United States and Canada

Rank r	$p-r$	x	y^*	f	p	q	p^*
United States							
1	5	38.52**	41.01**	49.02**	39.14**	39.02**	35.18**
2	4	14.89**	16.03**	23.18**	15.97**	15.48**	12.33**
3	3	13.05**	14.16**	21.27**	13.89**	13.42**	10.58**
4	2	11.79**	14.05**	13.08**	9.60**	9.03**	8.32**
5	1	4.90**	5.27**	5.16**	3.27*	2.82*	3.65**
Canada							
1	5	27.00**	27.01**	32.61**	19.92**	19.03**	21.31**
2	4	17.29**	18.02**	22.50**	11.31**	10.87**	11.32**
3	3	10.41**	11.11**	16.02**	4.93	4.65	4.57
4	2	7.76**	8.27**	12.35**	0.89	0.32	0.73
5	1	4.04**	4.15**	6.14**	0.31	0.30	0.48
Germany							
1	5	20.33**	22.00**	35.47**	26.94**	26.62**	26.05**
2	4	12.86**	12.56**	26.84**	17.38**	17.04**	16.30**
3	3	3.87	6.22	14.25**	5.48	5.78	2.75
4	2	3.81	5.37*	12.09**	5.48*	5.77*	2.75
5	1	2.22	1.60	5.02**	4.61**	4.49**	2.66

*(**) means rejection at the 10 (5) percent significance level. The test statistic is χ^2-distributed with $(p-r)$ degrees of freedom ($p=6$ being the number of variables in this context). The critical values can be found e.g. in Bamberg and Schittko (1979: 193).

Source: Own calculations.

As far as the tests for I(1) are concerned, the results of the multivariate stationarity tests are far less encouraging and rather strengthen the doubts shed on a feasible distinction between I(1) and I(2) for the time series used. Unsurprisingly, the null of a difference-stationary f is rejected throughout (Table 11). Even if one interpreted only the last two lines ($r=4, r=5$) (based on the assumption that the theory backing our analysis is formulated for levels of variables and that therefore most first differences should be stationary), all price indices are found to be I(2) in every country, although rejection is somewhat less clear for the United States. In Canada, even real exports seem to be I(2).

Table 11: Multivariate Tests for Difference-Stationarity in the Set (x, y^*, f, p, q, p^*) for Germany, the United States and Canada

Rank r	$p-r$	Δx	Δy^*	Δf	Δp	Δq	Δp^*
United States							
1	5	19.47**	9.78*	39.46**	31.24**	32.80**	37.85**
2	4	11.04**	3.42	31.03**	22.64**	24.76**	29.56**
3	3	6.81*	2.73	23.69**	15.10**	17.40**	22.41**
4	2	3.36	1.02	20.89**	14.04**	15.93**	19.04**
5	1	0.59	0.53	8.71**	3.09*	3.55*	8.28**
Canada							
1	5	22.92**	24.09**	50.68**	39.32**	36.15**	46.04**
2	4	19.05**	12.49**	39.03**	28.49**	24.63**	34.57**
3	3	13.27**	9.96**	28.20**	18.64**	14.43**	24.87**
4	2	6.44**	5.13*	14.81**	6.80**	6.97**	13.05**
5	1	4.22**	2.50	12.76**	6.80**	6.80**	10.75**
Germany							
1	5	23.01**	21.88**	51.34**	41.93**	42.54**	51.92**
2	4	9.77**	21.41**	34.59**	27.15**	27.23**	35.31**
3	3	3.38	7.18*	19.11**	11.90**	11.84**	19.56**
4	2	0.87	3.45	10.87**	10.65**	8.75**	11.24**
5	1	0.25	1.05	9.57**	7.39**	5.13**	9.17**

*(**) means rejection at the 10 (5) percent significance level. The test statistic is χ^2-distributed with $(p-r)$ degrees of freedom ($p=6$ being the number of variables in this context). The critical values can be found e.g. in Bamberg and Schittko (1979: 193).

Source: Own calculations.

According to these results, stationarity in levels does not seem to be a problem for the time series used here. In turn, the existence of a second unit root in the levels of the time series appears to be quite more common than according to univariate unit root or stationarity tests. The conclusion from these multivariate tests for stationarity is that they find a higher degree of integration than their univariate counterparts in many cases. Therefore the results of this sub-section do not corroborate the findings of the ADF and KPSS tests but rather intensify the doubts regarding the feasibility of a clear-cut distinction between (difference) stationarity and non-stationarity in empirical applications. As virtually all the time series turn out to be I(1) at least according to the ADF test, I maintain the hypothesis that cointegration analysis of the I(1) type is the appropriate tool of analyzing the economic relationships between the time series at hand, i.e. an empirical model with first differences as the dependent variables and long-run equilibrium relationships expressed in levels. The focus will be on the long-run economic relationships. The objective of the next section therefore consists of finding and identifying the cointegration relationships in the system.

5. The Long-Run Structure of the Aggregate Export System

5.1. The VECM As a Suitable Type of Empirical Model

Owing to the non-stationarity of the data, the demand and the supply of real exports are estimated in a vector error correction model (VECM) framework based on the procedure developed by Johansen (1988 and 1991) as well as Johansen and Juselius (1992 and 1994). This approach is suited to detect stationary linear combinations (i.e. cointegration relationships) between I(1) variables. These relationships are interpreted as the long-term economic equilibrium relationships and should therefore correspond to the hypotheses derived from theoretical considerations. The starting point is the multivariate vector autoregressive (VAR) model of the levels of the I(1) variables at lag length l:

[13a] $z_t = A_1 z_{t-1} + A_2 z_{t-2} + \ldots + A_l z_{t-l} + \mu + \delta t + u_t,$

with u_t being the $(n \times 1)$-vector of independently normally distributed errors, μ the $(n \times 1)$-vector of constant terms, δ the $(n \times 1)$ coefficient vector of a linear deterministic time trend, and A_i the $(n \times n)$-matrices of coefficients. This model is equivalent to the VECM (see e.g. Harris (1995: 77)):

[13b] $\Delta z_t = \Gamma_1^* \Delta z_{t-1} + \ldots + \Gamma_l^* \Delta z_{t-l+1} + \Pi^* z_{t-l} + \mu + \delta t + u_t,$

where $\Gamma_i^* = -(I - A_1 - \ldots - A_i),\ i = 1,2,\ldots,(l-1),$

and $\Pi^* = -(I - A_1 - \ldots - A_l)$,

By a slightly different transformation, equation [13a] can be transformed into a VECM dependent on the levels in period $(t-1)$, which is more useful for my purposes:

[13c] $\Delta z_t = \sum_{i=1}^{l-1} \Gamma_i \Delta z_{t-i} + \Pi z_{t-1} + \mu + \delta \cdot t + u_t$,

where $\Gamma_i \equiv -(A_{i+1} + \ldots + A_l), i = 1,2,\ldots,(l-1)$, and $\Pi \equiv -(I - A_1 - \ldots - A_l) = \Pi^*$.

z' is the $(T \times n)$-matrix of I(1) variables (T being the number of observations, n the number of I(1) variables), Δ the difference operator, Γ_i the $(n \times n)$-vector of short-run coefficients of the first differences of variables in period $t-i$. The z-variables with a time subscript thus represent the $(n \times 1)$-vector of realizations of all I(1) variables in the period indicated. In this initial representation I assume all I(1) variables are endogenous, i.e. $p = n$, where p in this context stands for the number of endogenous I(1) variables.

The economic hypotheses of interest are formulated as restrictions on Π, the $(n \times n)$-coefficient matrix of once-lagged levels, whereas the other parameters are allowed to vary without restrictions. The short-run part of the model is thus of reduced form, i.e. contemporary first differences of endogenous variables do not show up anywhere on the right-hand side of the p equations. As to matrix Π, the case where Π has reduced rank r, $0 < r < p$ is of particular interest. In this case Π can be decomposed into a $(p \times r)$-matrix α of adjustment coefficients and a $(p \times r)$-matrix β of cointegrating vectors such that $\Pi = \alpha\beta'$, as will be seen in more detail in the sub-section on the rank test below.[34] The

[34] The case $r = 0$ means that there is no stationary linear combination between the I(1) variables implying that a VECM does not yield any additional information compared to a vector auto-regressive model (VAR) in first differences. In turn, the case $r = p$ can only occur if all p variables are in fact stationary in levels. The matrix of cointegrating vectors, β, then reduces to a $(p \times p)$-identity matrix.

$(r \times 1)$-vector $\beta' z_{t-1}$ represents the r cointegration relationships each of which equals zero in equilibrium.

In my application, z consists of the variables typically used in the analysis of export demand, $x, q_{NCU}, p^{*}_{NCU}, y^{*}$, of the variables additionally needed to capture the supply of exports, respectively p alone and the set $p_{NCU}, \tilde{y}$, and of the globalization variable f. Altogether, z thus contains six or seven time series. There are at least two reasons why the single-equation error-correction model (SEECM) approach widely used in the empirical literature on the determinants of exports cannot yield satisfying results given the modeling strategy (to describe both supply *and* demand forces) and the choice of variables at hand. The first reason is that there is probably more than one cointegrating vector. In this case estimating a SEECM with Δx_t as the left-hand side variable would allow to capture only Π_1, the first row of the Π matrix. Referring to the case where z' is of dimension $T \times 6$ and assuming that there are two cointegration relationships ($r = 2$), neither of these would then emerge distinctly from the estimation since Π_1 is a linear combination of both vectors (Harris (1995: 63)). As can be seen in equation [14], the estimate showing up as a coefficient in front of each lagged variable in levels is a linear combination of two cointegrating vectors which cannot be singled out, such that the scalar $\Pi_1 z_{t-1}$ reads:

[14]

$$\Pi_1 z_{t-1} = \left[(\alpha_{11}\beta_{11} + \alpha_{12}\beta_{12}) \;\; (\alpha_{11}\beta_{21} + \alpha_{12}\beta_{22}) \;\; (\alpha_{11}\beta_{31} + \alpha_{12}\beta_{32}) \;\; (\alpha_{11}\beta_{41} + \alpha_{12}\beta_{42}) \;\; (\alpha_{11}\beta_{51} + \alpha_{12}\beta_{52}) \;\; (\alpha_{11}\beta_{61} + \alpha_{12}\beta_{62}) \right] \begin{bmatrix} x_{t-1} \\ q_{t-1} \\ p^{*}_{t-1} \\ y^{*}_{t-1} \\ f_{t-1} \\ p_{t-1} \end{bmatrix} =$$

$$= \alpha_{11}\left(\beta_{11}x_{t-1} + \beta_{21}q_{t-1} + \beta_{31}p^{*}_{t-1} + \beta_{41}y^{*}_{t-1} + \beta_{51}f_{t-1} + \beta_{61}p_{t-1}\right) + \alpha_{12}\left(\beta_{12}x_{t-1} + \beta_{22}q_{t-1} + \beta_{32}p^{*}_{t-1} + \beta_{42}y^{*}_{t-1} + \beta_{52}f_{t-1} + \beta_{62}p_{t-1}\right).$$

The second reason for the inappropriateness of the SEECM is that the right-hand side variables are probably not weakly exogenous, at least not all of them. In the context of the VECM, a variable is called weakly exogenous to the system if disequilibrium changes (deviations of the cointegration relations from zero) do not affect the variable, i.e. there is no feedback from the disequilibrium back on the level of this variable. By limiting the analysis to a SEECM with Δx_t as the left-hand side variable, one would waste information on the determination of all variables which are not weakly exogenous. As shown by Johansen (1992a), in this case the elements of the cointegrating vector β have a higher variance than in the VECM. In other words, the estimates of the long-run coefficients are not efficient, although they are at least consistent. Consistency of $\hat{\beta}$, i.e. $\underset{T \to \infty}{p\lim}\,\hat{\beta} = \beta$, arises thanks to the "superconsistency" property of the OLS estimator when non-stationary time series are cointegrated because then, as T grows larger, $\hat{\beta}$ converges to its true value at a faster rate than with stationary variables (Stock 1987). However, there is a bias in small samples which is likely to be relevant in my case ($T \approx 100$) and should therefore be avoided.[35]

The Johansen procedure with its VECM setting avoids these problems by allowing to determine the number of cointegrating vectors, to impose an economically meaningful structure on them by identifying restrictions and to formally check for weak exogeneity of variables to the system in the multivariate context. However, before the cointegration analysis itself can be carried out, the precise shape of the unrestricted initial VECM has to be decided on.

35 Inder (1993) finds by Monte Carlo simulations with I(1) variables in a bivariate rational distributed lag Model (RDLM) at lag length two that the ECM performs much better than the widely used two-step procedure proposed by Engle and Granger (1987) and the fully modified OLS estimator by Phillips and Hansen (1990). Yet even Inder reports parameter constellations for $T = 50$, where the estimate of the long-run coefficient is within 0.05 of the true value (=1) in only 50 or 60 percent of times (Inder, op. cit.: 62).

5.2. Deterministic Elements, Lag Length, and Dummy Variables of the VECM

Important preliminary specification choices regard assumptions on the deterministic elements (μ and δ in [11]) as well as the lag length l of the model, i.e. how many periods into the past are taken into account in the determination of z_t. The reason is that the number of cointegrating vectors, the estimates of the economically interesting long-run elasticities and, of course, the behavior of the residuals depends on these choices. To avoid wrong outcomes and to minimize arbitrariness, these choices should be sound and suitable in many respects.

5.2.1. Deterministic Elements

As to the deterministic structure of the VECM, five different settings are conceivable given the structure outlined in [13] (see e.g. Boswijk (1994: 57–58)). The most restrictive one ("model 1") is where one sets $\mu = 0$ and $\delta = 0$ by assumption. "Model 2" is less restrictive setting $\delta = 0$ and $\mu \neq 0$ but the intercept only shows up in the cointegration space (i.e. the I(1) variables cointegrate around a constant term). This is achieved by decomposing μ into

[15] $\mu = \alpha\mu_1 + \alpha_\perp\mu_2$,

and by then imposing $\mu_2 = 0$ (Juselius (1993: 10)). As $\alpha_\perp$ is the $(p \times (p-r))$-matrix orthogonal to α defining the space of the common stochastic trends and the intercept of the once-differenced data restrictions on parts of the Π matrix, this means that μ does not account for linear trends in the data in "model 2". In light of the reasoning on the deterministic structure of the time series composing the VECM in the section on unit root tests, models 1 and 2 are far too restrictive as both visual inspection and economic reasoning point to a deterministic trend in most of the time series used. Therefore, at least μ should remain unrestricted ($\mu_1 \neq 0, \mu_2 \neq 0$) allowing for a linear trend in the data and for intercepts in the cointegration relationships.[36] This is done in "model 3" which still maintains $\delta = 0$. "Model 4" is the first one to allow $\delta \neq 0$ in addition to an unrestricted μ but limits the deterministic trend to the cointegration space by setting $\delta_2 = 0$ after decomposing δ in an analogous manner to [15]. Finally, "model 5" is completely unrestricted with respect to deterministic elements as $\alpha_1, \alpha_2, \delta_1, \delta_2$ may all differ from zero thus accounting for a quadratic trend in the data. It is hard to imagine logarithms of economic variables growing at a deterministic quadratic pace in the long run although such a pattern might prevail for a certain period in time. Specifically, merchandise trade data for many countries exhibited such a pattern during the last 15 years of the twentieth century, and this higher speed in the expansion of world trade has become one stylized fact of globalization. As this effect is accounted for by the insertion of the globalization variable f into the VECM, I can be rather confident that the episode of (near-) quadratic growth in exports does not need to receive further attention in the deterministic part of the VECM such that "model 5" can be ruled out.

The decision has to be made between "model 3" and "model 4". The latter would be advisable if the structural relationships for quantities and prices of exports contained some deterministic increase or decrease not fully explained by the variables of the model. This is a relevant case as modeling exports with a time trend in the cointegration relationship is the textbook standard (see e.g. Whitley (1994: 91)) and has been widely used by applied researchers especially in cases where foreign output (rather than the volume of world trade) is the proxy for economic activity abroad (Döpke and Fischer (1994), Strauß (1998, 1999 and 2000)). In export quantity relationships the trend most commonly stands for the ongoing intensification of the worldwide division of labor

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[36] However, the latter are not estimated explicitly since only one $\hat{\mu}_i$-value is estimated per equation. This saves the VECM degrees of freedom but leaves a certain amount of ambiguity on how much of $\hat{\mu}_i$ belongs to the cointegration space and how much is the slope of a trend.

showing up in outsourcing, global sourcing and multinational networking encouraged by falling transport and communication costs and multilateral trade liberalization. Likewise, the deterministic trend may also show up in the export price relationship, but with a negative sign. However, the effects of globalization are captured here by variable f, which not only is economically more meaningful but also promises to yield a statistically superior specification since it is more flexible a proxy for globalization than the linear time trend.[37] This is why I do not see any further need of implementing a linear trend in the VECM at hand. But as there is little doubt as to the existence of linear trends in the data, the VECM should contain an unrestricted constant. "Model 3" is therefore chosen throughout the following sub-sections.[38]

5.2.2. The Appropriate Lag Length of the VECM and Necessary Impulse Dummy Variables

Turning to the lag length of the model, a good strategy is minimization of some information criterion subject to achieving Gaussian residuals. Information criteria have been developed as a decision device to solve the trade-off between improving the fit of the model (which would require additional lags) and granting a sufficiently high number of degrees of freedom (which would require a parsimonious parameterization) as one generally does not have observations galore in the analysis of quarterly data. In analogy to R^2, the goodness of fit can be expressed by the residual variance in the single-equation context and by the determinant of the $(p \times p)$-dimensional variance-covariance matrix Ω in the multi-equation context. It is straightforward to let $|\Omega|$ be part of the information

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[37] It is shown in Strauß (2002b) for Germany's aggregate volumes of exports and imports between 1975 and 2000 that replacing the linear trend by f produces a sensibly smaller standard deviation of the respective dependent variable in SEECMs.

[38] As a means of sensitivity analysis, the other models will be briefly looked at when it comes to the determination of the number of cointegrating vectors because it is sometimes advocated to carry out the according rank tests without economic priors on the deterministic structure of the model (e.g. Pantula (1989)).

criterion and to design it such that smaller values will be better than larger ones. To penalize over-parameterization, one may add some expression involving κ, the number of coefficients of the VECM. As over-parameterization is the worse the scarcer observations are, κ should be related to T in some way. I briefly discuss three of the most important information criteria that have emerged from the literature. The Akaike information criterion (AIC) is defined as (Akaike 1973)[39]

[16] $AIC = \ln|\hat{\Omega}| + 2\kappa / T$

The Schwarz information criterion (SC) is defined as (Schwarz 1978)

[17] $SC = \ln|\hat{\Omega}| + \kappa(\ln T) / T$

The Hannan-Quinn information criterion (HQC) is defined as (Hannan und Quinn 1979)

[18] $HQC = \ln|\hat{\Omega}| + 2\kappa[\ln(\ln T)] / T$

They sometimes yield contradictory results. This is not surprising if one answers the following questions:

1. Given T, can there be a range of incremental improvements of fit for which increasing κ by one leads to a reduction in one of the criteria (indicating that the additional lag should be incorporated into the model) whereas another one increases?
2. Assuming that the model is already over-parameterized according to all three criteria, which one increases most strongly, i.e. penalizes over-parameterization most vigorously?

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39 For reasons of comparability, I follow the unified notation from Doornik and Hendry (1997: 291).

The answers can be given by generalizing the formulas for all three information criteria (*IC*) to

[19] $IC_i = \ln|\hat{\Omega}| + \kappa c_i$, $i = AIC, SC, HQC$,

where c_i is the "penalizing factor" dependent on T responsible for an increase (or a slower decrease) in the criterion. I get

[20] $c_{AIC} = 2/T$, $c_{SC} = (\ln T)/T$, and $c_{HQC} = 2[\ln(\ln T)]/T$.

Referring to the case $T = 100$, one gets $c_{AIC} = 0.02$, $c_{SC} = 0.0461$, and $c_{HQC} = 0.0305$.

As the incremental improvement in fit is the same for all three criteria as κ rises by 1, question 2 boils down to finding the criterion for which c_i is highest. When comparing SC to AIC, SC penalizes more strongly as soon as $T \geq 8$, when comparing SC to HQC, SC penalizes more strongly for all relevant numbers of observations ($T > 1$), and when comparing HQC to AIC, HQC penalizes more strongly as soon as $T \geq 16$. This shows that for applications involving VECMs with as much as 6 variables (and thus T exceeding 16), the Schwarz criterion penalizes most strongly and the Akaike criterion least strongly.

As to the first question, SC rises while AIC shrinks for incremental reductions in $\ln|\hat{\Omega}|$ lying within the interval $[0.02;0.0461]$, SC rises while HQC shrinks for incremental reductions in $\ln|\hat{\Omega}|$ lying within the interval $[0.02;0.0305]$, and HQC rises while AIC shrinks for incremental reductions in $\ln|\hat{\Omega}|$ lying within the interval $[0.0305;0.0461]$. Of course one cannot assert that in general there are "decreasing returns" to adding one further lag $(l+1)$ to the VECM in [13c] as it depends on the AR($l+1$) coefficients and their standard deviations. But as the terms $z_{t-1}, \Delta z_{t-i}, \Delta z_{t-j}, i \neq j$ might not be completely independent from each other, it often happens that beyond some κ, any further increase is unable to achieve an incremental improvement in fit strong enough to lower the SC still further, whereas the minimum of, say, the AIC is found at a higher κ. As with respect to both questions, the Hannan-Quinn criterion lies in between the parsimonious Schwarz and the rather generous Akaike criterion, I choose the HQC as a guide in determining the lag length but as I have a high number of regressors, I additionally report the SC.

However, the soundest choice of the lag length becomes worthless when residuals are auto-correlated, heteroscedastic or not normally distributed as the critical values relevant for most of the structural tests below only apply in presence of Gaussian residuals. If residuals are not well behaved in that sense, one might add one or more additional lags for each variable and look if the problem disappears. Alternatively, an impulse or even a step dummy might be required if unique events in economic history have an impact on, say, exports that cannot be captured by the model. So the minimization of an information criterion might turn out to be overly simplistic. Rather, a minimization subject to Gaussian residuals might be required. The choice of the lag length has therefore to be discussed case by case.

First the lag length of the VECM for the United States is determined. As the presence of $\tilde{y}$ leads to huge residual autocorrelation problems and does not allow to accept any of the economic hypotheses of interest, it is dropped from the VECM.[40] For the United States as for the other two countries, all prices are in domestic currency units as this specification turns out to produce best results. The VECM thus is composed by the variables $\{x; q_{NCU}; y^*; p^*_{NCU}; p_{NCU}; f\}$. The observation period is 1975:1 to 2000:3. As lag lengths from 2 to 5 are inspected, the effective sample is chosen to be 76:2 to 00:3 for each lag length to grant comparability of the information criteria. First attempts without dummy variables show that at whatever lag length some of the equations have outliers in

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[40] Unlike the other countries, the long-run irrelevance of potential output cannot be demonstrated by a formal likelihood ratio test (data accept over-identifying null restriction on the $\hat{\beta}_{i,\tilde{y}}$-coefficient in each cointegration relationship) in the U.S. case.

77:1, 78:2, and 80:2 and require the insertion of the three impulse dummy variables $d771; d782; d802$ which take on the value 1 in the respective quarter and equal zero in all other quarters.[41] The results of the Hannan-Quinn and the Schwarz information criteria as well as the relevant residual tests are shown in Table 12.

Table 12: Determination of the Lag Length of the VECM: United States[a]

Lag length l	$\ln\lvert\hat{\Omega}\rvert$ [b]	HQC[c]	SC[d]	Auto-correlation[e]	ARCH (l)[f]
2	–60.48	–57.49	–55.98	[0.09], [0.01], [0.52]	0.3, 0.9, 2.3, 0.1, 4.2, 2.2
3	–61.13	–57.03	–54.96	[0.03], [0.24], [0.55]	1.2, 4.9, 0.6, 2.7, 4.8
4	–61.85	–56.63	–53.99	[0.00], [0.19], [0.27]	1.5, 8.0*, 4.1, 3.9, 1.3, 17.2***
5	–62.64	–56.30	–53.10	[0.00], [0.13], [0.40]	3.7, 5.0, 6.4, 4.7, 1.0, 13.3**

* (**, ***) means rejection at the 10% (5%, 1%) significance level. – [a]All variables are treated as being endogenous (p=6). The effective sample is 76:2–00:3. Impulse dummies are needed in the following quarters: 77:1, 78:2, 80:2. Probabilities are in square brackets. – [b]Log of determinant of the estimated variance-covariance matrix. – [c]Hannan-Quinn information criterion. – [d]Schwarz information criterion. – [e]Probabilities of the following tests for autocorrelation: Ljung-Box test of autocorrelation up to order 25 ($= T/4$), Breusch-Godfrey LM-tests for first- and fourth-order autocorrelation, respectively. – [f]The six values indicate the ARCH-statistic of the test for autoregressive conditional heteroscedasticity of up to order l in each of the six equations $\{\Delta x, \Delta q_{NCU}, \Delta y^*, \Delta p^*_{NCU}, \Delta p_{NCU}, \Delta f\}$. The statistic is χ^2-distributed with l degrees of freedom. The 10 % critical values for l=2, …, 5 are 4.61; 6.25; 7.78; 9.24, respectively.

The first column indicates the lag length of the VAR in levels. For example, $l = 2$ means that the short-run relationship only contains first differences back to period $(t-1)$. It is therefore useless to analyze the case $l = 1$ because the short-run dynamics of the model would then contain nothing else than the three

[41] The outlier in 77:1 is only due to equation Δf and can be dropped later when f will be set exogenous.

dummy variables. In the second column, it can be seen that the determinant of the estimated variance-covariance matrix, an overall measure of how big the errors are, gets smaller with each lag added. Yet these incremental improvements are found to be too small, i.e. to imply too high a cost in terms of degrees of freedom lost, according to both the HQC and the SC reported in columns 3 and 4. Pure minimization of the information criteria would therefore favor a model of lag length 2, the most parsimonious one among the models examined. However, this model leads to first-order residual autocorrelation as detected by the Breusch-Godfrey Lagrange Multiplier (LM-) test, which clearly rejects the null of no autocorrelation at the 5-percent significance level (see second number in column 5).

All tests for autocorrelation are of the LM type, i.e. they are based on the auxiliary regression of the residuals $\hat{u}_{i,t}$ on the original variables and lagged residuals, $\hat{u}_{i,t-j}$, in each equation *i*. In a single-equation model, the probability of the LM test for first-order autocorrelation is one minus the quantile corresponding to the *t*-value of the estimated coefficient of $\hat{u}_{i,t-1}$, multiplied by 2 (because $H_0 : \hat{u}_{t-1} = 0$ involves a two-sided test). In an equation system the relevance of first-order autocorrelation is tested for by a χ^2-test of the null hypothesis that the $\hat{u}_{i,t-1}$ have a zero coefficient in each equation *i*. The third number in column 5 is the probability resulting from the Breusch-Godfrey test for autocorrelation of up to order 4 (see e.g. Godfrey 1988), where $\hat{u}_{i,t-1}, \hat{u}_{i,t-2}, \hat{u}_{i,t-3}, \hat{u}_{i,t-4}$ are inserted in each equation of the auxiliary regression. The first number in column 5 represents the probability resulting from the Ljung-Box (LB-) test for autocorrelation of up to order $(T/4)$. The Ljung-Box statistic (Ljung and Box 1978) is calculated according to

$$[21]\quad LB(T/4) = T(T+2)\sum_{j=1}^{T/4} \frac{r_j}{T-j}, \text{ where } r_j = \frac{\sum_{t=j+1}^{T} \hat{u}_t \hat{u}_{t-j}}{\sum_{t=1}^{T} \hat{u}_t^2},$$

and is $\chi^2[p(T/4-k)]$-distributed, with k being the lag length in an autoregressive model. The null hypothesis of this test is that all the $\hat{u}_{i,t-j}$ are insignificant in each equation i.

The rather low probability of the LB test (0.09) for $l=2$ in the model for the United States is probably due to the significance of $\hat{u}_{i,t-1}$. Increasing the lag length from 2 to 3 seems to resolve the problem of first-order autocorrelation but introduces some autocorrelation beyond order 4, which becomes intolerable as l is further increased. Moreover, there is some evidence for heteroscedasticity if l exceeds 3 and the information criteria decrease further, as mentioned above. I decide to choose $l=3$ thus ensuring absence of first-order autocorrelation at the cost of some additional higher-order autocorrelation. As a general observation applying to the VECMs for each country, residuals behave better and better as more and more identifying restrictions, testable over-identifying restrictions and (very few) arbitrary restrictions are imposed on the models. Specifically, one of the first steps in restricting the model, setting f exogenous in the VECM (giving it the form $n=6, p=5$), attenuates the problem of higher-order autocorrelation. For $l=3$, the probability of the LB-test rises from 0.03 to 0.19 solely by setting f exogenous.

Homoscedasticity in the residuals is tested for with the help of linear autoregressive conditional heteroscedasticity model (ARCH model) proposed by Engle (1982). To see whether the OLS assumption of a constant residual variance holds, the squared estimated residual in period t is regressed on its own past observations and on a constant. Here the lag length of the ARCH model is chosen to be l, the same as in the VECM. For each equation in the latter, a univariate LM-test (called ARCH test) of the form

[22] $$H_0 : \gamma_1 = \gamma_2 = ...\gamma_l = 0$$

is carried out, where the γ_i are the coefficients of the regression just described. The R^2 of this regression enters the test statistic which is computed as TR^2 and which is asymptotically $\chi^2(l)$-distributed under the null. The last column in Table 12 shows rounded values of the test statistics for each of the six equations. As one can see, the null of no ARCH effects cannot be rejected at the 10 percent significance level in any equation for lag lengths 2 and 3. At higher lag lengths, some evidence for conditional heteroscedasticity is found in the equations for Δq_{NCU} and Δf. This further backs the choice of $l = 3$ rather than a higher lag length.

In the Canadian case, the problem of autocorrelation is more pervasive. The results for $p = 6$ are shown in Table 13. The VECM contains two impulse dummies: $d771$ takes the value of 1 in 77:1 and the value of zero in all other quarters. It is used to address a strong outlier in equation Δf and will be dropped later, when f is set exogenous. The dummy $d824$ serves to meet outliers in the export volume and some other equations. Again, the information criteria (HQC and SC) favor the shortest reasonable lag length ($l = 2$). First-order and fourth-order autocorrelation is not so much of a problem at the 5 percent significance level, except for $l = 3$. However, there is higher-order autocorrelation as indicated by the LB-test which consistently rejects the hypothesis of no autocorrelation. Neither a huge number of attempts to vary the set of impulse dummies nor an increase of the lag length succeed in solving this problem. But it turns out that just as in the U.S. case, part of this problem is due to f being endogenous. When f is exogenous, the probability of the LB-test for $l = 2$ increases to 0.04 and additionally dropping $d771$ has it rise to 0.07. For higher lag lengths, the LB-test results remain less encouraging. As the ARCH-test indicates no heteroscedasticity at the 5 percent level (although the decision is tight in the equation for the change in U.S. industrial production, Δy^*), a lag length of 2 is considered to be the best choice for the VECM of Canadian aggregate exports.

Table 13: Determination of the Lag Length of the VECM: Canada[a]

Lag length l	$\ln\lvert\hat{\Omega}\rvert$ [b]	HQC[c]	SC[d]	Auto-correlation[e]	ARCH (l)[f]
2	–59.22	–56.45	–55.04	[0.00], [0.35], [0.08]	1.1, 1.8, 5.8*, 2.6, 0.5, 1.9
3	–59.61	–55.73	–53.76	[0.00], [0.00], [0.06]	2.6, 1.7, 8.0**, 2.8, 2.6, 2.5
4	–60.32	–56.32	–52.80	[0.00], [0.20], [0.08]	2.7, 4.1, 4.9, 9.2*, 4.0, 2.9
5	–61.02	–54.92	–51.83	[0.00], [0.15], [0.31]	5.1, 2.1, 1.2, 5.9, 8.2, 3.6

* (**, ***) means rejection at the 10% (5%, 1%) significance level. – [a]All variables are treated as being endogenous (p=6). The effective sample is 76:2–00:3. Impulse dummies are needed in the following quarters: 77:1, 82:4. Probabilities are in square brackets. – [b]Log of determinant of the estimated variance-covariance matrix. – [c]Hannan-Quinn information criterion. – [d]Schwarz information criterion. – [e]Probabilities of the following tests for autocorrelation: Ljung-Box test of autocorrelation up to order 25 ($= T/4$), Breusch-Godfrey LM-tests for first- and fourth-order autocorrelation, respectively. – [f]The six values indicate the ARCH-statistic of the test for autoregressive conditional heteroscedasticity of up to order l in each of the six equations $\{\Delta x, \Delta q_{NCU}, \Delta y^*, \Delta p^*_{NCU}, \Delta p_{NCU}, \Delta f\}$. The statistic is χ^2-distributed with l degrees of freedom. The 10 % critical values for l=2, …, 5 are 4.61; 6.25; 7.78; 9.24, respectively.

Finally, the lag length for Germany is chosen as $l = 2$ (see Table 14) along similar lines as those outlined for the United States and for Canada. Problems of heteroskedaticity as well as residual autocorrelation of order one and four are best avoided by setting $l = 2$, and even the LB-test does not support a higher lag length to meet the quite persistent problem of higher-order autocorrelation. Unlike the Canadian case, the higher-order autocorrelation does not stem from equation Δf alone but also from equation Δp^*_{NCU}. This is probably not due to the presence of exchange rates in this variable but rather comes from foreign price series themselves because for any lag length, autocorrelation in the foreign price equation is much worse when all prices in the system are denominated in a trade-weighted basket of foreign currencies. This is one reason why in the German VECM, too, all prices are denominated in units of the national currency.[42] Foreign prices will turn out to be exogenous when hypotheses on the vector of loading coefficients are tested so that autocorrelation of higher order will not be detected any more in the final long-run specification of Germany's VECM.

Four impulse dummy variables are needed to address strong outliers in the export equation and some of them are also significant in other equations. The dummies can well be justified by German economic history in which reunification was the most important event. After inner-German borders had broken down, the introduction of the D-Mark in Eastern Germany led to an export boom from West to East Germany. These exports were incorporated in x until the end of 1990 but were not reflected by the proxy of foreign economic activity, y^*. As flows of goods and services from West to East Germany are ignored by x beginning in 91:1 with the transition from West German to "all" German statistics, the level of x falls back to about pre-unification-boom levels given a virtually absent export basis in East Germany. This is why there is no need to account for a break in the long-run levels of the data e.g. by a step dummy, which has a pleasant side-effect: As only impulse dummies are used in all models, the asymptotic distributions of the various tests presented below are unaffected and the unknown critical values for models with dummies come quite close to those tabled in the literature for cases without dummy (e.g. Osterwald-Lenum 1992, Hansen and Juselius 1995, MacKinnon et al. 1999) for a sufficiently high number of observations. A strong negative outlier in export volumes occurred in 84:2 when hundreds of thousands of metalworkers were on strike. Finally, the second oil price shock caused the residual variance to grow larger in a number of equations ($\Delta x, \Delta q_{NCU}, \Delta p_{NCU}$), and setting an impulse dummy for quarter 80:1 seems to be the most parsimonious way to avoid residual heteroscedasticity.

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42 Another reason is that virtually all testable restrictions presented below are more compatible with the data when prices are in units of the national currency.

Table 14: Determination of the Lag Length of the VECM: Germany[a]

Lag length l	$\ln\lvert\hat{\Omega}\rvert$ [b]	HQC[c]	SC[d]	Auto-correlation[e]	ARCH (l)[f]
2	–64.30	–61.11	–59.49	[0.01], [0.39], [0.41]	2.0, 1.5, 3.3, 1.3, 3.1, 1.7
3	–64.97	–60.64	–58.46	[0.00], [0.17], [0.50]	2.1, 2.4, 8.1**, 7.4*, 6.3*, 0.6
4	–65.68	–60.23	–57.48	[0.00], [0.10], [0.37]	2.4, 2.9, 18.5***, 6.5, 4.8, 9.3*
5	–66.52	–59.93	–56.61	[0.00], [0.81], [0.01]	10.0*, 1.8, 20.0***, 5.3, 1.5, 2.1

* (**, ***) means rejection at the 10% (5%, 1%) significance level. – [a]All variables are treated as being endogenous (p=6). The effective sample is 76:2–00:3. Impulse dummies are needed in the following quarters: 80:1, 84:2, 90:3, 90:4. Probabilities are in square brackets. – [b]Log of determinant of the estimated variance-covariance matrix. – [c]Hannan-Quinn information criterion. – [d]Schwarz information criterion. – [e]Probabilities of the following tests for autocorrelation: Ljung-Box test of autocorrelation up to order 25 ($=T/4$), Breusch-Godfrey LM-tests for first- and fourth-order autocorrelation, respectively. – [f]The six values indicate the ARCH-statistic of the test for autoregressive conditional heteroscedasticity of up to order l in each of the six equations $\{\Delta x, \Delta q_{NCU}, \Delta y^*, \Delta p^*_{NCU}, \Delta p_{NCU}, \Delta f\}$. The statistic is χ^2-distributed with l degrees of freedom. The 10 % critical values for l=2, …, 5 are 4.61; 6.25; 7.78; 9.24, respectively.

5.3. The Number of Cointegration Relationships in the System

In this section the intuition behind the two different tests for cointegration rank presented in Johansen (1988) is briefly outlined, then the results from the application of the rank tests to my export models for the United States, Canada and Germany are presented. Last but not least, the roots of the companion matrix are analyzed in order to obtain further information on the number of cointegrating vectors, and a final decision is taken for each of the three countries in the sample.

5.3.1 Johansen's Trace and λ_{max} – Tests

To see whether the levels of variables are cointegrated, it makes sense to "split" the VECM in equation [13b] into two separate regressions. The first one is a VAR in first differences regressing Δz_t on the lagged first differences, the deterministic terms (and the dummy variables) producing the $(n \times 1)$-vector of residuals, R_{0t}, in each period, the second one regresses the levels z_{t-k} on the same regressors producing the $(n \times 1)$-vector of residuals, R_{kt}. If there were no cointegration relations at all, the second regression would not add any explanatory power to the VAR in first differences but if at least some of the I(1)-variables cointegrate, the reduced rank regression

[23] $$R_{0t} = \Pi\, R_{kt} + error = \alpha\beta' R_{kt} + error$$

contains a non-zero coefficient matrix. As already alluded to above, the cointegration matrix has been split into the $(p \times r)$-matrix α and the $(p \times r)$-matrix β, with $r \leq n$ being the cointegration rank, i.e. the number of cointegrating vectors in the system. This splitting is meaningful because the first r rows of β (i.e. the r cointegrating vectors, with r still to be determined) correspond to the eigenvectors of S_{kk}, the product moment matrix from the residuals R_{kt}. Notation for the product moment matrices is

[24] $$S_{ij} = T^{-1} \sum_{t=1}^{T} R_{it} R'_{jt},\ i, j = 0, k.$$

Conceptually, the solution can be divided into two steps. In the first step, estimates for α are found for given β with $\hat{\beta}$ defined by the ordinary-least-squares (OLS) result $\hat{\beta} = (X' X)^{-1} X' y$:

[25] $$\hat{\alpha}(\beta) = S_{0k}\beta(\beta' S_{kk}\beta)^{-1}.$$

In the second step, β is found by solving the eigenvalue problem

[26] $\left|\lambda S_{kk} - S_{k0}S_{00}^{-1}S_{0k}\right| = 0$

There are p eigenvalues which can be put in a decreasing order, $1 > \hat{\lambda}_1 > \hat{\lambda}_2 > \ldots > \hat{\lambda}_p > 0$. The maximum likelihood estimates for β are just the first r rows of the $(p \times p)$-matrix of the normalized eigenvectors corresponding to these eigenvalues. Intuitively speaking, r is kind of a frontier that separates $\hat{\lambda}_i$-values close to zero (the last $(n-r)$ ones) from those clearly above zero. Empirically, the near-zero eigenvalues are matched by $(n-r)$ near-zero columns in the matrix of loading coefficients, α. As a consequence, the $(n-r)$ last eigenvectors do not matter for the long-run relationships as they are multiplied by zeros. What matters are the r cointegrating vectors. The value of the maximized likelihood function subject to the rank r amounts to

[27] $L_{\max}^{-2/T} = |S_{00}| \prod_{i=1}^{r} \left(1 - \hat{\lambda}_i\right)$

As one can see, the presence of huge $\hat{\lambda}_i$-values makes the right-hand side of [27] become small and $L_{\max}$ grow large. Both statistical rank tests suggested by Johansen (1988) and applied here are based upon these maximized likelihood values. The first one is the so-called trace test based on the (positive) trace statistic *trs* defined as

[28] $trs(r) = -T \sum_{i=r+1}^{p} \ln\left(1 - \hat{\lambda}_i\right).$

In this test the null hypothesis is $H_0 : \hat{\lambda}_{r+1} = \hat{\lambda}_{r+2} = \ldots = \hat{\lambda}_p = 0$, the alternative hypothesis $H_1 : \hat{\lambda}_{r+1} \neq 0$. Finding the correct number of cointegrating vectors thus implies a sequence of tests and one plausibly starts with the hypothesis that is "easiest" to reject, i.e. with $r = 0$. If the null is rejected, H_1 is considered to hold and prevails as a pre-condition in the next test, the one of the hypothesis that all but the largest $\hat{\lambda}_i$-values equal zero. If one rejects again, the test is carried out for all but the two largest $\hat{\lambda}_i$-values, and so on. The sequence stops when the null cannot be rejected for the first time, e.g. in the test involving all but the three largest $\hat{\lambda}_i$-values. Then the test result is $r = 3$. The null is rejected for too high values of the statistic, which occurs when at least one of the eigenvalues in H_0 is huge, as one can see both in [27] and in [28]. The asymptotic distributions and corresponding critical values are derived in Johansen (1988, 1991, 1994) for various cases that differ from one another with respect to the assumptions on the deterministic components. In general, critical values are the higher the more deterministic elements are in the equation, and they are still higher if the constant or the trend are restricted to the cointegration space because keeping them out of the short-run model involves one additional restriction.

The second rank test is the so-called $\lambda_{\max}$-test with the $\lambda_{\max}$-statistic defined as

[29] $\lambda_{\max}(r) = -T \ln\left(1 - \hat{\lambda}_{r+1}\right).$

As the statistic suggests, it is tested here whether $\hat{\lambda}_{r+1}$, the next smaller neighbour of $\hat{\lambda}_r$, equals zero given that $\hat{\lambda}_r$ does not. Again, it is advisable to start with $r = 0$ in order to check first whether the biggest eigenvalue is significant and thus if there is any cointegration in the VECM at all. The sequence stops when rejection fails for the first time. Critical values are also derived in Johansen (1988, 1991, 1994).

5.3.2 Results from Johansen's Rank Tests

I now turn to the results of the rank tests. For the United States, both the trace test and the $\lambda_{\max}$-test successively reject the null hypothesis of zero, then at most one, then at most two cointegrating vectors.[43] The first null that cannot be

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[43] The $\lambda_{\max}$-test rejects the null hypothesis of at most two cointegrating vectors only at the 10 percent, not at the 5 percent significance level.

rejected at the 10 percent level is $r \leq 3$ (Table 15). However, it has been noted in the literature that the Johansen procedure over-rejects the null in small samples (Reimers 1992) and that the finite-sample bias is a positive function of $T/(T-pl)$ (Cheung and Lai 1993). To address this problem, I adjust the trace statistics in the way suggested by Reimers and also document the corrected trace values. According to the latter, $r \leq 2$ is the first null that is not rejected. Therefore, I consider there to be two cointegrating vectors in the VECM.

Table 15: Determination of Cointegration Rank: United States

Rank r [a]	$p-r$ [a]	Eigen-value	Trace statistic	Cor-rected trace[b]	Critical value trace OL92[c]	Critical value trace MK99[d]	λ_{max} -statistic	Critical value λ_{max} OL92[c]	Critical value λ_{max} MK99[d]
0	6	0.425	146.51*	120.14*	89.37	92.61	55.27*	36.76	38.15
1	5	0.317	91.23*	74.81*	64.74	67.36	38.14*	30.90	31.39
2	4	0.229	53.09*	43.53	43.84	46.18	26.03*	24.73	25.34
3	3	0.147	27.06	22.19	26.70	29.03	15.87	18.60	19.25
4	2	0.088	11.20	9.18	13.31	15.87	9.24	12.07	13.05
5	1	0.019	1.96	1.61	2.71	6.59	1.96	2.69	6.59

* denotes rejection at the 10% significance level. – [a]r is the number of cointegrating vectors, p the number of enogenous variables. For any given r, the null hypothesis of the trace test is that the $p-r$ smallest eigenvalues of the cointegration matrix Π equal zero against the alternative that all eigenvalues are different from zero ($r = n$, i.e. matrix Π has full rank). The λ_{max}-test tests the null of r cointegrating vectors ($(r+1)th$ eigenvalue equals zero) against the alternative of $(r+1)$ cointegrating vectors (the $(r+2)th$ and all further eigenvalues equal zero). – [b]Equals $[(T-pl)/T]$ times the trace statistic according to the correction proposed by Reimers' (1992). The adjustment factor for the U.S. VECM is $(100-6\cdot 3)/100 = 0.82$. – [c]10 percent critical value for VECM with unrestricted constant taken from Osterwald-Lenum (1992: 468). – [d]10 percent critical value for VECM with unrestricted constant derived from Monte Carlo simulations with larger sample size by MacKinnon et al. (1999).

In case of the Canadian VECM, just as in the U.S. example, the finding of $r = 2$ is not straightforward, the test results are even more contradictory. For instance, the uncorrected trace test points to $r = 3$, although the null of at most two cointegration relationships is only rejected at the 10 percent, not at the 5 percent significance level (critical value according to MacKinnon et al. (1999) is 49.64). In turn, the $\lambda_{\max}$-test would rather indicate $r = 1$. However, one should take into account that the assumption of normally distributed residuals is violated in the initial VECM for Canada.[44] As the trace test is more robust than the $\lambda_{\max}$-test with respect to both skewness and excess kurtosis in residuals (Cheung and Lai 1993), the trace test is given more weigth in this case. The Reimers-corrected trace statistic for $r = 2$ amounts to 67.03, a value well above the 10 percent critical value given in Osterwald-Lenum (1992) and only a little smaller than the one in MacKinnon et al. (1999). Although Reimers' correction is a step in the right direction as the sign of the finite-sample bias in the Johansen procedure is not controversial at all, it is not clear that dividing the trace statistic by $T/(T - pl)$ is an optimal correction.[45] Therefore (and for reasons already mentioned, e.g. the presence of impulse dummies), the critical values should rather serve as an indication for the unknown number of cointegrating vectors than as a precise test. Given that at the one hand, $r = 0$ is rejected very clearly, but that on the other hand, the test statistics decay quite slowly relative to their critical values in the range from one to three vectors, it seems advisable to allow for more than one stationary linear combination in the system. Yet as the null hypothesis $r \leq 2$ is not rejected at the 5 percent level by any of the rank tests discussed, I consider that restricting the system to two cointegration relationships is the best choice.

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44 Some excess kurtosis prevails in equations Δy^* and Δf. Both U.S. industrial production and the trade intensity of world production will be set exogenous later in the analysis, and the problem of non-normality detected by the multivariate normality test will disappear.

45 Doubts on the optimality of Reimers' correction are raised by Doornik and Hendry (1997: 225).

Table 16: Determination of Cointegration Rank: Canada

Rank r [a]	$p-r$ [a]	Eigen-value	Trace statistic	Cor-rected trace[b]	Critical value trace OL92[c]	Critical value trace MK99[d]	λ_{max} -statistic	Critical value λ_{max} OL92[c]	Critical value λ_{max} MK99[d]
0	6	0.400	128.08*	113.02*	89.37	92.61	52.12*	36.76	38.15
1	5	0.229	75.96*	67.03(*)	64.74	67.36	26.47	30.90	31.39
2	4	0.192	49.49*	43.67	43.84	46.18	21.68	24.73	25.34
3	3	0.136	27.81	24.54	26.70	29.03	14.86	18.60	19.25
4	2	0.077	12.95	11.43	13.31	15.87	8.18	12.07	13.05
5	1	0.046	4.76	4.20	2.71	6.59	4.76	2.69	6.59

* denotes rejection at the 10% significance level. – [a] r is the number of cointegrating vectors, p the number of enogenous variables. For any given r, the null hypothesis of the trace test is that the $p-r$ smallest eigenvalues of the cointegration matrix Π equal zero against the alternative that all eigenvalues are different from zero ($r=n$, i.e. matrix Π has full rank). The λ_{max}-test tests the null of r cointegrating vectors ($(r+1)th$ eigenvalue equals zero) against the alternative of $(r+1)$ cointegrating vectors (the $(r+2)th$ and all further eigenvalues equal zero). – [b]Equals $[(T-pl)/T]$ times the trace statistic according to the correction proposed by Reimers' (1992). The adjustment factor for the Canadian VECM is $(102-6\cdot 2)/100=0.8824$. – [c]10 percent critical value for VECM with unrestricted constant taken from Osterwald-Lenum (1992: 468). – [d]10 percent critical value for VECM with unrestricted constant derived from Monte Carlo simulations with larger sample size by MacKinnon et al. (1999).

How many vectors are found in the system of Germany's exports? Both the uncorrected trace statistic and the λ_{max}-statistic lead to rejection of $H_0: r \le 1$ and to non-rejection of $H_0: r \le 2$. Once again, the λ_{max}-statistic should not be paid too much attention due to non-normality in the distribution of residuals (that will get settled once f is set exogenous). Then a decision has to be taken between the competing outcomes "two cointegrating vectors" according to the uncorrected trace statistic and "one cointegrating vector" according to the corrected trace statistic. Using the "margin of judgment" in the same direction as I did for the other two countries, i.e. arguing in favour of a rather small number of cointegration relationships, would clearly oblige me to decide $r=1$.

Nevertheless, I decide to again restrict the VECM to $r = 2$ on the ground that the 10 percent critical values of the trace test lie roughly in the middle of the interval spanned by the corrected and the uncorrected trace statistics. An additional, more pragmatic argument is to support a maximum of international comparability in my specification choices as long as this is compatible with the data.

Table 17: Determination of Cointegration Rank: Germany

Rank r [a]	$p-r$ [a]	Eigen-value	Trace statistic	Cor-rected trace[b]	Critical value trace OL92[c]	Critical value trace MK99[d]	λ_{max} -statistic	Critical value λ_{max} OL92[c]	Critical value λ_{max} MK99[d]
0	6	0.459	132.45*	116.56*	89.37	92.61	61.48*	36.76	38.15
1	5	0.273	70.97*	62.45	64.74	67.36	31.82*	30.90	31.39
2	4	0.178	39.15	34.45	43.84	46.18	19.55	24.73	25.34
3	3	0.127	19.60	17.25	26.70	29.03	13.61	18.60	19.25
4	2	0.058	5.99	5.27	13.31	15.87	5.96	12.07	13.05
5	1	0.000	0.03	0.03	2.71	6.59	0.03	2.69	6.59

* denotes rejection at the 10% significance level. – [a]r is the number of cointegrating vectors, p the number of enogenous variables. For any given r, the null hypothesis of the trace test is that the $p-r$ smallest eigenvalues of the cointegration matrix Π equal zero against the alternative that all eigenvalues are different from zero ($r = n$, i.e. matrix Π has full rank). The λ_{max}-test tests the null of r cointegrating vectors ($(r+1)th$ eigenvalue equals zero) against the alternative of $(r+1)$ cointegrating vectors (the $(r+2)th$ and all further eigenvalues equal zero). – [b]Equals $[(T-pl)/T]$ times the trace statistic according to the correction proposed by Reimers' (1992). The adjustment factor for the German VECM is $(100-6\cdot 2)/100 = 0.88$. – [c]10 percent critical value for VECM with unrestricted constant taken from Osterwald-Lenum (1992: 468). – [d]10 percent critical value for VECM with unrestricted constant derived from Monte Carlo simulations with larger sample size by MacKinnon et al. (1999).

5.3.3. The Roots of the Companion Matrix

As the results from the different rank tests are not very clear-cut, with the models for the United States and for Canada allowing for two or three cointegrating vectors, the German model for one or two vectors, one should use additional information concerning the most plausible assumption on the number of stationary linear combinations in the three VECMs. Such additional information can be extracted from the eigenvalues (i.e., roots) of the companion matrix, A. This matrix comprises all the $(n \times n)$ parameter matrices $A_i, i = 1,...,l$, from [13a]. It has the form

$$[30] \quad A = \begin{bmatrix} A_1 & A_2 & \dots & A_{l-1} & A_l \\ I_n & 0 & \dots & 0 & 0 \\ 0 & I_n & \dots & 0 & 0 \\ \vdots & & \ddots & & \vdots \\ 0 & 0 & & I_n & 0 \end{bmatrix},$$

where I_n is the n-dimensional identity matrix which is of dimension $(nk \times nk)$. Thus A is (12×12) for Germany and Canada and (18×18) for the United States as $n = p = 6$ and $k = 2$ and $k = 3$, respectively. As a consequence the companion matrix has 12 and 18 eigenvalues, respectively. The look on these eigenvalues allows to check whether the VAR model in levels [13a] converges in the long run. The number of roots close to the unit circle corresponds to the number of I(1)-processes within Π and thus equals $(n - r)$ (Juselius 1993: 14). Although no statistical test is carried out to answer the question whether a root is or is not significantly below one, the analysis of on the eigenvalues serves as a "cross-check" to the formal rank test results.

An additional advantage of the analysis is that it may detect explosive processes, which indicate that the model chosen is inadequate. This is the case when one or more real roots of A exceed one. Then one or more variables might in fact be integrated of order two. In the present case ($p = 6$) the roots near one are 1.0084, 0.9575, 0.9575, 0.9356, and 0.9356 for the United States; 1.0120, 0.9543, and 0.9543 for Canada; 1.0034, 0.9756, 0.9217, and 9217 for Germany. This points to 1, 3, and 2 cointegrating vectors, respectively. This result is probably due to the near-I(2) behaviour of the globalization variable. As the other variables are still assumed to be I(1), the I(2)-analysis suggested by Juselius (1993) in presence of several I(2) variables would not be most appropriate here.

Rather, the therapy chosen consists of setting f exogenous to the VECM. This appears to be fully justified on economic grounds as one can assume that neither of the three countries may determine long-run trends in world trade and production by its mere export volumes.[46] The statistical evidence for f being exogenous is mixed, however. Under the restriction of two cointegration relationships, the likelihood ratio statistic of the null hypothesis that both loading coefficients in equation Δf equal zero is 9.06 for the United States, 18.32 for Canada, and 0.37 for Germany. The corresponding probabilities drawn from the χ^2-distribution with two degrees of freedom are 0.83, 0.01 and 0.00, respectively. Non-convergence in case of an endogenous f and thus model instability are the most compelling reasons for setting the trade intensity of world GDP exogenous but not the only ones. It also makes sense from an economic point of view as it is not very plausible that the export performance of a single nation, even of a big one, has repercussions on trends in world trade in the long run. Among all six variables of the VECM, the decision of setting one of them exogenous purely on economic grounds seems most justified for f. This is noteworthy because the decision comes at the price of some statistical arbitrariness.

As one can see, all roots are inside the unit circle when the model is modified in this way (Table 18). Moreover, the case for $r = 2$ becomes stronger for the United States and Germany as there seem to be three I(1)-processes in the system of five endogenous I(1)-variables.

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46 Remember that in case of the United States, f does not contain domestic exports and GDP.

Table 18: The Seven Biggest Roots of the Companion Matrix in each VECM with Exogenous Globalization Variable ($p = 5$)[a]

United States	Canada	Germany
0.9981	0.9565	0.9819
0.9981	0.9565	0.8802
0.9381	0.7904	0.8802
0.8530	0.7340	0.7850
0.8530	0.7340	0.7850
0.5707	0.5898	0.7063
0.5707	0.5898	0.4084

[a]The number of roots close to one is an indication of $(n-r)$, the number of non-stationary relations in matrix Π. These numbers most probably are 3, 2 and 3 for the United States, Canada, and Germany, respectively.

5.3.4. Conclusions for the Number of Cointegration Relationships

To sum up the discussion on the cointegration rank in my models of aggregate exports, Table 19 gives an overview of the results. A unique number of cointegrating vectors does not result from any of the three models under investigation but the conclusion $r = 2$ is clearly the most plausible one for the United States and Germany as not only the eigenvalues of the companion matrix but also the trace statistic and the adjusted trace statistic, respectively, support this result. The rank of Π for Canada seems to be a limit case between $r = 2$ and $r = 3$ and would merit deeper investigation. But as it is not possible to achieve results any better than those presented below from the perspective of structural long-run analysis[47], I stick to the decision $r = 2$ to ensure international comparability of my results. Anyway, among the two errors of

47 The fact that the "rest of the world" is limited to the United States in the Canadian model might play a role in explaining why the rank test results diverge from those for other countries. Many attempts to identify structural relationship for Canada under the restriction of $r = 3$ have been undertaken and are available from the author upon request. Besides the long-run export supply and demand relationships found for all three countries, natural "candidates" for further relationships include a (trend-)stationary real effective exchange rate of the Canadian dollar and some form of purchasing power parity between the United States and Canada.

choosing r too small by one or of choosing it too high by one, the latter is probably more problematic as the cointegration space would then contain "one non-stationary relation […] thus invalidating the stationary inference" (Juselius (1993: 14)). The rank analysis finished, the task of the next section is to implement the view that the two stationary long-run relationships can be interpreted as export demand and export supply. To achieve this, the identification problem has to be solved.

Table 19: Cointegration Ranks According to Different Tests and Indicators: United States, Canada, and Germany[a]

Indicator	United States	Canada	Germany
Uncorrected Trace Statistic[b]	3	3	2
Corrected Trace Statistic[b]	2	2	1
$\lambda_{\max}$-Statistic[b]	3	1	2
Roots of Companion Matrix ($p = 6$)[c,d]	1	3	2
Roots of Companion Matrix ($p = 5$)[c,e]	2	3	2
Decision	**2**	**2**	**2**

[a]The numbers in the table correspond to the rank of the cointegration matrix, Π. – [b]The corresponding test statistics and critical values are reported in Tables 15, 16, and 17. – [c]Number of roots near the unit circle (see text for description). – [d]All I(1) variables are taken as endogenous. – [e]The globalization variable f is set exogenous to ensure that all roots of the companion matrix be inside the unit circle. The other five I(1) variables remain endogenous.

5.4. Identifying the Long-Run Relationships: Problem and Strategy

The Johansen procedure to determine the number of cointegrating relationships only informs on *how many unique* cointegrating vectors span the cointegration space but does not necessarily deliver unique estimates of these vectors in the reduced-rank regression. As any linear combination of I(0) variables is itself I(0), one of my two vectors might in fact be a linear combination of the two unique structural vectors (Johansen and Juselius 1992: 222).[48] In order to eliminate this possibility and to extract the two structural relationships "hidden" in the cointegration space, I need additional assumptions, so-called identifying restrictions. In this chapter I argue that it is sound to interpret the two cointegration relationships as structural long-run relationships of export demand and supply, and the restrictions will be set accordingly.

5.4.1. The Identification Problem Applied to the Example of Aggregate Exports

The interaction of two distinct groups of economic agents, demanders and suppliers, produces a pair of values $(x_t, q_{NCU,t})$ (labeled $y_{1,t}, y_{2,t}$ in the following) which might be represented in a price-quantity diagram. One dot corresponds to each period and results from the optimizing behavior of demand and supply given the same period's realizations in the other variables ($y^*, p^*_{NCU}, p_{NCU}, f$, labelled $x_1,...,x_4$ in the following).[49] Changes in these other four variables shift the structural demand and supply curves around. The problem of identification arises because it is impossible to "recognize" two distinct curves (demand and supply) in the data cloud without further assumptions when one knows the realizations of the predetermined variables. This becomes clear when one tries to express the structural coefficients in terms of the parameters of the reduced form as is done in the following. Labeling the first cointegration relationship "demand" and the second one "supply", the structural long-run system of export quantities and prices is written as

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48 Instead of the interesting structural long-run parameters β and loading coefficients α, what one might see on the regression output is $\alpha^* \beta'^* = (\alpha\xi^{-1})(\xi\beta')$, with ξ being any $r \times r$ non-singular matrix (Harris 1995: 95).

49 This contingency does not mean that the levels of foreign production, foreign and domestic prices as well as the globalization proxy are exogenous to the whole VECM. Rather, they are considered as being predetermined in the *reading* of the cointegration relationships as export demand and supply. Whether deviations from the long-run relationships have repercussions on these four variables remains to be checked (see tests for weak exogeneity below), and any feedback will be duly taken into account.

$$[31]\quad \begin{aligned} \beta_{11}y_{1t} + \beta_{12}y_{2t} + \gamma_{11}x_{1t} + \gamma_{12}x_{2t} + \gamma_{13}x_{3t} + \gamma_{14}x_{4t} = u_{1t} \\ \beta_{21}y_{1t} + \beta_{22}y_{2t} + \gamma_{21}x_{1t} + \gamma_{22}x_{2t} + \gamma_{23}x_{3t} + \gamma_{24}x_{4t} = u_{2t} \end{aligned},$$

or, more compactly, as

[32] $By_t + Cx_t = u_t$.

B is the (2×2)-matrix of slope coefficients for export quantities and prices, C is the (2×4)-matrix of the predetermined variables, which shift the demand and supply curves. u_{1t} and u_{2t} are structural residuals which can be interpreted as the "clean disequilibrium" (Hansen and Juselius 1995: 22), i.e. the deviation of export volumes and prices from their respective equilibrium levels. The reduced form reads

[33] $y_t = Rx_t + v_t$, with $R = -B^{-1}C$ and $v_t = B^{-1}u_t$.

The solution of the reduced form in terms of the structural coefficients is:

[34]

$$y_t = -1/\Delta \begin{bmatrix} (\gamma_{11}-\beta_{12}\gamma_{21})(\gamma_{12}-\beta_{12}\gamma_{22})(\gamma_{13}-\beta_{12}\gamma_{23}) & (\gamma_{14}-\beta_{12}\gamma_{24}) \\ (-\beta_{21}\gamma_{11}+\gamma_{21})(-\beta_{21}\gamma_{12}+\gamma_{22})(-\beta_{21}\gamma_{13}+\gamma_{23}) & (-\beta_{21}\gamma_{14}+\gamma_{24}) \end{bmatrix} [x_{1t}x_{2t}x_{3t}x_{4t}]'$$
$$+1/\Delta \begin{bmatrix} \beta_{22} - \beta_{12} \\ -\beta_{21}\beta_{11} \end{bmatrix} \begin{bmatrix} u_{1t} \\ u_{2t} \end{bmatrix}, \text{ where } \Delta = \beta_{11}\beta_{22} - \beta_{12}\beta_{21}$$

As R is of order (2×4), estimation of the reduced form yields the 8 coefficients

$$[35]\quad R = \begin{bmatrix} r_{11} & r_{12} & r_{13} & r_{14} \\ r_{21} & r_{22} & r_{23} & r_{24} \end{bmatrix}.$$

The problem is that B and C together contain 12 structural parameters. Two of the four missing parameters are obtained by normalization, e.g. $\beta_{11} = 1, \beta_{22} = 1$, which allows to express the equilibrium level of exports conditional on the five other variables in equation 1 and the equilibrium level of export prices conditional on the five other variables in equation 2. There are two structural parameters left which are not determined by either reduced-form coefficients or normalization, and the other structural variables (apart from $\beta_{11}=1, \beta_{22}=1$) can only be expressed as a function of these two "missing" parameters. Thus the solution to the system is not unique but situated on a plane in the R^2-space, the system is under-determined or "not identified". To get a unique solution, two restrictions on structural parameters are required. Intuitively speaking, one can only be sure about the slope of the demand and supply curves after specifying one shift parameter in the demand and one in the supply relationship by assumption.[50] For computational purposes, any such two restrictions would lead to a unique solution (as long as they are not linear combinations of one another), but a sound analysis would always try to derive such restrictions from economic theory. In general, zero restrictions (a variable does not appear in a structural relationship), homogeneity restrictions (two variables have the same coefficients, e.g. restriction of the production-cost elasticity of export prices to one) or add-up restrictions (two long-run elasticities add up to one or to the parameter of a third variable) are used.

The solution of the problem is first presented in a "technical" manner, then the identification restrictions chosen will be motivated by economic reasoning. The formal requirement is reflected by the rank condition of identifiability (Johnston and DiNardo 1997: 310–312). From [33] it follows that

[36] $BR + C = 0$ or $AW = 0$, where $A = [B \quad C], W' = [R \quad I_k]$.

Let G be the number of endogenous variables in the system (here: 2), g the number of endogenous variables in an equation (here: 2 in each), K the number of predetermined variables in the system (here: 4), k the number of predetermined variables in the an equation. In the first equation, the four reduced-form

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50 Identification would also be yielded by restricting, say, the long-run price elasticity of export supply ($-\beta_{22}/\beta_{21}$), but this does not make much sense in the present context as the aim is to estimate the price elasticities on both sides of the markets and to examine the widely used infinitely-elastic-supply assumption.

parameters and the normalization condition allow to determine five of the six structural coefficients. With a_{ij} designing the elements of matrix A, I have

[37] $a_1 W = 0$

The identifying restriction is set by adding a column vector Φ_1 to matrix W that also satisfies

[38] $a_1 \Phi_1 = 0$.

While the dimension of W is $(G+K)\times K$ (here: 6×4), the one of Φ_1 is $(G+K)\times 1$ as the number of columns corresponds to the number of identifying restrictions imposed on the first equation $(\#)$.[51] For example, a zero restriction on domestic production costs in the export demand relationship, $\gamma_{13}=0$, is implemented by setting $\Phi_1' = [0 \quad 0 \quad 0 \quad 0 \quad 1 \quad 0]$. Combining [37] and [38] yields the complete set of restrictions:

[39] $a_1 [W \quad \Phi_1] = 0$.

Matrix $[W \quad \Phi_1]$ has rank $(K{+}\#)$, which equals five in the example at hand. With $\beta_{11}=1$ by normalization, there are now five equations and the five structural coefficients can be determined uniquely. Thus the necessary condition of identification,

[40] $K{+}\# \geq G+K-1$,

is now fulfilled for the first cointegration relationship. The computation for the structural coefficients is done in terms of the reduced-form coefficients:

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51 In the following, # stands for “number of restrictions” in a broader sense. It will become clear from the context which restrictions this refers to.

[41] $a_1[W \quad \Phi_1] = [1 \quad \beta_{12} \quad \gamma_{11} \quad \gamma_{12} \quad \gamma_{13} \quad \gamma_{41}] \begin{bmatrix} r_{11} & r_{12} & r_{13} & r_{14} & 0 \\ r_{21} & r_{22} & r_{23} & r_{24} & 0 \\ 1 & 0 & 0 & 0 & 0 \\ 0 & 1 & 0 & 0 & 0 \\ 0 & 0 & 1 & 0 & 1 \\ 0 & 0 & 0 & 1 & 0 \end{bmatrix} = [0 \quad 0 \quad 0 \quad 0 \quad 0].$

The solution is:

[42] $\gamma_{13} = 0; \beta_{12} = -r_{13}/r_{23}; \gamma_{11} = r_{21}(r_{13}/r_{23}) - r_{11}; \gamma_{12} = r_{22}(r_{13}/r_{23}) - r_{12}; \gamma_{14} = r_{24}(r_{13}/r_{23}) - r_{14}$.

The identification of the second structural relationship works very much like the one of first relationship. Again there are six structural coefficients one of which is set by normalization (the coefficient of export prices: $\beta_{22} = 1$), but there are only four equations as W is of rank four. Now I restrict the coefficient of foreign industrial production to zero ($\gamma_{21} = 0$) in order to obtain a unique solution for the structural long-run supply parameters. The computational exercise resembles the one in [41] with a_1 replaced by $a_2 = [\beta_{21} \quad 1 \quad \gamma_{21} \quad \gamma_{22} \quad \gamma_{23} \quad \gamma_{24}]$ and Φ_1 by Φ_2, which reads $\Phi'_2 = [0 \quad 0 \quad 1 \quad 0 \quad 0 \quad 0]$. The solution in terms of reduced-form coefficients is

[43] $\gamma_{21} = 0; \beta_{21} = -r_{21}/r_{11}; \gamma_{22} = r_{12}(r_{21}/r_{11}) - r_{22}; \gamma_{23} = r_{13}(r_{21}/r_{11}) - r_{23}; \gamma_{24} = r_{14}(r_{21}/r_{11}) - r_{24}$.

As only zero restrictions have been used, the necessary condition of identifiability may also be expressed by the order condition which states that the number of predetermined variables excluded from the equation considered (*k*) must be at least as great as the number of endogenous variables in this equation (*g*) less one (Hansen 1993: 176).

A still easier check of the rank condition [40] being fulfilled is delivered by Farebrother (1971) who proves that [40] holds if and only if

[44] $rank(A\Phi_i) = G - 1, \ \forall i, i = 1,...,G,$

Passing by this criterion makes it unnecessary to deal with the reduced-form parameters at least at the stage of identification. Applied to the first and second structural equation in my example ($G = 2$), this condition reads

[45a] $rank(A\Phi_1) = rank\begin{bmatrix} \gamma_{13} \\ \gamma_{23} \end{bmatrix} = rank\begin{bmatrix} 0 \\ \gamma_{23} \end{bmatrix} = 1 = G - 1.$ [52]

[45b] $rank(A\Phi_2) = rank\begin{bmatrix} \gamma_{11} \\ \gamma_{21} \end{bmatrix} = rank\begin{bmatrix} \gamma_{11} \\ 0 \end{bmatrix} = 1 = G - 1.$

The latter version of the rank condition is particularly meaningful in the context of cointegration analysis. As one can see, multiplication of the second row in *A*, a_2, with Φ_1 produces a value different from zero as does multiplication of the first row in *A*, a_1, with Φ_2 such that the respective products $A\Phi_i$, *i*=1, 2, have the required rank. Moreover, the vectors $A\Phi_1$ and $A\Phi_2$ are not zero vectors or linear combinations from one another, which is important because it makes sure that a variable does not completely drop out of the cointegration space. A dropout would reduce the number of variables without solving the identification problem. In the context of cointegration analysis, linear economic hypotheses to be tested against the data may be put on each cointegrating vector β_i such that the $(p \times r)$-matrix of all cointegrating variables can be represented as

[46] $\beta = (H_1\psi_1, ..., H_r\psi_r)$,

where ψ_i is the vector of the freely estimated coefficients in column vector β_i. Then the first equation is identified if

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52 If three cointegration vectors had been found in the VECM, *G* would be 3 and two linearly independent restrictions would have been required for identification. $A\Phi_1$ would then be a (3×2)-matrix with a zero vector (rather than a scalar) in the first row and a square matrix of order $(G - 1)$ containing the coefficients of all those variables which have been excluded from the first equation. This square matrix is non-singular (and the first equation identified), in general.

[47] $rank(\Phi_1'\beta_1,...,\Phi_1'\beta_r) = rank(\Phi_1'H_1\psi_1,...,\Phi_1'H_r\psi_r) = r-1$.

As a consequence, "no linear combination of $\beta_2,...\beta_r$ can produce a vector that 'looks like' the coefficients of the first relation, that is, satisfies the restrictions defining the first relation." (Johansen and Juselius 1994: 15).

5.4.2. The Identifying Restriction for the Demand Vector

As already mentioned in the general discussion of the identification problem, the identifying restrictions used here to single out the long-run demand for and the long-run supply of aggregate exports are $\gamma_{13} = 0$ and $\gamma_{21} = 0$, respectively. These restrictions are well justified on economic grounds. The first one states that a change in domestic producer prices has no direct long-run effect on export demand. This makes sense because the price variable foreign customers care about is the index of export prices in foreign currency units. The two sources of change in this variable are reflected in the remaining two price variables of the demand relationship: q_{NCU} to reflect changes in the price of exports in domestic currency units and $p^*_{NCU} = p^* - w$ to reflect changes in the nominal effective exchange rate. The restriction is supported by the choice of the static model of export demand outlined in section 2 in which domestic producer prices do not appear. This does not mean that the quantity of exports demanded is completely independent from domestic prices or costs, first because there might be short-run effects, second because a change in the domestic producer price index probably acts on the export supply side and may induce a change in export prices[53] which then has an impact on demand. The exclusion of the direct long-run effect by restriction can be "translated" economically by stating that a change in the

[53] In the Goldstein/Khan (1978) model, an increase in domestic producer prices *ceteris paribus* has exporters shift part of their supply from foreign to domestic markets as the relative profitability of exporting (the ratio of export prices to domestic prices) deteriorates. In the Dornbusch (1987) model, an increase in domestic producer prices directly leads to higher export prices set by the exporting firm.

domestic producer price index may lead to a movement *along* the long-run export demand curve but does not shift this curve.

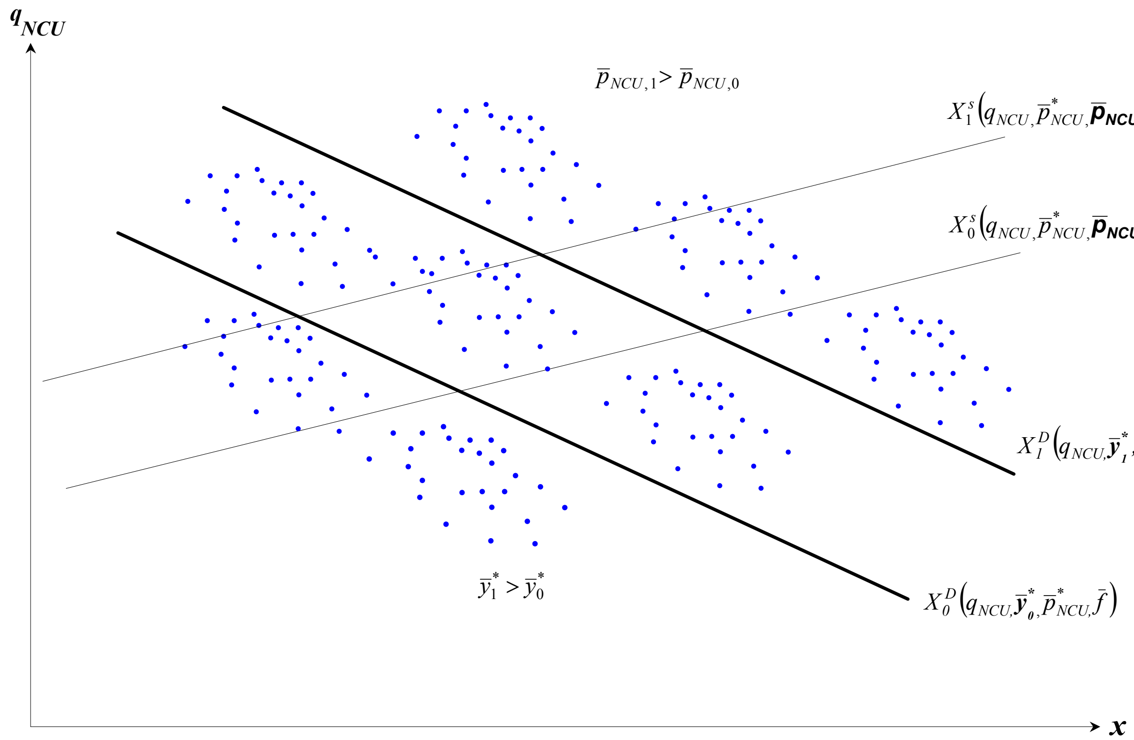
5.4.3. The Identifying Restriction for the Supply Vector

The second identifying restriction states that a change in industrial production abroad has no direct long-run effect on export supply. The economic rationale is that aggregate foreign production (or income) is the classical scale variable in all models of aggregate export demand and does not appear in most export supply models. This is particularly true for the Dornbusch (1987) model which puts the focus on price-setting behavior, not on quantities. But even in the model by Goldstein and Khan (1978) stressing that providing the market with more exports may involve higher production costs and thus export prices even in the long run, it is more reasonable to assume an indirect link between foreign production and export supply, not a direct one: Higher foreign production increases the demand for exports and only this change in demand leads to an adjustment by exporters. In other words, the second identifying restriction has it that a change in foreign economic activity leads to a shift in the long-run demand curve *along* the given supply curve but does not shift this supply curve.

The distinction between movements *on* the curve and *shifts* of the curve contributes to a more literal understanding of the notion of identification. Assuming for a moment that all disturbances to the system arise from changes in foreign output, then one concludes by virtue of the second identifying restriction that there exists a single long-run supply curve but as many demand curves as there are changes in output. Then each shift in the demand curve leads to a different long-run equilibrium in export price and quantity point, and many shifts produce many points which all lie on that single supply curve. To put it differently, changes in foreign production *"trace out"* the supply curve. Thanks to the second identifying restriction the researcher is able to discern the supply curve from an amorphous cloud of historical export price and quantity observations.

Likewise, if all disturbances were changes in domestic producer prices, the supply curve would shift a lot but all equilibria would be situated on a single demand curve, as is postulated by the first identifying restriction. Implementing the two long-run restrictions imposes a structure on the cointegration space as the supply curve and the demand curve emerge from the "fog" of historical data. This recognition process is illustrated in Figure *4*, where the long-run effect of a change in foreign industrial production (with $\bar{y}_1^* > \bar{y}_0^*$) and the one of a change in domestic producer prices (with $p_{NCU,1} > p_{NCU,0}$) are shown in a stylized manner: An increase in y^* shifts the demand curve outwards while leaving the supply curve unaffected, an increase in p_{NCU} shifts the supply curve inwards without affecting the demand curve, whereas changes in x and q_{NCU} trigger movements *on* the curve. The bars above symbols indicate shift variables.

Figure 4: The Intuition Behind Identification



5.5. Finding the Structural Long-Run Relations

To gauge the "usefulness" of the identification strategy just outlined, the identified relationships are searched for one by one before the exactly identified long-run model is presented at a glance. The first question in this sub-section is whether the restriction just identifying the long-run demand (supply) relationship holds in both vectors spanning the cointegration space. This allows to see whether one of the two cointegration relationships detected by the rank tests (or even both) is (are) in fact a linear combination of the structural demand and supply vectors. The assumption tested is more restrictive than the one made in the second type of tests, where I ask whether a vector can be detected in the cointegration space that contains the characteristics of the long-run demand (supply) relationship. In this second type of tests the other vector, presumably the supply (demand) vector, remains unrestricted.

The first question is answered with the $®_4$ type of hypothesis test outlined in Johansen and Juselius (1992: 225). Under the null hypothesis (called $®_4$), the unrestricted (6×2)-matrix of cointegrating vectors β (from the VECM with exogenous globalization) can be factorized into a $(n \times s)$ restriction matrix H_4 and a $(s \times r)$-matrix φ of freely estimated long-run coefficients,

[48] $®_4: \beta = H_4\varphi, \qquad H_4(n \times s), \varphi(s \times r), \qquad r \leq s \leq n,$

where $n = 6, r = 2$, and $s = n - \#$ is the number of unrestricted parameters per vector. The lines of matrix H_4 are in the order $(x, q_{NCU}, y^*, p^*_{NCU}, p_{NCU}, f)$. First the test is performed for the restriction identifying the demand relationship (superscript *D*). As this is a zero restriction on domestic prices, imposing it also on the second vector is equivalent to dropping p_{NCU} from the cointegration space.[54] That drop is reflected by the zero row vector in the fifth line of H_4^D in [49a]. By analogy, as the supply vector is identified by a zero restriction on foreign production, its simultaneous imposition on both vectors makes the third line of the corresponding restriction matrix H_4^S in [49b] become zero and y^* drop from the system:

$$[49a]\quad H_4^D = \begin{bmatrix} 1 & 0 & 0 & 0 & 0 \\ 0 & 1 & 0 & 0 & 0 \\ 0 & 0 & 1 & 0 & 0 \\ 0 & 0 & 0 & 1 & 0 \\ 0 & 0 & 0 & 0 & 0 \\ 0 & 0 & 0 & 0 & 1 \end{bmatrix}, \qquad [49b]\quad H_4^S = \begin{bmatrix} 1 & 0 & 0 & 0 & 0 \\ 0 & 1 & 0 & 0 & 0 \\ 0 & 0 & 0 & 0 & 0 \\ 0 & 0 & 1 & 0 & 0 \\ 0 & 0 & 0 & 1 & 0 \\ 0 & 0 & 0 & 0 & 1 \end{bmatrix}.$$

The appropriate test is a likelihood ratio test, where the ratios in the test statistic contain eigenvalues from the reduced rank regression under ®$_4$ in the numerator and those from the unrestricted reduced rank regression in the denominator (Johansen 1991). The statistic is asymptotically χ^2-distributed with $r \cdot \#$ degrees of freedom (here: 2 as the number of restrictions imposed on each vector, $\#$, is one). For each country ®$_4$ is clearly rejected at the 1 percent significance level both in the test of the restriction identifying demand (H_4^D) and in the test of the restriction identifying supply. (H_4^S) (see upper half of Table 20). This is reassuring because non-rejection would have pointed to misspecification as either p_{NCU} or y^* could then be dropped from the long-run model.[55] For the sake of completeness, this use of ®$_4$ as a misspecification test is extended to the other four variables of the system, x, q_{NCU}, p^*_{NCU}, f (results

[54] In this respect my application differs from the procedures described in Johansen and Juselius (1992) as they search for a purchasing power parity relation in each vector, which implies a homogeneity restriction on domestic and foreign prices rather than zero restrictions.

[55] Domestic potential GDP, considered as a crucial scale variable for the volume of exports supplied in Goldstein and Khan (1978), was dropped from an earlier specification of the VECM (not reported here), *inter alia* because a ®$_4$-type of test dramatically failed to reject the irrelevance of the variable.

not reported in Table 20): For Germany and Canada, all eight null hypotheses are rejected with probabilities [0.00] or [0.01], for the United States the null is rejected in three out of the four cases but dropping f from the cointegration space is "accepted" surprisingly well by the data ($\chi^2(2) = 0.01$). Nonetheless, the globalization variable is kept not only because it is economically motivated but also because this economic relevance will be reflected in the statistical tests once over-identifying restrictions are put on the demand vector. This will become clear in the next paragraphs.

Turning to the second type of tests, their null hypothesis, labeled $®_6$ by Johansen and Juselius (1992: 225), states that a certain set of restrictions holds for one part of the cointegration space, while the other part remains unrestricted.[56] The test is useful when one looks for one of the r cointegrating vectors, e.g. an export demand (supply) relationship derived from economic theory. Yet, as it is impossible to test an identifying restriction statistically, testing $®_6$ is not feasible for the respective just identified versions of the demand and supply vectors. This is why for each vector one over-identifying restriction is considered. To select it, I anticipate on the results from restrictions on each cointegrating vector (shown in Tables 21, 22, and 23) in order to pick the least controversial over-identifying restriction for each country and market side. As to demand, this is the constant-returns-to-scale (CRS) hypothesis for the demand vector, a homogeneity restriction (1,–1) on the coefficients of export volumes and foreign production. As to supply, there is no uniformly accepted over-identifying restriction. The homogeneity restriction on export prices and domestic producer prices resulting both from the mark-up model without conjectural

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[56] In Johansen and Juselius (1992), a third case, $®_5$, is considered which I do not follow here. It is the hypothesis that one or more (r_1) vectors are *completely* known. They use this hypothesis for testing e.g. whether the pure purchasing power parity relation is present in one of the cointegrating vectors implying two homogeneity restrictions (between the domestic and the foreign price level and between price levels and the nominal exchange rate) as well as zero restrictions on all other I(1) variables in the cointegration space.

variations (Dornbusch 1987) and from the relative-profitability concept (Goldstein and Khan 1978) is implemented for Canada and Germany, while a zero restriction on the globalization parameter quite evident in the unrestricted reduced-rank regression is used for the United States.

Formally, ®$_6$ partitions the r-dimensional cointegration space into two groups containing r_1 and r_2 vectors, respectively ($r = r_1 + r_2$), and imposes a set of restrictions on the first group according to

[50] ®$_6$: $\beta = (H_6\varphi, \beta_2)$, with dimensions $H_6(n \times s), \varphi(s \times r_1), \beta_2(n \times r_2)$,

where matrix or vector β_2 contains the unrestricted column(s) of β. In the present example I have $r_1 = r_2 = 1$ and $s = 4$. Again two distinct tests are carried out, the first one to find the over-identified demand vector, the second one to find the respective over-identifying supply vector. For the same order of variables appearing in the lines of H as above, $(x, q_{NCU}, y^*, p^*_{NCU}, p_{NCU}, f)$, the corresponding restriction matrices read:

[51a] $H_6^D = \begin{bmatrix} 1 & 0 & 0 & 0 \\ 0 & 1 & 0 & 0 \\ -1 & 0 & 0 & 0 \\ 0 & 0 & 1 & 0 \\ 0 & 0 & 0 & 0 \\ 0 & 0 & 0 & 1 \end{bmatrix}$; [51b] $H_6^{S,j} = \begin{bmatrix} 1 & 0 & 0 & 0 \\ 0 & 1 & 0 & 0 \\ 0 & 0 & 0 & 0 \\ 0 & 0 & 1 & 0 \\ 0 & -1 & 0 & 0 \\ 0 & 0 & 0 & 1 \end{bmatrix}$, $j = 1,2$; $H_6^{S,US} = \begin{bmatrix} 1 & 0 & 0 & 0 \\ 0 & 1 & 0 & 0 \\ 0 & 0 & 0 & 0 \\ 0 & 0 & 1 & 0 \\ 0 & 0 & 0 & 1 \\ 0 & 0 & 0 & 0 \end{bmatrix}$,

where j=1, 2, stands for Canada (1) or Germany (2). The likelihood function is maximized by a switching algorithm (Johansen and Juselius 1992: 233–235); the appropriate likelihood ratio test statistic contains eigenvalues from the restricted first subgroup and the unrestricted second subgroup in the numerator as well as eigenvalues from the unrestricted reduced rank regression in the denominator. It is asymptotically χ^2-distributed with $(n - s - r_2)r_1$ degrees of freedom. The

Table 20: Tests on Single (Over-)Identified Lon-Run Relationships in the Cointegration Space (r=2)

Hypothesis	LR-Statistic[a]	DF[b]	Probability
Hypotheses of type ®$_4$ [c]			
United States			
Demand identification (p_{NCU} redundant)	22.86	2	[0.00]
Supply identification (y* redundant)	13.56	2	[0.00]
Canada			
Demand identification (p_{NCU} redundant)	11.83	2	[0.00]
Supply identification (y* redundant)	9.47	2	[0.01]
Germany			
Demand identification (p_{NCU} redundant)	15.57	2	[0.00]
Supply identification (y* redundant)	22.87	2	[0.00]
Hypotheses of type ®$_6$ [d]			
United States			
Export demand with CRS[e]	1.95	1	[0.16]
Export supply without globalization variable[f]	0.00	1	[0.96]
Horizontal export supply curve[h]	3.43	1	[0.06]
Canada			
Export demand with CRS[e]	0.07	1	[0.79]
Export supply with relative profitability[g]	0.21	1	[0.64]
Horizontal export supply curve[h]	4.06	1	[0.04]
Germany			
Export demand with CRS[e]	0.93	1	[0.34]
Export supply with relative profitability[g]	0.01	1	[0.92]
Horizontal export supply curve[h]	13.39	1	[0.00]

[a]Likelihood ratio test statistic which is χ^2 – distributed with the number of degrees of freedom indicated in the third column. – [b]Number of degrees of freedom. It is r·# (number of cointegration vectors times number of restrictions) for hypotheses of type ®$_4$ and $(n-s-r_2)\cdot r_1$ for hypotheses of type ®$_6$ (r_1 is the number of cointegration vectors in the subgroup, i.e. 1 if one looks for one single vector). – [c]®$_4$ states that a particular restriction holds in each vector of the cointegration space. The aim is to judge whether the unrestricted stationary long-run relationships are in fact linear combinations of the structural economic relationships. As both demand and supply are identified by zero restriction, ®$_4$ ultimately tests for exclusion from the system of p_{NCU} and y*, respectively. – [d]®$_6$ states that a vector or a set of vectors satisfying specific assumptions (but leaving some parameters to be estimated freely) is in the cointegration space. Therefore the latter is partitioned into a subset of r_1 vectors satisfying the restrictions of interest and the $r_2(=r-r_1)$ other vectors. The LR test checks whether this partitioning is supported by the data. – [e]Export demand is identified by restricting the coefficient of domestic prices to zero; as an over-identifying restriction, the coefficient of foreign production is set equal to one corresponding to constant returns to scale (CRS) in the production function of the rest of the world. – [f]Export supply is identified by restricting the coefficient of foreign production to zero. In this just identified specification, the coefficient of f is very close to zero and insignificant. – [g] In the just identified supply relationship, the coefficients of export prices and domestic prices are set to 1 and –1, respectively, reflecting that an increase in q_{NCU} has the same long-run effect on the export volume supplied as an equally-sized decrease in p_{NCU}. – [h]Export supply is identified by restricting the influence of foreign production to zero. Then the coefficient of export volumes is set equal to zero and the vector is normalized to export prices.

results demonstrate that the data fit the assumed economic long-run relationships as none of the hypotheses sketched in [51a] and [51b] are rejected for any country (lower half of Table 20).

Based on the identifying restriction for the supply vector (i.e. the exclusion of y^*), Table 20 also reports the results of the over-identifying restriction of export supply being infinitely price-elastic in the long run, a wide-spread assumption in the trade literature that will become important later in this section. This assumption is implemented by a zero restriction on the volume of exports in the supply vector. The restriction is rejected for every country at the 10 percent significance level, especially clearly so for Germany. Yet at the 5 percent level, it cannot be rejected for the United States and is close to non-rejection for Canada.

To conclude, the restriction identifying the demand vector does not hold in the other vector and vice versa for the supply vector. Moreover, when searched for separately, there is strong evidence for an export demand vector characterized by constant returns to scale in aggregate foreign output and for an export supply vector characterized by a one-to-one co-movement of export prices and producer prices in the long-rung (except for the United States). These findings may help justify the identification strategy developed in the last section and pave the way to simultaneous testing of over-identifying demand and supply restrictions. This is done in the next sub-section.

5.6. Testing For Over-Identifying Restrictions in the Cointegrating Vectors and for Weak Exogeneity

In this section I impose the restrictions identifying long-run export demand and supply jointly upon the cointegration space. The result serves as a basis for the examination of various other economic hypotheses which are then tested as over-identifying restrictions. First hypotheses on coefficients of the cointegrating vector are presented, then restrictions on the loading coefficients are checked to test for weak exogeneity of variables and thus for possibilities of slimming the VECM in a further step.

5.6.1. Presentation of the Test Results

The detailed results for all tests discussed in 5.6. are reported in the long tables of the appendix sub-section 5.6.7. (Tables 21, 22, and 23). As all the tests carried out throughout section 5.6. are documented in these tables, it is necessary to explain how they are organized. Each table contains more than a dozen numbered horizontal "blocks". Each block regroups four lines and reports the test results for one specific (or several joint) over-identifying restriction(s). Each table starts off reporting the results of the just identified system of export demand and supply in block 1. It serves as a starting point from which both coefficient tests on the cointegrating vectors and exogeneity tests are implemented. Furthermore, looking at the long-run model in the respective block 1 allows some preliminary international comparisons. Blocks 2 through 10 (United States: 9) show over-identifying restrictions on elements of the cointegrating vectors. Thereafter, results of tests for week exogeneity are reported. Finally, at the end of each table, the "successful" restrictions on each beta-vector are combined with the non-rejected exogeneity restrictions to yield the best restricted long-run model available for each country. Only then will a detailed comparison between the countries in the sample be carried out.

As to the first type of tests, restrictions on each beta-vector are set according to formula (6) in Johansen and Juselius (1994: 14):

[52] $\beta = (H_1\varphi_1,...,H_r\varphi_r)$, $H_i(n \times s_i), \varphi_i(s_i \times 1), \forall i, i = 1,...,r$,

where in my case $r = 2$. H_1 is the matrix of restrictions imposed on the export demand relationship, H_2 the restriction matrix for the export supply relationship and φ_1, φ_2 are the corresponding vectors of freely estimated long-run coefficients.

The left half of each table shows the beta-coefficients and corresponding *t*-values (in brackets) obtained from division of the coefficient by its standard deviation. In each block the upper half contains the values of the first vector, the demand vector, which is normalized on the quantity of exports, whereas the lower half reports the coefficients of the supply vector, normalized on the level of export prices in national currency units. Multiplying each coefficient with the log of the corresponding I(1) variable in quarter $(t-1)$ and summing up over the line yields the "clean" disequilibrium, i.e. the deviation of export volumes from their long-run level dictated by the demand relationship at the end of that quarter (which is zero if export demand is in equilibrium) and the deviation of export prices from their long-run level dictated by the supply relationship. After bringing all variables except for x_{t-1} to the right hand-side, the long-run export demand for the United States reads

[53] $x_{t-1} = -1.47 q_{NCU,t-1} + 1.85 y^*_{t-1} + 1.16 p^*_{NCU,t-1} - 0.00 f_{t-1} + error_{t-1}$.

Hence, multiplying each figure in the left half of the tables by –1 gives the long-run elasticity of aggregate real export demand to the respective determinant (upper half of a block) and the long-run elasticity of the export price level with respect to its determinants (lower half of the block). The concept of a long-run equilibrium level of export prices is an indication of what the long-run price level of exports should be *given* the realizations of the other variables and should not be mixed up with a reaction function of exporters. To restore the supply equilibrium after a disturbance, it could be that other variables than export prices themselves adjust.

It is the values in the central right part of each table that show how strongly and in which direction an endogenous I(1)-variable changes in quarter *t* when export quantities and prices deviate positively from their equilibrium levels in

$(t-1)$, and whether these changes are significant (*t*-values in brackets). The upper half of each block reports the loading coefficients and their *t*-values corresponding to demand disequilibria, the lower half the ones corresponding to supply disequilibria. For example, the numbers 0.05 and –0.18 in Table 21, block 1, column Δy^*, mean that if U.S. exports exceed their long-run demand by, say, 1 percent, this leads to an increase in foreign industrial production by 0.05 percent one quarter later; accordingly, export prices 3 percent in excess of their long-run level provoke a decrease in foreign production in the order of $0.18 \times 3\% = 0.54\%$. If quantities and prices are instead below their long-run levels, the change in y^* has the opposite sign, of course. The *t*-values of loading coefficients 0.05 (3.04) and –0.18 (–3.42) indicate that the reactions are significant implying a feedback from American exports to international economic activity.

However, foreign production is not the only variable to react to a demand disequilibrium. As one recognizes reading the first two lines of block 1 from the left $(...\Delta x)$ to the right $(...\Delta f)$, almost one fifth (0.19) of, say, an excess of exports over the long-run level of demand is corrected by a decrease in export volumes themselves one quarter later. Foreign prices shrink, too, although the effect is at the limit of insignificance. The reason for the reduction in Δp^*_{NCU} is either an appreciation of the dollar or a rebate offered by foreign competitors as excess U.S. exports boost world supply and thus exert some pressure on prices.

What happens in case of a supply disequilibrium in U.S. exports? When export prices do not correspond to their long-run level expressed by the error term in the second cointegrating relationship, foreign production again is not the only variable to adjust. However, export prices themselves do not seem to be involved in the error-correction mechanism. Rather, domestic prices are lifted when export prices are in excess of domestic export supply capacities, preventing American firms from withdrawing too many goods and services from the domestic market in favour of profitable foreign markets. At the same time,

Δp^{*}_{NCU} decreases quite strongly, i.e. the dollar appreciates or foreign competitors reduce their prices, which seems economically implausible to me. It will be seen later whether it is accepted by the data to suppress this channel by setting Δp^{*}_{NCU} exogenous.

As the trade intensity of global production is set exogenous from the outset, there are only five endogenous variables and thus five loading coefficients per cointegration relation. In case of Canada, however, $p = 5$ in the just-identified VECM implies some problems. Especially the residuals do not satisfy the required assumptions as they exhibit non-normality and serial correlation. Neither increasing the lag length of the model nor implementing further impulse dummy variables helps correcting these deficiencies. Yet the problems disappear when setting y^{*} exogenous, probably thanks to the contemporaneous change in foreign (i.e. U.S.) production, $\left(\Delta y_t^{*}\right)$, the coefficient of which is significant and has the expected positive sign. Just as for the exogeneity of f in the Canadian and the U.S. cases, I disregard the result of the statistical test for weak exogeneity (explained below), which would not allow for this rescue measure. However, as Canada is a small open economy, there is a strong economic prior in favor of assuming that changes in the country's exports do not affect the overall level of production in the rest of the world and even in the United States.

Finally, the right end of Tables 21, 22, and 23 indicate results from likelihood ratio (LR) tests and multivariate residual tests (normality and autocorrelation of first and higher order). The former indicate the appropriateness of the over-identifying restriction(s) in the corresponding block. In case of restrictions on each of the beta-vectors, the relevant LR test statistic is based on the log of the ratio between the likelihood from the restricted reduced-rank regression (numerator) and the likelihood from the unrestricted one (denominator). The statistic is asymptocically χ^2-distributed (Johansen and Juselius 1994: 16). The number of degrees of freedom, ν, is:

[54] $$\nu = \sum_{i=1}^{r} n-(r-1)-s_i = \sum_{i=1}^{r} k_i-(r-1)\,.$$

The left half of [54] is taken from Johansen and Juselius (op. cit.: 24). Using the number of total restrictions in each cointegrating vector, k_i, with $k_i = n - s_i$, the right half shows that ν equals the total number of *over-identifying* restrictions imposed on the system as a whole. Thereby the contribution of each vector i to ν is k_i less the number of necessary identifying restrictions per vector $(r-1)$. The three figures in the penultimate columns of Tables 21–23 show for each block the value of the LR-statistic, the degrees of freedom ν and the significance level (SL), i.e. the error probability of rejecting the set of over-identifying restrictions in that block. The last column reports probabilities from various residual tests for autocorrelation and normality. Tests for hetero-scedasticity are briefly mentioned at the end of the footnotes in each table.

5.6.2. The Just-Identified Cointegrating Vectors

As far as the export demand relationship of the just-identified model is concerned, demand for U.S. exports turns out to be clearly more price-elastic than demand for Canadian and German exports. The long-run elasticity of American exports with respect to both own export prices and foreign producer prices in U.S. dollars exceeds one, they do not for Germany and Canada. The point estimator of the own-price elasticity is higher than the cross-price and/or exchange-rate elasticity only for the United States, not for the other two countries. Moreover, only for Germany is the long-run elasticity with respect to foreign production close to one (1.15) as theory predicts is consistent with constant returns to scale in foreign production. For Canada it is lower than one, for the United States it is higher, both pointing to increasing returns to scale (see section 2). Yet one can observe that the responsiveness to globalization is low in countries with a high production elasticity and high for countries with a low production elasticity. The coefficient of f is even insignificant for the United States and exceeds one for Canada, suggesting an increasing Canadian share in world exports in the long run, which is hard to believe. One would rather expect a declining tendency in the world market share for each of the three countries as more and more newly industrialized countries, e.g. China, enter the international division of labor. Although there should be no multicollinearity between y^* and f because the latter is already normalized by world production, maybe the positive deterministic trend characterizing time series makes it difficult to get the precise effects separated in the estimation. As it is hard to imagine that the three countries are this different as to their participation in world trade and as to the production function of their trading partners, it will be useful to impose some over-identifying restriction on the production coefficient in order to obtain a more reliable gauge of the globalization effect.

5.6.3. Over-Identifying Restrictions on Cointegrating Vectors (β)

As a result of this discussion, the following over-identifying restrictions are imposed one by one upon the just-identified demand vector (expressed as elasticities of export volumes):

– CRS: unit elasticity with respect to y^*, i.e. constant returns to scale (CRS) in foreign production
– REER: equally-sized but opposite-signed elasticities with respect to q_{NCU} and p^*_{NCU}, i.e. using the real effective exchange rate (REER) as a single price variable is appropriate to model export demand
– CMS: unit elasticity with respect to f, i.e. constant market share (CMS) of the country's exports in world export demand.

The test results for these restrictions are reported in blocks 2, 3, and 4, respectively. In addition to REER, a unit price elasticity of export demand (UPD) is tested (and not rejected) for Canada (Table 22, block 16).

Turning to the second vector, the long-run supply curve is upward-sloping as predicted by textbook microeconomic theory for Germany. The coefficient implies that a one-percent price increase leads to quite strong a reation of export supply in the long-run (3.2%≈(1/0.31)%) but supply is not infinitely price-elastic. However, the finding of a negative price-elasticity of export supply for Canada and the United States is puzzling. It is true that average costs deline with quantities in industries with fixed costs allowing for a falling price as quantities expand, but this reasoning holds for a given capital stock and only as long as capacity is not binding. At the aggregate level, the expansion of exports over time involves costly investment into new capacity, thus lowering supply prices as exports increase can hardly be conceived as a genuine reaction by suppliers. Yet exporters may be unable to set their export prices because the latter are dictated by a world market characterized by new entrants at lower marginal costs, the only means to keep pace with the falling price trend might be to invest in new capacity if this is associated with a lower marginal cost. A look at the corresponding loading coefficient $\hat{\alpha}_{22}$ for the two countries with "wrongly-signed" supply relationships reveals that they have the lowest *t*-values in absolute terms, putting weight on the idea of the export price level being exogenous to the supply vector. Another, more technical explanation could be that the same supply identification strategy does not fit all three countries in my sample. Therefore, alternative identifying restrictions or specifications will be discussed with regard to their effect on the slope of the supply curve in sub-section 5.7.

The second detail of interest in the supply relationship is the long-run elasticity of export prices with respect to both domestic and foreign producer prices. As pointed out by Dornbusch (1987), domestic marginal producer costs (empirically best reflected by the PPI) are crucial but whether the mark-up is constant or whether it varies allowing for some reaction to foreign price changes depends on the structure of the theoretical model. A variable mark-up seems to be reflected by U.S. data as the long-run coefficient of domestic producer prices is well below one but seems to add up to about one with the coefficient of foreign prices. In case of Canada, the question is left to further testing of over-identifying restrictions because on the one hand, the coefficient of domestic prices is itself close to one, but on the other hand, it adds up to a value slightly above one with the coefficient of foreign prices. Besides an equilibrium relationship between domestic and foreign prices, a positive value of the p^*_{NCU}-parameter may also reflect an incomplete exchange-rate pass-through on export prices as export prices in national currency units can be lifted in presence of a permanent depreciation of the domestic currency. Against this background, the negative p^*_{NCU}-parameter for Germany is not compatible with either case of export price formation in the Dornbusch model which does not care about the volume of exports. In the German case, however, the supply vector from Table 23, block 1, may be re-normalized by export volumes to become

[55] $$x = 3.22 q_{NCU} - 3.19 p_{NCU} + 1.10 p^*_{NCU} + 0.94 f \text{ ,}$$

which has a more intuitive interpretation: an increase in prices set by foreign competitors allows for an expansion of the volume of exports supplied. This expansion occurs along the increasing long-run supply curve. As a by-product of equation [55], one may want to recognize that in presence of a non-horizontal long-run export supply curve the constant mark-up hypothesis by Dornbusch (1987) and relative-profitability hypothesis by Goldstein and Khan (1978) are tested for by the same over-identifying restriction on the domestic producer price elasticity, i.e. the q_{NCU}-parameter and the p_{NCU}-parameter having opposite signs but being equal in absolute terms. What the constant mark-up hypothesis postulates in addition to relative profitability, is no influence from foreign prices on export prices.[57]

[57] In Goldstein and Khan (1978) the structural export supply equation does not contain a foreign price index either but unlike Dornbusch (1987), this absence does not result from a

The third point of attention concerns the globalization effect on export prices. For Germany, the expected negative sign is confirmed by the data, for the United States there seems to be no influence at all, and for Canada the coefficient is even positive. But a closer look shows that this is not an additional puzzle but ultimately has to do with the “wrong” slope of the supply curve. To see this, re-normalize Canadian export supply by x as done for Germany in [55] to get

[56] $x = -2.08 q_{NCU} + 1.90 p_{NCU} + 0.48 p^{*}_{NCU} + 1.19 f$,

i.e. an expression quite similar to the one in [55] were it not for the “wrong” signs of the export price and domestic producer price parameters. These do not affect the validity of the simple-mark-up hypothesis but the one of the relative-profitability assumption as now the parameter of relative profitability as a whole ($q_{NCU} - p_{NCU}$) has the wrong sign.

The analysis of the just-identified supply vector underlines the interest in the following over-identifying restrictions upon the just-identified supply vector (expressed as elasticities of export prices):

– RP: unit elasticity with respect to p_{NCU}, i.e. relative profitability (RP) is one relevant price variable
– CM: Like RP *plus* zero restriction on p^{*}_{NCU}, i.e. export prices are obtained by a constant mark-up (CM) on marginal cost (proxied by p_{NCU})
– NG: zero restriction on f, i.e. no globalization (NG) influence on export prices
– HS: zero restriction on x, i.e. infinitely price-elastic export supply implying a horizontal supply (HS) curve

specific behavioral assumption within a profit-maximization framework, so their model would not explicitly “forbid” a foreign price variable.

– CMSS: unit elasticity with respect to f (only for Canada and Germany), i.e. supply keeps pace with the rising trade intensity of global production and aims at a constant market share from the supply-side (CMSS).

The test results for these restrictions are reported in blocks 5 through 8 (Table 21) or 5 through 9 (Tables 22 and 23).

The final specification of the cointegration space for the United States contains CRS, REER, and NG, not to forget the just-identifying restrictions. With variables in the order $(x, q_{NCU}, y^*, p^*_{NCU}, p_{NCU}, f)$ corresponding to the rows, the corresponding restriction matrices for the demand and the supply vector, H^D_{US} and H^S_{US} have the form

$$[57] \quad H^D_{US} = \begin{bmatrix} 1 & 0 & 0 \\ 0 & 1 & 0 \\ -1 & 0 & 0 \\ 0 & -1 & 0 \\ 0 & 0 & 0 \\ 0 & 0 & 1 \end{bmatrix}; \quad H^S_{US} = \begin{bmatrix} 1 & 0 & 0 & 0 \\ 0 & 1 & 0 & 0 \\ 0 & 0 & 0 & 0 \\ 0 & 0 & 1 & 0 \\ 0 & 0 & 0 & 1 \\ 0 & 0 & 0 & 0 \end{bmatrix}; \chi^2(3) = 4.69[0.20].$$

In the long-run model for Canada, CRS, REER, UPD, RP and CMSS hold, which brings the restriction matrices into the form

$$[58] \; H^D_{CAN} = \begin{bmatrix} 1 & 0 \\ 1 & 0 \\ -1 & 0 \\ -1 & 0 \\ 0 & 0 \\ 0 & 1 \end{bmatrix}; \; H^S_{CAN} = \begin{bmatrix} 1 & 0 & 0 \\ 0 & 1 & 0 \\ 0 & 0 & 0 \\ 0 & 0 & 1 \\ 0 & -1 & 0 \\ -1 & 0 & 0 \end{bmatrix}; \gamma^2(5) = 4.73[0.45].$$

Finally, for Germany, CRS, RP, and CMSS cannot be rejected, which gives

$$[59]\ H_{GER}^{D} = \begin{bmatrix} 1 & 0 & 0 & 0 \\ 0 & 1 & 0 & 0 \\ -1 & 0 & 0 & 0 \\ 0 & 0 & 1 & 0 \\ 0 & 0 & 0 & 0 \\ 0 & 0 & 0 & 1 \end{bmatrix};\ H_{GER}^{S} = \begin{bmatrix} 1 & 0 & 0 \\ 0 & 1 & 0 \\ 0 & 0 & 0 \\ 0 & 0 & 1 \\ 0 & -1 & 0 \\ -1 & 0 & 0 \end{bmatrix};\ \chi^2(3) = 1.28[0.73].$$

Based on these test results, I can classify the hypotheses under consideration into

– those "accepted" for each country: CRS

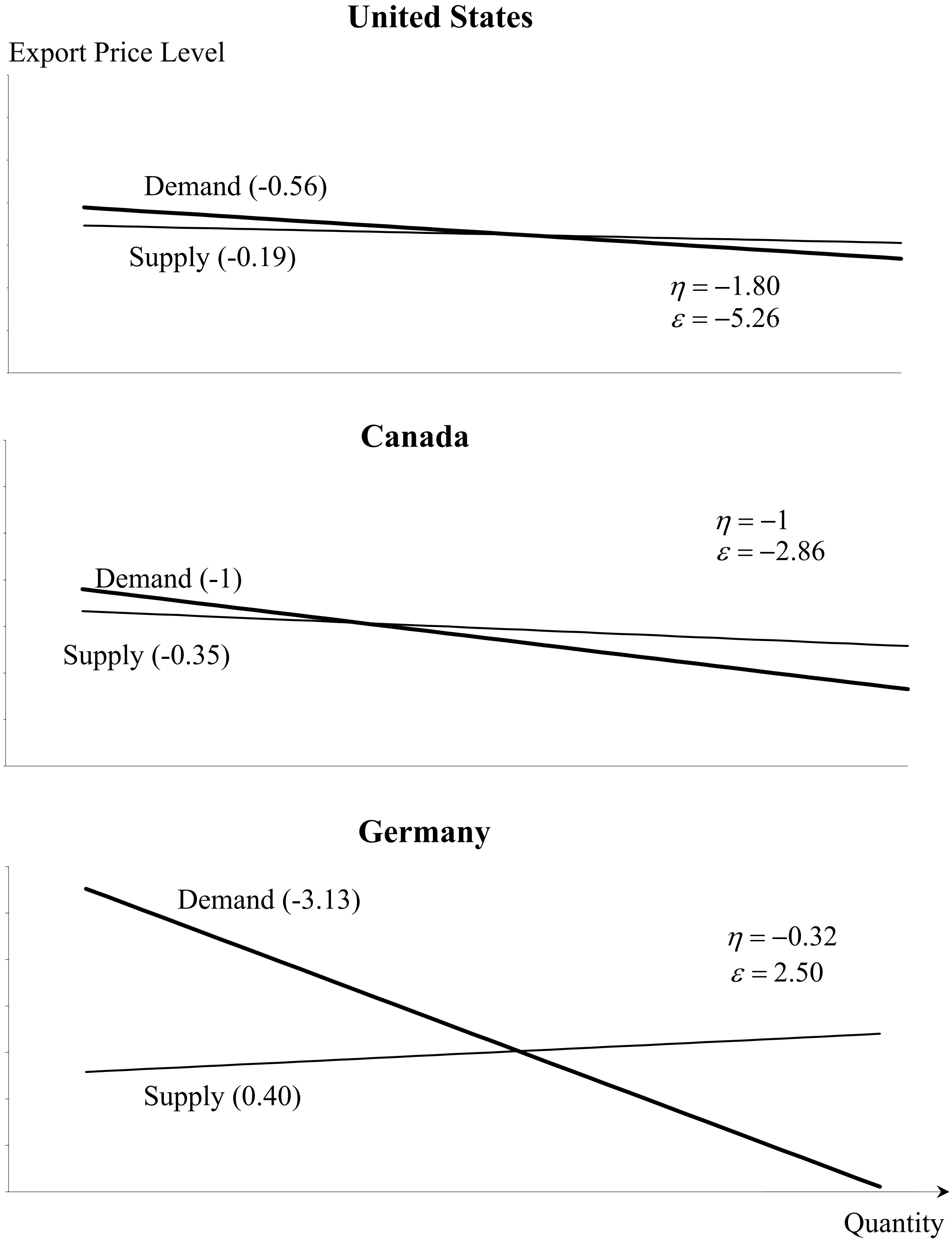
– those accepted for one or two countries: REER, (UPD), RP, NG, CMSS

– those rejected throughout: CMS, CM, HS

5.6.4. The Loading Coefficients as An Indication of How Disequilibria Are Corrected

Next I am interested in the loading coefficients that belong to the cointegrating vectors and in the related question of weak exogeneity of one or several I(1) variables. What is the expected sign for the ten (Canada: eight) loading coefficients of the VECM? First consider a demand disequilibrium. Let it be characterized by export volumes exceeding long-run export demand, i.e. by a point northeast to the long-run equilibrium, off the demand curve but on the supply curve, in the conventional scheme of downward-sloping demand and upward-sloping supply as it empirically holds for Germany (Figure 5). Then a stabilizing reaction would be a decrease in both export quantities and prices. Negative loading coefficients ($\alpha_{11} < 0$, $\alpha_{21} < 0$) would be a most obvious way to trigger such an adjustment. When the supply curve is falling (although less strongly so than the demand curve), as in the Canadian and the U.S. cases, the initial demand disequilibrium is southeast to the intersection of the curves implying, *ceteris paribus*, that export prices should instead rise ($\alpha_{21} > 0$) in

Figure 5: Long-run Slope Properties of Empirical Export Demand and Supply Curves[a]



[a]The actual slope of each curve is shown in brackets after the name of the curve. The long-run elasticities of demand (supply) $r_i(\varepsilon)$, are the inverses of the respective figures in brackets and are also reported.

order to guarantee a movement back into equilibrium along the supply curve. In turn, a supply disequilibrium (e.g. export prices above their long-run level determined by the second cointegration relationship) may be conceived as a point on the demand curve northwest of the equilibrium. A stabilizing reaction now consists of falling export prices and rising volumes ($\alpha_{12} > 0$, $\alpha_{22} < 0$), as long as the supply curve is not falling steeplier than the demand curve.

If the four loading coefficients in the error correction equations of export quantities and prices do not satisfy these sign requirements, this need not mean the system is unstable because some of the shift parameters are endogenous, too, and might help restore the equilibrium. As can be seen in the empirical application presented here, considering the loading coefficients of the just-identified long-run models (first blocks of Tables 21, 22, and 23), export *volumes* adjust to demand disequilibria in all countries but the role volumes play in the adjustment of supply disequilibria is not clear-cut in my sample: the coefficient has the expected positive sign but is at the limit of insignificance for Germany while being zero for the United States and significantly negative for Canada. Moreover, export *prices* do not appear to play any significant role in the correction of both demand and supply disequilibria, at least in the VECMs without over-identifying restrictions.[58]

As to the expected reaction of the shift parameters, first look at the case of a demand disequilibrium, again. Excess exports can be viewed as additional inputs for foreign firms and should therefore stimulate foreign production ($\alpha_{31} > 0$). Excess exports put pressure on prices of inputs and output set by local producers on foreign export markets, so $\alpha_{41} < 0$. Facing falling prices abroad, exporters shift part of their supply from export markets to the domestic market (shift described by $\alpha_{11} < 0$) thereby exerting pressure on domestic sales prices. One may therefore expect $\alpha_{51} < 0$. Empirically, however, this negative effect is only

[58] The loading coefficient in the ECM of export prices for Canada is a limit case ($t = -1.64$).

observed for Canada, probably due to the highly integrated North-American market (facilitating goods arbitrage) and to the high degree of openness of the country. Openness augments the effect on domestic prices of arbitrage between external and domestic markets.[59] Foreign prices are also found to react significantly only for Canada. The loading of industrial production abroad has the expected positive sign for the United States and Germany and is zero for Canada due to the *a priori* exogeneity restriction.

In case of a supply disequilibrium taking the form of excess export prices, foreign production is expected to be negatively affected as one of its inputs is unusually expensive ($\alpha_{32} < 0$). As foreign producers then recur to more foreign inputs, the latter should get more expensive ($\alpha_{42} > 0$). In a similar vein, attracted by the high price signal from export markets, domestic producers may want to shift part of their domestic supply to international markets, and the relative scarcity at home drives domestic producer prices up ($\alpha_{52} > 0$).

My empirical finding is that inter-country differences are bigger for the reaction to supply disequilibria than for the one to demand disequilibria. For Germany one observes neither domestic nor foreign producer price movements in response to excessive export prices, at least not from the error-correction terms of the just identified VECM. For the United States and for Canada, there are significant price responses which have the expected sign except for Δp^*_{NCU} in the U.S. model. Finally, foreign production, where taken into account, reacts significantly to a supply disequilibrium showing the expected negative sign for the United States but a positive sign for Germany.

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59 In 1999, the share of exports of goods and services in real GDP amounted to 43 percent for Canada but only to 30 percent for Germany and 12 percent for the United States (OECD 2002b).

5.6.5. The Concept of Weak Exogeneity

From the definition of weak exogeneity given in Engle et al. (1983: 282), two necessary conditions can be derived in the context of cointegration (Harris 1995: 99): (i) The economically interesting long-run coefficients contained in β are determined only by the conditional model, not by the marginal model, and (ii) the parameters in the conditional and in the marginal model must not be subject to the same restrictions, which is fulfilled by Gaussian errors. The decomposition of the original VECM [13c] into a conditional and a marginal model follows Johansen (1992b: 321). It is useful to rearrange the $(p \times 1)$-vector z_t of I(1) variables into x_t, the variable that will be supposed to be exogenous to the cointegration space, and the $((p-1) \times 1$-vector y_t of endogenous variables such that $z_t' = [y_t \quad x_t]'$.[60] The conditional model is the VECM determining Δy_t *given* Δx_t and reads (for lag length $l = 2$)

[60a] $\Delta y_t = \omega \Delta x_t + (\Gamma_{y1} - \omega \Gamma_{x1}) \Delta z_{t-1} + (\alpha_y - \omega \alpha_x) \beta' z_{t-1} + u_{yt} - \omega u_{xt}$,

whereas the marginal model determines Δx_t using the last equation of the VECM [13c]:

[60b] $\Delta x_t = \Gamma_{x1} \Delta z_{t-1} + \alpha_x \beta' z_{t-1} + u_{xt}$,

and the new coefficient ω is obtained as

[60c] $\omega = \Omega_{yx} \Omega_{xx}^{-1}$,

i.e. with the help of the auto-covariance matrix of the marginal model and the cross-covariances between the marginal and the conditional model.

Estimating [60a] and [60b] as a system is equivalent to estimating [13c] because $\Gamma_1 = \begin{bmatrix} \Gamma_{y1} \\ \Gamma_{x1} \end{bmatrix}$ and $\Pi = \alpha\beta' = \begin{bmatrix} \alpha_y \\ \alpha_x \end{bmatrix} \beta'$ and substituting [60b] into [60a]

60 Throughout 5.6.5., x stands for the vector of exogenous I(1) variables and y for the vector of endogenous I(1) variables.

makes all ω-terms drop out. Searching for weak exogeneity consists of checking whether the whole line of loading coeffients in the marginal model is zero, i.e. $\alpha_{x,j} = 0, \forall j, j = 1,...,r$. Economically speaking, a variable is weakly exogenous with respect to the long-run economic system, when it is not affected by any long-run disequilibrium ($\beta' z_{t-1} \neq 0$).[61] Generalizing the above example, x_t may be a $(p_x \times 1)$-vector (rather than a scalar), y_t a vector of order p_y with dimensions adding up ($p_x + p_y = p$).The null hypothesis of the test for weak exogeneity states

[61] $H_0 : \alpha_x = 0$.

As the number of endogenous variables exceeds the number of cointegration relationships in my case, the appropriate framework of analyzing the conditional and the partial model is reduced-rank regression. Once more the test statistic is of the likelihood-ratio type. Assume the exogeneity of the i-th variable to be under investigation. Let $\tilde{\lambda}_i$ denote the r non-zero eigenvalues of the reduced-rank regression under the null and $\hat{\lambda}_i$ those of the unrestricted reduced-rank regression. Then the test statistic is

[62] $T \sum_{j=1}^{r} \ln\left\{\left(1 - \tilde{\lambda}_{ij}\right) / \left(1 - \hat{\lambda}_{ij}\right)\right\}$

and is distributed as $\chi^2(r \cdot p_x)$. In the present context, this means that two degrees of freedom have to be taken into account for each variable assumed to be weakly exogenous.

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61 I only ask for one or several *lines* of loading coefficients (rather than single values) being insignificant. It would also be meaningful to check whether $\alpha_{x,j} = 0$ for some j but not for all of them, i.e. to look whether a variable is weakly exogenous with respect to only one cointegration vector (e.g. export demand).

5.6.6. Testing for Weak Exogeneity in the Aggregate Export System

To test the hypothesis of weak exogeneity in the just identified model according to [61], row restrictions are placed on the (5×2) (Canada: 4×2) matrix of loading coefficients in the just identified reduced rank model. Technically, this is achieved by specifying a $(p \times (p - \#))$ matrix A of linear restrictions, which reduces α to the $((p - \#) \times r)$ matrix α_0 of rows that are non zero under the null (Harris 1995: 101). The symbol # equals the number of row restrictions imposed on α and thus the number of variables that are weakly exogenous under the null. The null hypothesis then amounts to [63a]. Yet in the software package used to carry out my cointegration analysis (CATS in RATS, Version 5.0), an alternative specification is used to express the same restrictions which consists of finding a $(p \times \#)$ matrix B satisfying [63b]. Obviously, matrix B is orthogonal to A. Both notations will be presented in the examples below.

[63a] $H_0 : \alpha = A\alpha_0$; [63b] $H_0 : B'\alpha = 0$.

The results of the just identified VECM make it worthwhile to test for weak exogeneity of (a) export prices, (b) foreign producer prices, (c) domestic producer prices, (d) both export and foreign producer prices as well as (e) both foreign and domestic producer prices. In the U.S. VECM, I check for (a) through (d), in the Canadian one for (a) through (c) and in the German one for (a) through (c) and (e). With the error correction equations appearing in the familiar order $(\Delta x_t, \Delta q_{NCU,t}, \Delta y_t^*, \Delta p_{NCU,t}^*, \Delta p_{NCU,t})$, the corresponding restriction matrices are:[62]

[62] To obtain the matrices for Canada, eliminate those columns in A which have a non-zero entry in the third row, then eliminate the third lines in all A and B matrices.

[64] (a) $B=\begin{bmatrix}0\\1\\0\\0\\0\end{bmatrix}$; $A=\begin{bmatrix}1&0&0&0\\0&0&0&0\\0&1&0&0\\0&0&1&0\\0&0&0&1\end{bmatrix}$; (b) $B=\begin{bmatrix}0\\0\\0\\1\\0\end{bmatrix}$; $A=\begin{bmatrix}1&0&0&0\\0&1&0&0\\0&0&1&0\\0&0&0&0\\0&0&0&1\end{bmatrix}$;

(c) $B=\begin{bmatrix}0\\0\\0\\0\\1\end{bmatrix}$; $A=\begin{bmatrix}1&0&0&0\\0&1&0&0\\0&0&1&0\\0&0&0&1\\0&0&0&0\end{bmatrix}$; (d) $B=\begin{bmatrix}0&0\\1&0\\0&0\\0&1\\0&0\end{bmatrix}$; $A=\begin{bmatrix}1&0&0\\0&0&0\\0&1&0\\0&0&0\\0&0&1\end{bmatrix}$;

(e) $B=\begin{bmatrix}0&0\\0&0\\0&0\\1&0\\0&1\end{bmatrix}$; $A=\begin{bmatrix}1&0&0\\0&1&0\\0&0&1\\0&0&0\\0&0&0\end{bmatrix}$.

The test results can be found in blocks 10–13 of Table 21, blocks 11–13 of Table 22, and blocks 11–14 of Table 23. The hypothesis of weakly exogenous export prices in domestic currency units cannot be rejected for any country. The evidence for exogenous foreign prices and exchange rates (p^*_{NCU}) is not so clear-cut when the hypothesis is applied to the version of the VECM without over-identifying restrictions: the hypothesis may pass at the 5 percent significance level for all three countries but is rejected at the 10 percent level for Canada and verging at the limit of rejection for the United States, while being well accepted for Germany. The difference between Europe's biggest economy and both North-American states is still larger when it comes to domestic producer prices. For the former they seem as exogenous as one can imagine ($\chi^2(2) = 0.01$), for the latter two the null hypothesis is rejected at the 5 percent level. As to the combined hypotheses (d) and (e), they are well supported by the U.S. (d) and the German data (e), respectively.

If one came to ask which error-correction equation to remove "legitimately" from the just identified VECM at this stage of the discussion, one would tend to decide against export prices in all three cases, further to both other price variables in the German case and maybe against p^*_{NCU} in the American case. Before doing so, however, an important sensitivity check is recommendable. Actually, it makes sense to look whether the exogeneity results change once all non-rejected over-identifying β-restrictions lie on the cointegration space. As loading coefficients are part of the short-run adjustment, putting a clear long-run economic structure on export demand and supply may well result in another adjustment path. In this respect, the Canadian case is interesting: alternative testing of (a) and (b) in the fully over-identified model leads to virtually the same LR-test result, i.e. the "endogeneity advantage" of foreign prices over Canadian export prices has disappeared. For the sake of plausibility, the econometrician may want to and is now allowed to prefer a final specification with weakly exogenous foreign prices (block 16) to the alternative of exogenous export prices because the variables of the system have been chosen to explain the interaction of demand and supply on export markets rather than the structural determinants of foreign prices or exchange rates.

As far as the U.S. model is concerned, the LR test results are less sensitive insofar as the decision on whether or not to exclude the $\Delta p^*_{NCU,t}$-equation from the VECM remains as tight as before: hypothesis (d) is attributed a probability of 0.12 (block 16), when formulated on the over-identified model, and one of 0.24 (block 13) for the just identified one.[63] So unlike the analysis of the just identified model, there is now scope for considering foreign prices as weakly exogenous in the U.S. and Canadian models, let alone in the German one, where both domestic and foreign producer prices are put exogenous without the slightest difficulty. However, as it seems difficult to imagine a market without

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63 When testing for weak exogeneity of the foreign price index alone, non-rejection is more difficult to obtain in the over-identified model (probability of 0.07 in block 15).

any price adjustments to long-run economic conditions, the error-correction equation for the export price change is kept in the German model although its highest *t*-value in absolute terms is slightly below the 10 percent critical value and although the joint hypotheses of three over-identifying restrictions and exogeneity of all three price indices cannot be rejected by the LR test (result not reported in Table 23). But it will be seen in the partial model of the section 6 that the index of export prices has an important (and then significant) role to play.

The joint tests on α- and β-restrictions are now complete. The economically "best" results are those of block 16 in both Tables 21 and 22 as well as block 14 in Table 23.

5.6.7. Appendix to Sub-Section 5.6.: Detailed Estimation and Test Results for the United States, Canada, and Germany

Table 21: Cointegration Relations and Loading Coefficients for Various Overidentifying Restrictions: United States[a]

No.[b]	Cointegration coefficients (and t-values)						Loading coefficients (and t-values) in equation …					Test for Restrictions[c]	Residual Tests[d]
	x	q_{NCU}	y^*	p^*_{NCU}	p_{NCU}	f	$\ldots\Delta x$	$\ldots\Delta q_{NCU}$	$\ldots\Delta y^*$	$\ldots\Delta p^*_{NCU}$	$\ldots\Delta p_{NCU}$	χ^2-statistic (DF) [SL]	L-B (25) LM(1) LM (4) Normality
1	1	1.47	–1.85	–1.16	0	0.00	–0.19	0.01	0.05	–0.08	–0.01	—	[0.40]
	—	(9.40)	(–6.11)	(–9.23)	—	(0.02)	(–4.70)	(0.78)	(3.04)	(–1.52)	(–0.87)	—	[0.09]
	0.18	1	0	–0.16	–0.76	–0.01	0.01	–0.06	–0.18	–0.52	0.18	—	[0.46]
	(3.49)	—	—	(–2.47)	(–17.34)	(–0.09)	(0.07)	(–1.17)	(–3.42)	(–2.79)	(3.17)		[0.16]
2	1	1.28	–1	–1.34	0	–0.64	–0.24	0.02	0.07	0.03	–0.04	1.95	[0.37]
	—	(9.88)	—	(–14.15)	—	(–8.06)	(–4.49)	(0.91)	(3.61)	(0.36)	(–1.87)	1	[0.17]
	0.30	1	0	–0.38	–0.64	–0.22	0.04	–0.09	–0.23	–0.41	0.22	[0.16]	[0.28]
	(7.21)	—	—	(–6.52)	(–14.79)	(–2.82)	(0.28)	(–1.44)	(–3.95)	(–1.87)	(3.30)		[0.15]
3	1	1.28	–1.20	–1.28	0	–0.48	–0.23	0.02	0.06	–0.00	–0.03	1.25	[0.37]
	—	(11.86)	(–9.56)	(–11.86)	—	(–3.64)	(–4.79)	(0.96)	(3.37)	(–0.06)	(–1.41)	1	[0.15]
	0.26	1	0	–0.30	–0.69	–0.15	0.03	–0.07	–0.21	–0.43	0.21	[0.26]	[0.30]
	(5.82)	—	—	(–4.87)	(–15.29)	(–1.74)	(0.17)	(–1.30)	(–3.81)	(–2.07)	(3.35)		[0.17]
4	1	1.19	–0.54	–1.44	0	–1	–0.29	0.04	0.15	0.14	–0.11	4.25	[0.37]
	—	(10.31)	(–6.34)	(–15.19)	—	—	(–3.08)	(1.29)	(4.25)	(1.08)	(–2.78)	1	[0.17]
	0.47	1	0	–0.66	–0.47	–0.51	0.16	–0.12	–0.32	–0.37	0.28	[0.04]	[0.23]
	(13.74)	—	—	(–11.35)	(–9.59)	(–8.91)	(0.77)	(–1.54)	(–4.26)	(–1.28)	(3.29)		[0.17]
5	1	1.55	–2.76	–0.90	0	0.66	–0.12	0.01	0.03	–0.13	–0.01	4.50	[0.39]
	—	(7.81)	(–6.89)	(–5.65)	—	(2.06)	(–3.67)	(1.09)	(2.55)	(–2.96)	(–0.36)	1	[0.05]
	0.01	1	0	0.18	–1	0.29	0.14	–0.01	–0.10	–0.40	0.10	[0.03]	[0.64]
	(0.12)	—	—	(4.00)	—	(2.88)	(1.47)	(–0.21)	(–2.76)	(–3.27)	(2.62)		[0.19]

Table 21 continued

No.[b]	Cointegration coefficients (and *t*-values)						Loading coefficients (and *t*-values) in equation …					Test for Restrictions[c]	Residual Tests[d]
	x	q_{NCU}	y^*	p^*_{NCU}	p_{NCU}	f	$\ldots\Delta x$	$\ldots\Delta q_{NCU}$	$\ldots\Delta y^*$	$\ldots\Delta p^*_{NCU}$	$\ldots\Delta p_{NCU}$	χ^2-statistic (DF) [SL]	L-B (25) LM(1) LM (4) Normality
6	1	1.11	–2.00	–0.89	0	0.06	–0.19	0.02	0.04	–0.09	–0.01	9.63	[0.44]
	—	(7.34)	(–6.54)	(–7.06)	—	(0.25)	(–4.54)	(1.53)	(2.61)	(–1.60)	(–0.45)	2	[0.05]
	0.21	1	0	0	–1	–0.01	0.09	0.01	–0.11	–0.17	0.10	[0.01]	[0.52]
	(6.21)	—	—	—	—	(–0.14)	(1.08)	(0.18)	(–3.82)	(–1.60)	(3.15)		[0.23]
7	1	1.47	–1.87	–1.16	0	0.02	–0.19	0.01	0.04	–0.08	–0.01	0.00	[0.40]
	—	(9.44)	(–6.33)	(–9.21)	—	(0.09)	(–4.68)	(0.78)	(3.03)	(–1.55)	(–0.85)	1	[0.09]
	0.17	1	0	–0.16	–0.77	0	0.01	–0.06	–0.18	–0.52	0.18	[0.96]	[0.47]
	(13.31)	—	—	(–4.36)	(–22.56)	—	(0.09)	(–1.16)	(–3.39)	(–2.82)	(3.17)		[0.16]
8	1	1.70	–2.84	–0.99	0	0.77	–0.12	0.01	0.03	–0.10	–0.01	3.43	[0.40]
	—	(8.07)	(–6.78)	(–5.75)	—	(2.30)	(–3.75)	(0.96)	(2.98)	(–2.65)	(–0.50)	1	[0.05]
	0	1	0	0.11	–0.92	0.30	0.16	–0.02	–0.10	–0.50	0.13	[0.06]	[0.68]
	—	—	—	(2.95)	(–22.46)	(10.79)	(1.40)	(–0.57)	(–2.38)	(–3.41)	(2.82)		[0.17]
9	1	1.41	–1	–1.41	0	–0.63	–0.22	–0.00	0.04	–0.02	–0.02	4.69	[0.41]
	—	(15.12)	—	(–15.12)	—	(–7.99)	(–5.01)	(–0.03)	(2.31)	(–0.40)	(–0.92)	3	[0.20]
	0.17	1	0	–0.18	–0.74	0	0.05	–0.07	–0.18	–0.57	0.19	[0.20]	[0.41]
	(14.42)	—	—	(–5.03)	(–22.97)	—	(0.32)	(–1.25)	(–3.18)	(–2.91)	(3.09)		[0.12]
10	1	1.48	–1.84	–1.17	0	–0.01	–0.19	0	0.04	–0.10	–0.02	1.48	[0.35]
	—	(9.46)	(–6.05)	(–9.36)	—	(–0.03)	(–4.69)	—	(2.80)	(–1.85)	(–1.19)	2	[0.09]
	0.17	1	0	–0.14	–0.79	0.02	0.01	0	–0.16	–0.44	0.21	[0.48]	[0.46]
	(3.21)	—	—	(–2.00)	(–17.09)	(0.17)	(0.09)	—	(–3.14)	(–2.44)	(3.77)		[0.15]
11	1	1.32	–1.03	–1.40	0	–0.58	–0.21	0.02	0.09	0	–0.06	4.36	[0.39]
	—	(8.91)	(–5.12)	(–12.18)	—	(–3.31)	(–3.62)	(1.10)	(4.24)	—	(–2.68)	2	[0.09]
	0.35	1	0	–0.49	–0.58	–0.28	0.06	–0.06	–0.24	0	0.27	[0.11]	[0.24]
	(7.40)	—	—	(–7.38)	(–11.35)	(–3.19)	(0.38)	(–0.95)	(–3.83)	—	(3.96)		[0.25]

Table 21 continued

No.[b]	Cointegration coefficients (and *t*-values)						Loading coefficients (and *t*-values) in equation …					Test for Restrictions[c]	Residual Tests[d]
	x	q_{NCU}	y^*	p^*_{NCU}	p_{NCU}	f	$\ldots\Delta x$	$\ldots\Delta q_{NCU}$	$\ldots\Delta y^*$	$\ldots\Delta p^*_{NCU}$	$\ldots\Delta p_{NCU}$	χ^2-statistic (DF) [SL]	L-B (25) LM(1) LM (4) Normality
12	1	1.47	–1.95	–1.11	0	0.06	–0.19	0.01	0.04	–0.12	0	7.21	[0.21]
	—	(9.43)	(–6.19)	(–8.81)	—	(0.24)	(–4.56)	(0.73)	(2.43)	(–2.30)	—	2	[0.06]
	0.17	1	0	–0.10	–0.79	–0.01	0.01	–0.12	–0.19	–0.69	0	[0.03]	[0.45]
	(2.83)	—	—	(–1.35)	(–15.43)	(–0.08)	(0.11)	(–2.60)	(–4.12)	(–4.28)	—		[0.17]
13	1	1.33	–0.99	–1.42	0	–0.61	–0.21	0	0.09	0	–0.08	5.50	[0.39]
	—	(8.96)	(–4.90)	(–12.32)	—	(–3.46)	(–3.62)	—	(3.96)	—	(–3.18)	4	[0.10]
	0.34	1	0	–0.47	–0.59	–0.27	0.06	0	–0.22	0	0.30	[0.24]	[0.24]
	(7.26)	—	—	(–7.14)	(–11.59)	(–3.03)	(0.35)	—	(–3.62)	—	(4.46)		[0.21]
14	1	1.42	–1	–1.42	0	–0.63	–0.22	0	0.04	–0.02	–0.02	6.04	[0.38]
	—	(12.70)	—	(–12.70)	—	(–3.65)	(–4.97)	—	(2.36)	(–0.37)	(–0.92)	5	[0.19]
	0.18	1	0	–0.17	–0.75	0	0.02	0	–0.16	–0.48	0.23	[0.30]	[0.36]
	(12.60)	—	—	(–7.08)	(–15.77)	—	(0.13)	—	(–2.91)	(–2.45)	(3.81)		[0.12]
15	1	2.74	–1	–2.74	0	0.04	–0.12	0.02	0.08	0	–0.09	10.15	[0.45]
	—	(13.15)	—	(–13.15)	—	(0.25)	(–1.97)	(1.11)	(3.68)	—	(–3.93)	5	[0.07]
	0.21	1	0	–0.62	–0.39	0	0.33	–0.11	–0.30	0	0.44	[0.07]	[0.38]
	(11.11)	—	—	(–11.14)	(–8.77)	—	(1.18)	(–1.15)	(–3.02)	—	(4.10)		[0.40]
16	1	1.80	–1	–1.80	0	–0.43	–0.15	0	0.06	0	–0.06	11.33	[0.40]
	—	(15.06)	—	(–15.06)	—	(–7.89)	(–3.49)	—	(3.51)	—	(–3.49)	7	[0.06]
	0.19	1	0	–0.35	–0.62	0	0.17	0	–0.17	0	0.35	[0.12]	[0.37]
	(14.58)	—	—	(–4.83)	(–23.53)	—	(0.92)	—	(–2.39)	—	(4.77)		[0.34]

Table 21 continued

[a]Under the restriction of cointegration rank 2, the table shows the transposed (6×2)-matrix of cointegrating vectors $\hat{\beta}'$ with elements $\hat{\beta}'_{ij}$ and the corresponding transposed (5×2)-matrix of loading coefficients $\hat{\alpha}'$ with elements $\hat{\alpha}'_{ij}$. The subscript i stands for the cointegrating vector with $i = 1$ the demand vector or its loading, $i = 2$ the supply vector or its loading, whereas j stands for the variable (as ordered in the table). – [b]The number in the first column corresponds to the following specifications:

1 – Just identified model: Domestic producer prices irrelevant for export demand ($\hat{\beta}'_{15} = 0$), foreign production irrelevant for export supply in the long run ($\hat{\beta}'_{23} = 0$);

2 – Constant returns to scale in foreign production ($\hat{\beta}'_{13} = -1$);

3 – "Real effective exchange rate"-hypothesis ($\hat{\beta}'_{12} = -\hat{\beta}'_{14}$);

4 – Demand for exports keeps pace with globalization ($\hat{\beta}'_{16} = -1$) (rejected);

5 – Relative profitability determines export supply ($\hat{\beta}'_{25} = -\hat{\beta}'_{22}$) (rejected);

6 – Simple markup-pricing by export suppliers ($\hat{\beta}'_{25} = -\hat{\beta}'_{22}$, $\hat{\beta}'_{24} = 0$) (rejected);

7 – Globalization has no influence on export prices and thus on export supply ($\hat{\beta}'_{26} = 0$);

8 – Long-run export supply is infinitely price-elastic ($\hat{\beta}'_{21} = 0$) (rejected);

9 – All not rejected restrictions on cointegration vectors (see 2, 3 and 7) together;

10 – In the just identified model (see 1), level of export prices is exogenous ($\hat{\alpha}'_{12} = \hat{\alpha}'_{22} = 0$);

11 – In the just identified model, foreign producer prices are exogenous ($\hat{\alpha}'_{14} = \hat{\alpha}'_{24} = 0$);

12 – In the just identified model, domestic producer prices are exogenous ($\hat{\alpha}'_{15} = \hat{\alpha}'_{25} = 0$) (rejected);

13 – In the just identified model, both export prices and foreign prices are exogenous (see 10 and 11);

14 – All not rejected restrictions on cointegration vectors plus exogeneity of export prices (2, 3, 7 and 10);

15 – All not rejected restrictions on cointegration vectors plus exogeneity of foreign prices (2, 3, 7 and 11) (rejected);

16 – All not rejected restrictions on cointegration vectors plus exogeneity of both export and foreign producer prices (2, 3, 7, 10 and 11).

– [c]The validity of the various restrictions is tested for with the classical χ^2-distributed likelihood ratio statistic with DF degrees of freedom and significance level SL (in square brackets) for the restriction(s) in this bloc. – [d]The last column contains significance levels of the following residual tests: L-B (25) is the Ljung-Box test for serial correlation of up to order T/4 (Ljung and Box 1978), LM (1) and LM (4) are tests for serial correlation of order one and up to order 4, respectively (e.g. Breusch 1988), and Normality is the multivariate normality test proposed by Doornik and Hansen (1994). As to residual heteroscedasticity (not reported), it is not found at the 5 percent significance level for any of the specifications shown according to univariate ARCH(2)-tests (Engle 1982) for each equation.

Table 22: Cointegration Relations and Loading Coefficients for Various Overidentifying Restrictions: Canada[a]

No.[b]	Cointegration coefficients (and *t*-values)						Loading coefficients (and *t*-values) in equation …				Test for Restrictions[c]	Residual Tests[d]
	x	q_{NCU}	y^*	p^*_{NCU}	p_{NCU}	f	$\ldots\Delta x$	$\ldots\Delta q_{NCU}$	$\ldots\Delta p^*_{NCU}$	$\ldots\Delta p_{NCU}$	χ^2-statistic (DF) [SL]	L-B (25) LM(1) LM (4) Normality
1	1	0.43	–0.59	–0.55	0	–1.06	–0.29	–0.06	–0.11	–0.06	—	[0.11]
	—	(2.34)	(–6.13)	(–3.89)	—	(–3.50)	(–4.29)	(–1.64)	(–1.76)	(–2.75)	—	[0.67]
	0.48	1	0	–0.23	–0.91	–0.57	–0.22	0.06	0.16	0.11	—	[0.51]
	(4.69)	—	—	(–2.47)	(–8.00)	(–3.33)	(–2.44)	(1.11)	(2.02)	(3.43)		[0.09]
2	1	0.39	–1	–0.53	0	–0.65	–0.30	–0.06	–0.10	–0.06	0.07	[0.12]
	—	(2.18)	—	(–3.87)	—	(–8.17)	(–4.43)	(–1.43)	(–1.71)	(–2.61)	1	[0.63]
	0.47	1	0	–0.22	–0.91	–0.54	–0.22	0.06	0.17	0.11	[0.79]	[0.51]
	(4.65)	—	—	(–2.48)	(–8.02)	(–3.25)	(–2.37)	(1.07)	(2.05)	(3.43)		[0.08]
3	1	0.72	–1.26	–0.72	0	–0.38	–0.21	–0.08	–0.09	–0.06	1.01	[0.08]
	—	(4.99)	(–6.66)	(–4.99)	—	(–2.08)	(–3.48)	(–2.49)	(–1.71)	(–2.89)	1	[0.69]
	0.57	1	0	–0.22	–0.97	–0.70	–0.24	0.06	0.14	0.10	[0.32]	[0.47]
	(5.27)	—	—	(–2.19)	(–8.12)	(–3.87)	(–2.75)	(1.29)	(1.89)	(3.35)		[0.24]
4	1	0.02	–0.73	–0.28	0	–1	–0.34	–0.00	–0.07	–0.02	2.57	[0.17]
	—	(0.10)	(–8.04)	(–1.88)	—	—	(–4.90)	(–0.08)	(–1.19)	(–0.91)	1	[0.33]
	0.28	1	0	–0.25	–0.79	–0.23	–0.28	0.03	0.16	0.10	[0.11]	[0.40]
	(3.07)	—	—	(–2.85)	(–7.17)	(–1.54)	(–2.80)	(0.46)	(1.75)	(2.61)		[0.04]
5	1	0.49	–1.07	–0.60	0	–0.56	–0.28	–0.07	–0.10	–0.06	0.21	[0.10]
	—	(2.72)	(–6.13)	(–4.28)	—	(–3.30)	(–4.26)	(–1.72)	(–1.66)	(–2.71)	1	[0.69]
	0.54	1	0	–0.18	–1	–0.66	–0.22	0.07	0.15	0.11	[0.64]	[0.47]
	(5.36)	—	—	(–3.65)	—	(–3.99)	(–2.55)	(1.38)	(1.92)	(3.57)		[0.14]
6	1	0.36	–0.77	–0.57	0	–0.82	–0.37	–0.03	–0.03	–0.02	4.21	[0.20]
	—	(2.49)	(–5.30)	(–4.97)	—	(–5.37)	(–4.86)	(–0.59)	(–0.46)	(–0.61)	2	[0.37]
	0.16	1	0	0	–1	–0.10	–0.16	0.11	0.12	0.09	[0.12]	[0.52]
	(2.44)	—	—	—	—	(–0.69)	(–1.80)	(0.11)	(1.43)	(2.77)		[0.04]

Table 22 continued

No.[b]	Cointegration coefficients (and t-values)						Loading coefficients (and t-values) in equation …				Test for Restrictions[c]	Residual Tests[d]
	x	q_{NCU}	y^*	p^*_{NCU}	p_{NCU}	f	$\ldots\Delta x$	$\ldots\Delta q_{NCU}$	$\ldots\Delta p^*_{NCU}$	$\ldots\Delta p_{NCU}$	χ^2-statistic (DF) [SL]	L-B (25) LM(1) LM (4) Normality
7	1	0.11	–0.80	–0.34	0	–0.90	–0.38	–0.03	–0.07	–0.03	2.78	[0.19]
	—	(0.66)	(–5.54)	(–2.65)	—	(–6.56)	(–4.93)	(–0.64)	(–0.96)	(–1.03)	1	[0.29]
	0.14	1	0	–0.24	–0.74	0	–0.25	–0.01	0.14	0.07	[0.10]	[0.49]
	(3.89)	—	—	(–2.53)	(–6.89)	—	(–2.52)	(–0.25)	(1.59)	(1.96)		[0.05]
8	1	–0.04	–0.66	–0.25	0	–1.04	–0.43	–0.02	–0.04	–0.01	4.21	[0.19]
	—	(–0.25)	(–5.21)	(–2.01)	—	(–7.62)	(–4.94)	(–0.32)	(–0.48)	(–0.29)	1	[0.22]
	0	1	0	–0.24	–0.67	0.21	–0.30	–0.03	0.12	0.06	[0.04]	[0.42]
	—	—	—	(–2.42)	(–6.28)	(3.01)	(–2.83)	(–0.47)	(1.29)	(1.56)		[0.06]
9	1	0.33	–1.00	–0.48	0	–0.67	–0.32	–0.05	–0.10	–0.06	0.51	[0.15]
	—	(1.84)	(–6.09)	(–3.51)	—	(–4.27)	(–4.53)	(–1.33)	(–1.63)	(–2.31)	1	[0.53]
	1	2.71	0	–0.66	–2.29	–1	–0.08	0.01	0.06	0.04	[0.48]	[0.55]
	—	(6.67)	—	(–2.65)	(–6.37)	—	(–2.30)	(0.60)	(2.05)	(2.99)		[0.06]
10	1	0.75	–1	–0.75	0	–0.64	–0.23	–0.07	–0.07	–0.04	3.61	[0.13]
	—	(5.94)	—	(–5.94)	—	(–7.16)	(–3.73)	(–2.06)	(–1.36)	(–2.02)	4	[0.50]
	0.42	1	0	–0.16	–1	–0.42	–0.20	0.04	0.15	0.08	[0.46]	[0.58]
	(5.44)	—	—	(–3.29)	—	(–5.44)	(–2.41)	(0.83)	(2.07)	(3.05)		[0.18]
11	1	0.15	–0.89	–0.38	0	–0.79	–0.36	0	–0.06	–0.03	1.82	[0.22]
	—	(0.83)	(–5.22)	(–2.72)	—	(–4.73)	(–5.05)	—	(–1.00)	(–1.34)	2	[0.43]
	0.33	1	0	–0.25	–0.80	–0.33	–0.23	0	0.14	0.09	[0.40]	[0.44]
	(3.41)	—	—	(–2.86)	(–7.51)	(–2.03)	(–2.26)	—	(1.59)	(2.57)		[0.03]
12	1	0.75	–1.12	–0.83	0	–0.41	–0.24	–0.04	0	–0.04	5.25	[0.07]
	—	(3.49)	(–5.53)	(–4.98)	—	(–2.06)	(–4.01)	(–1.27)	—	(–1.90)	2	[0.57]
	0.64	1	0	–0.07	–1.14	–0.88	–0.20	0.03	0	0.08	[0.07]	[0.22]
	(4.98)	—	—	(–0.65)	(–7.99)	(–4.08)	(–2.44)	(0.71)	—	(2.90)		[0.26]
13	1	–1.01	–0.33	0.44	0	–1.57	–0.79	–0.04	–0.02	0	6.69	[0.18]
	—	(–8.78)	(–3.47)	(3.89)	—	(–8.65)	(–4.74)	(–0.45)	(–0.14)	—	2	[0.17]
	–0.33	1	0	–0.45	–0.24	0.76	–1.11	–0.14	0.03	0	[0.04]	[0.32]
	(–5.51)	—	—	(–6.95)	(–3.69)	(5.41)	(–4.28)	(–0.90)	(0.12)	—		[0.19]

Table 22 continued

No.[b]	Cointegration coefficients (and *t*-values)						Loading coefficients (and *t*-values) in equation …				Test for Restrictions[c]	Residual Tests[d]
	x	q_{NCU}	y^*	p^*_{NCU}	p_{NCU}	f	$\ldots\Delta x$	$\ldots\Delta q_{NCU}$	$\ldots\Delta p^*_{NCU}$	$\ldots\Delta p_{NCU}$	χ^2-statistic (DF) [SL]	L-B (25) LM(1) LM (4) Normality
14	1	0.73	–1	–0.73	0	–0.66	–0.27	0	–0.01	–0.02	10.02	[0.16]
	—	(5.08)	—	(–5.08)	—	(–6.80)	(–4.35)	—	(–0.17)	(–0.75)	6	[0.22]
	0.40	1	0	–0.15	–1	–0.40	–0.27	0	0.11	0.08	[0.12]	[0.29]
	(5.18)	—	—	(–3.02)	—	(–5.18)	(–2.41)	—	(1.51)	(2.59)		[0.08]
15	1	0.99	–1	–0.99	0	–0.53	–0.19	–0.05	0	–0.03	10.35	[0.12]
	—	(6.94)	—	(–6.94)	—	(–5.25)	(–3.39)	(–1.61)	—	(–1.66)	6	[0.27]
	0.35	1	0	–0.09	–1	–0.35	–0.11	–0.03	0	0.08	[0.11]	[0.37]
	(4.48)	—	—	(–1.78)	—	(–4.48)	(–1.14)	(–0.53)	—	(2.56)		[0.08]
16	1	1	–1	–1	0	–0.52	–0.19	–0.05	0	–0.03	10.35	[0.12]
	—	—	—	—	—	(–6.78)	(–3.37)	(–1.63)	—	(–1.67)	7	[0.27]
	0.35	1	0	–0.09	–1	–0.35	–0.10	–0.03	0	0.08	[0.17]	[0.38]
	(4.51)	—	—	(–1.83)	—	(–4.51)	(–1.09)	(–0.53)	—	(2.57)		[0.08]
17	1	0	–0.73	–0.26	0	–1	–0.32	0.00	–0.08	–0.03	2.70	[0.16]
	—	—	(–9.29)	(–6.09)	—	—	(–4.77)	(0.01)	(–1.41)	(–1.14)	3	[0.37]
	1	2.83	0	–0.72	–2.31	–1	–0.10	0.02	0.06	0.04	[0.44]	[0.38]
	—	(7.29)	—	(–3.14)	(–6.63)	—	(–2.76)	(0.73)	(1.83)	(2.88)		[0.03]
18	1	0	–0.72	–0.27	0	–1	–0.32	0	–0.08	–0.03	3.13	[0.20]
	—	—	(–9.27)	(–6.30)	—	—	(–4.80)	—	(–1.35)	(–1.12)	5	[0.35]
	1	2.98	0	–0.76	–2.37	–1	–0.09	0	0.04	0.03	[0.68]	[0.36]
	—	(7.46)	—	(–3.26)	(–6.61)	—	(–2.51)	—	(1.40)	(2.46)		[0.03]
19	1	0	–0.67	–0.33	0	–1	–0.32	0	0	–0.01	10.58	[0.12]
	—	—	(–8.31)	(–6.94)	—	—	(–4.71)	—	—	(–0.51)	7	[0.19]
	1	3.54	0	–0.53	–3.06	–1	–0.07	0	0	0.03	[0.16]	[0.14]
	—	(6.93)	—	(–1.81)	(–6.79)	—	(–2.30)	—	—	(2.52)		[0.03]

Table 22 continued

[a]Unlike in the VECMs for the United States and Germany, here the level of foreign production, y^*, is considered as exogenous from the outset, i.e. $p = 4$, to obtain sensible results. Under the restriction of cointegration rank 2, the table shows the transposed (6×2)-matrix of cointegrating vectors $\hat{\beta}'$ with elements $\hat{\beta}'_{ij}$ and the corresponding transposed (4×2)-matrix of loading coefficients $\hat{\alpha}'$ with elements $\hat{\alpha}'_{ij}$. The subscript i stands for the cointegrating vector with $i = 1$ the demand vector or its loading, $i = 2$ the supply vector or its loading, whereas j stands for the variable (as ordered in the table). – [b]The number in the first column corresponds to the following specifications:

1 – Just identified model: Domestic producer prices irrelevant for export demand ($\hat{\beta}'_{15} = 0$), foreign production irrelevant for export supply in the long run ($\hat{\beta}'_{23} = 0$);

2 – Constant returns to scale in foreign production ($\hat{\beta}'_{13} = -1$);

3 – "Real effective exchange rate"-hypothesis ($\hat{\beta}'_{12} = -\hat{\beta}'_{14}$);

4 – Demand for exports keeps pace with globalization ($\hat{\beta}'_{16} = -1$);

5 – Relative profitability determines export supply ($\hat{\beta}'_{25} = -\hat{\beta}'_{22}$);

6 – Simple markup-pricing by export suppliers ($\hat{\beta}'_{25} = -\hat{\beta}'_{22}$, $\hat{\beta}'_{24} = 0$);

7 – Globalization has no influence on export prices and thus on export supply ($\hat{\beta}'_{26} = 0$) (rejected);

8 – Long-run export supply is infinitely price-elastic ($\hat{\beta}'_{21} = 0$) (rejected);

9 – Export supply keeps pace with globalization ($\hat{\beta}'_{21} = -\hat{\beta}'_{26}$);

10 – Best not rejected restrictions on cointegration vectors (see 2, 3, 5 and 9) together;

11 – In the just identified model (see 1), level of export prices is exogenous ($\hat{\alpha}'_{12} = \hat{\alpha}'_{22} = 0$);

12 – In the just identified model, foreign producer prices are exogenous ($\hat{\alpha}'_{13} = \hat{\alpha}'_{23} = 0$) (rejected);

13 – In the just identified model, domestic producer prices are exogenous ($\hat{\alpha}'_{14} = \hat{\alpha}'_{24} = 0$) (rejected);

14 – Best not rejected restrictions on cointegration vectors plus exogeneity of export prices (2, 3, 5, 9 and 11);

15 – Best not rejected restrictions on cointegration vectors plus exogeneity of foreign prices (2, 3, 5, 9 and 12);

16 – Like 15 plus unit-price elasticity of export demand ($\hat{\beta}'_{12} = -\hat{\beta}'_{14} = 1$); specification robust against dropping each of its overidentifying restrictions individually;

17 – "Alternative specification": Demand and supply keep pace with globalization ($\hat{\beta}'_{16} = 0$ and $\hat{\beta}'_{26} = 0$), no own-price elasticity of export demand ($\hat{\beta}'_{12} = 0$);

18 – "Alternative specification" (see 17) plus exogeneity of export prices;

19 – "Alternative specification" plus exogeneity of both export and foreign producer prices.

– [c]The validity of the various restrictions is tested for with the classical χ^2-distributed likelihood ratio statistic with DF degrees of freedom and significance level SL (in square brackets) for the restriction(s) in this bloc. – [d]The last column contains significance levels of the following residual tests: L-B (25) is the Ljung-Box test for serial correlation of up to order T/4 (Ljung and Box 1978), LM (1) and LM (4) are tests for serial correlation of order one and up to order 4, respectively (e.g. Breusch 1988), and Normality is the multivariate normality test proposed by Doornik and Hansen (1994). As to residual heteroscedasticity (not reported), it is generally not found at the 5 percent significance level according to univariate ARCH(2)-tests (Engle 1982). The only exception is equation Δp_{NCU} in specification 8.

Table 23: Cointegration Relations and Loading Coefficients for Various Overidentifying Restrictions: Germany[a]

No.[b]	Cointegration coefficients (and *t*-values)						Loading coefficients (and *t*-values) in equation …					Test for Restrictions[c]	Residual Tests[d]
	x	q_{NCU}	y^*	p^*_{NCU}	p_{NCU}	f	$\ldots\Delta x$	$\ldots\Delta q_{NCU}$	$\ldots\Delta y^*$	$\ldots\Delta p^*_{NCU}$	$\ldots\Delta p_{NCU}$	χ^2-statistic (DF) [SL]	L-B (25) LM(1) LM (4) Normality
1	1	0.37	–1.15	–0.78	0	–0.43	–0.31	–0.03	0.15	–0.04	0.00	—	[0.01]
	—	(2.13)	(–9.92)	(–7.37)	—	(–6.00)	(–3.18)	(–1.23)	(3.34)	(–0.35)	(0.09)	—	[0.09]
	–0.31	1	0	0.34	–0.99	0.29	0.29	–0.06	0.36	0.04	0.01	—	[0.52]
	(–8.46)	—	—	(5.46)	(–12.38)	(6.21)	(1.58)	(–1.49)	(4.37)	(0.21)	(0.09)		[0.15]
2	1	0.26	–1	–0.74	0	–0.51	–0.29	–0.03	0.17	–0.03	0.01	0.93	[0.01]
	—	(2.03)	—	(–8.04)	—	(–13.15)	(–2.70)	(–1.34)	(3.41)	(–0.27)	(0.15)	1	[0.10]
	–0.37	1	0	0.37	–0.96	0.36	0.25	–0.06	0.37	0.04	0.01	[0.34]	[0.50]
	(–12.83)	—	—	(5.46)	(–12.31)	(8.48)	(1.35)	(–1.49)	(4.36)	(0.18)	(0.09)		[0.14]
3	1	1.26	–1.57	–1.26	0	–0.30	–0.35	–0.02	0.10	–0.04	–0.00	16.72	[0.03]
	—	(13.03)	(–16.13)	(–13.03)	—	(–3.43)	(–4.20)	(–1.00)	(2.71)	(–0.49)	(–0.06)	1	[0.02]
	–0.23	1	0	0.63	–1.45	0.24	–0.00	–0.04	0.22	0.00	0.01	[0.00]	[0.60]
	(–4.24)	—	—	(6.51)	(–14.24)	(3.21)	(–0.02)	(–1.52)	(4.01)	(0.01)	(0.17)		[0.14]
4	1	0.11	–0.25	–0.75	0	–1	–1.29	–0.05	0.84	–0.18	–0.02	15.27	[0.01]
	—	(1.32)	(–6.02)	(–11.29)	—	—	(–2.62)	(–0.48)	(3.96)	(–0.35)	(–0.16)	1	[0.10]
	–3.87	1	0	2.96	–1.10	4.41	–0.24	–0.01	0.20	–0.04	–0.01	[0.00]	[0.50]
	(–44.53)	—	—	(10.22)	(–4.00)	(73.47)	(–2.09)	(–0.45)	(4.16)	(–0.30)	(–0.19)		[0.14]
5	1	0.38	–1.16	–0.79	0	–0.43	–0.31	–0.03	0.15	–0.04	0.00	0.01	[0.01]
	—	(2.95)	(–10.69)	(–9.86)	—	(–6.16)	(–3.22)	(–1.23)	(3.33)	(–0.36)	(0.09)	1	[0.09]
	–0.31	1	0	0.35	–1	0.29	0.28	–0.06	0.36	0.04	0.01	[0.92]	[0.53]
	(–8.69)	—	—	(11.73)	—	(6.23)	(1.56)	(–1.50)	(4.36)	(0.20)	(0.09)		[0.15]

Table 23 continued

No.[b]	Cointegration coefficients (and *t*-values)						Loading coefficients (and *t*-values) in equation …					Test for Restrictions[c]	Residual Tests[d]
	x	q_{NCU}	y^*	p^*_{NCU}	p_{NCU}	f	$\ldots\Delta x$	$\ldots\Delta q_{NCU}$	$\ldots\Delta y^*$	$\ldots\Delta p^*_{NCU}$	$\ldots\Delta p_{NCU}$	χ^2-statistic (DF) [SL]	L-B (25) LM(1) LM (4) Normality
6	1	0.36	–1.09	–0.78	0	–0.46	–0.44	0.01	–0.01	–0.05	–0.00	17.11	[0.01]
	—	(1.96)	(–8.43)	(–7.05)	—	(–5.80)	(–7.99)	(0.73)	(–0.20)	(–0.89)	(–0.04)	2	[0.46]
	0.26	1	0	0	–1	–0.28	–0.01	–0.04	–0.07	–0.01	0.01	[0.00]	[0.55]
	(4.11)	—	—	—	—	(–2.40)	(–0.14)	(–2.62)	(–2.06)	(–0.17)	(0.25)		[0.19]
7	1	0.42	–1.20	–0.80	0	–0.41	–0.44	–0.01	–0.01	–0.06	0.00	13.39	[0.01]
	—	(2.29)	(–9.36)	(–7.23)	—	(–5.26)	(–8.01)	(–0.47)	(–0.34)	(–0.93)	(0.14)	1	[0.21]
	0	1	0	0.30	–1.22	–0.01	0.01	–0.06	0.04	0.01	0.02	[0.00]	[0.63]
	—	—	—	(2.78)	(–9.37)	(–0.34)	(0.10)	(–2.46)	(0.66)	(0.12)	(0.69)		[0.39]
8	1	0.45	–1.24	–0.82	0	–0.38	–0.44	–0.01	0.00	–0.06	0.00	12.52	[0.01]
	—	(2.48)	(–9.93)	(–7.44)	—	(–5.04)	(–7.55)	(–0.79)	(0.01)	(–0.91)	(0.18)	1	[0.16]
	–0.04	1	0	0.29	–1.16	0	0.01	–0.07	0.07	–0.00	0.02	[0.00]	[0.60]
	(–1.65)	—	—	(3.24)	(–10.49)	—	(0.11)	(–2.51)	(1.06)	(–0.02)	(0.58)		[0.39]
9	1	0.39	–1.13	–0.79	0	–0.47	–0.33	–0.03	0.15	–0.04	0.00	0.59	[0.01]
	—	(2.24)	(–9.83)	(–7.45)	—	(–7.27)	(–3.29)	(–1.17)	(3.31)	(–0.33)	(0.10)	1	[0.07]
	–0.34	1	0	0.37	–1.01	0.34	0.23	–0.05	0.34	0.05	0.01	[0.44]	[0.51]
	(–8.69)	—	—	(5.42)	(–11.52)	(8.69)	(1.34)	(–1.41)	(4.35)	(0.25)	(0.13)		[0.08]
10	1	0.31	–1	–0.77	0	–0.53	–0.31	–0.03	0.17	–0.03	0.01	1.28	[0.01]
	—	(2.03)	—	(–8.04)	—	(–13.15)	(–2.87)	(–1.30)	(3.38)	(–0.28)	(0.15)	3	[0.09]
	–0.39	1	0	0.41	–1	0.39	0.20	–0.05	0.35	0.03	0.01	[0.73]	[0.53]
	(–12.71)	—	—	(5.46)	—	(12.71)	(1.15)	(–1.44)	(4.35)	(0.18)	(0.10)		[0.14]

Table 23 continued

No.[b]	Cointegration coefficients (and *t*-values)						Loading coefficients (and *t*-values) in equation …					Test for Restrictions[c]	Residual Tests[d]
	x	q_{NCU}	y^*	p^*_{NCU}	p_{NCU}	f	$\ldots\Delta x$	$\ldots\Delta q_{NCU}$	$\ldots\Delta y^*$	$\ldots\Delta p^*_{NCU}$	$\ldots\Delta p_{NCU}$	χ^2-statistic (DF) [SL]	L-B (25) LM (1) LM (4) Normality
11	1	0.38	–1.15	–0.80	0	–0.43	–0.30	–0.02	0.15	0	0.01	0.92	[0.02]
	—	(2.18)	(–9.72)	(–7.46)	—	(–5.90)	(–3.06)	(–1.03)	(3.36)	—	(0.26)	2	[0.09]
	–0.32	1	0	0.36	–1.00	0.30	0.28	–0.06	0.36	0	–0.00	[0.63]	[0.53]
	(–8.37)	—	—	(5.61)	(–12.17)	(6.15)	(1.57)	(–1.63)	(4.37)	—	(–0.03)		[0.19]
12	1	0.37	–1.15	–0.78	0	–0.43	–0.31	–0.03	0.15	–0.04	0	0.01	[0.01]
	—	(2.13)	(–9.92)	(–7.37)	—	(–6.00)	(–3.17)	(–1.28)	(3.35)	(–0.40)	—	2	[0.09]
	–0.31	1	0	0.35	–0.99	0.29	0.28	–0.06	0.36	0.03	0	[1.00]	[0.53]
	(–8.49)	—	—	(5.48)	(–12.22)	(6.10)	(1.57)	(–1.54)	(4.37)	(0.16)	—		[0.14]
13	1	0.37	–1.15	–0.78	0	–0.43	–0.29	0	0.17	0.03	0.02	1.47	[0.01]
	—	(2.12)	(–9.84)	(–7.36)	—	(–6.00)	(–3.02)	—	(3.82)	(0.29)	(0.63)	2	[0.07]
	–0.34	1	0	0.34	–0.98	0.32	0.32	0	0.39	0.19	0.04	[0.48]	[0.43]
	(–8.59)	—	—	(5.10)	(–11.52)	(6.32)	(1.85)	—	(4.97)	(0.99)	(0.76)		[0.15]
14	1	0.32	–1	–0.78	0	–0.53	–0.31	–0.03	0.17	0	0	2.72	[0.02]
	—	(3.41)	—	(–11.10)	—	(–22.13)	(–2.82)	(–1.24)	(3.39)	—	—	7	[0.08]
	–0.40	1	0	0.42	–1	0.40	0.20	–0.06	0.35	0	0	[0.91]	[0.55]
	(–12.87)	—	—	(14.31)	—	(12.87)	(1.14)	(–1.56)	(4.36)	—	—		[0.15]

Table 23 continued

[a]Under the restriction of cointegration rank 2, the table shows the transposed (6×2)-matrix of cointegrating vectors $\hat{\beta}'$ with elements $\hat{\beta}'_{ij}$ and the corresponding transposed (5×2)-matrix of loading coefficients $\hat{\alpha}'$ with elements $\hat{\alpha}'_{ij}$. The subscript i stands for the cointegrating vector with $i = 1$ the demand vector or its loading, $i = 2$ the supply vector or its loading, whereas j stands for the variable (as ordered in the table). – [b]The number in the first column corresponds to the following specifications:

1 – Just identified model: Domestic producer prices irrelevant for export demand ($\hat{\beta}'_{15} = 0$), foreign production irrelevant for export supply in the long run ($\hat{\beta}'_{23} = 0$);
2 – Constant returns to scale in foreign production ($\hat{\beta}'_{13} = -1$);
3 – "Real effective exchange rate"-hypothesis ($\hat{\beta}'_{12} = -\hat{\beta}'_{14}$) (rejected);
4 – Demand for exports keeps pace with globalization ($\hat{\beta}'_{16} = -1$) (rejected);
5 – Relative profitability determines export supply ($\hat{\beta}'_{25} = -\hat{\beta}'_{22}$);
6 – Simple markup-pricing by export suppliers ($\hat{\beta}'_{25} = -\hat{\beta}'_{22}$, $\hat{\beta}'_{24} = 0$) (rejected);
7 – Long-run export supply is infinitely price-elastic ($\hat{\beta}'_{21} = 0$) (rejected);
8 – Globalization has no influence on export prices and thus on export supply ($\hat{\beta}'_{26} = 0$) (rejected);
9 – Export supply keeps pace with globalization ($\hat{\beta}'_{21} = -\hat{\beta}'_{26}$);
10 – All not rejected restrictions on cointegration vectors (see 2, 6 and 9) together;
11 – In the just identified model (see 1), foreign producer prices are exogenous ($\hat{\alpha}'_{14} = \hat{\alpha}'_{24} = 0$);
12 – In the just identified model, domestic producer prices are exogenous ($\hat{\alpha}'_{15} = \hat{\alpha}'_{25} = 0$);
13 – In the just identified model, export prices are exogenous ($\hat{\alpha}'_{12} = \hat{\alpha}'_{22} = 0$);
14 – All not rejected restrictions on cointegration vectors plus exogeneity of domestic and foreign producer prices (2, 5, 6, 11 and 12).

– [c]The validity of the various restrictions is tested for with the classical χ^2-distributed likelihood ratio statistic with DF degrees of freedom and significance level SL (in square brackets) for the restriction(s) in this bloc. – [d]The last column contains significance levels of the following residual tests: L-B (25) is the Ljung-Box test for serial correlation of up to order T/4 (Ljung and Box 1978), LM (1) and LM (4) are tests for serial correlation of order one and up to order 4, respectively (e.g. Breusch 1988), and Normality is the multivariate normality test proposed by Doornik and Hansen (1994). As to residual heteroscedasticity (not reported), it is not found at the 5 percent significance level for any of the specifications shown according to univariate ARCH(2)-tests (Engle 1982) for each equation. At the 10 percent level, there is some evidence of heteroscedastic residuals only in the Δy^*-equation in a few specifications.

5.7. Alternative Identification and Specification

An oddity the U.S. results share with the Canadian ones is the falling long-run export supply curve. The various over-identifying restrictions have not been able to remove this feature. This is why I now explore two alternative specifications, the first one concerning the identification of the supply relationship, the second one concerning the whole model.

5.7.1. United States

As to the first alternative, I start with the VECM for the United States subject to $r = 2$ without any identifying restrictions. The cointegration vectors normalized to export quantities and prices read:

N°	x	q_{NCU}	y^*	p^*_{NCU}	p_{NCU}	f
1	1	–4.87	–4.69	–1.54	6.57	0.08
2	0.32	1	–0.36	–0.34	–0.55	–0.01

After imposing the restriction indentifying demand on the first vector, the system becomes:

N°	x	q_{NCU}	y^*	p^*_{NCU}	p_{NCU}	f
1	1	1.47	–1.85	–1.16	0	0.00
2	–0.08	1	0.65	0.16	–1.15	–0.01

In this unshaped form, the not-yet identified “supply” vector (n°2) exhibits a most welcome feature: it is horizontal or slightly upward-sloping (depending on the standard deviations not yet computed). This does not change when the demand vector (n°1) is brought into its final form already familiar from Table 21, blocs 9 and 16 by adding the CRS and REER hypotheses (LR test: $\chi^2(3) = 2.54, [0.47]$):

N°	x	q_{NCU}	y^*	p^*_{NCU}	p_{NCU}	f
1	1	1.38	–1	–1.38	0	–0.65
2	–0.17	1	0.78	0.29	–1.28	0.05

A most pragmatic idea now is to choose the identifying restriction for the second vector such as to "hurt" the data the least possible, which apparently implies putting a zero restriction on the globalization variable ($\beta_{62} = 0$, compare with Table 21, block 7) thereby obtaining (LR test: $\chi^2(2) = 2.54,[0.28]$):

N°	x	q_{NCU}	y^*	p^*_{NCU}	p_{NCU}	f
1	1	1.38	–1	–1.38	0	–0.65
2	–0.08	1	0.63	0.16	–1.15	**0**

As restrictions are now lying on each of the beta-vectors, the degrees of freedom equal the number of over-identifying restrictions (Johansen and Juselius 1994: 24), i.e. two. The standard error of the export volume coefficient in the supply vector is almost as high as the coefficient itself (0.06), but before giving up the assumption of a positive slope of the export supply curve, I impose two over-identifying restrictions on the producer price parameters making them compatible with both the simple-markup hypothesis in Dornbusch (1987) and with the concept of relative profitability of export supply advocated by Goldstein and Khan (1978). The system then becomes (LR test: $\chi^2(4) = 3.82,[0.43]$):

N°	x	q_{NCU}	y^*	p^*_{NCU}	p_{NCU}	f
1	1	1.41	–1	–1.41	0	–0.63
2	0.01 (0.03)	1	0.45	0	–1	0

The standard deviation of $\hat{\beta}_{21}$ is indicated in brackets and points to insignificance of the estimate. It is easily set to zero (LR test: $\chi^2(5) = 3.85,[0.57]$). When weak exogeneity restrictions on export and foreign producer prices are added, the final over-identified structure is obtained. Table 24 shows the result in block “3 USA” after reporting both the just-identified and the final over-identified models already discussed (blocks “1 USA” and “2 USA”). As one can see, the alternative identification leads to results much better accepted by the data.

However, the theoretical interpretation for the negative influence of growth in foreign production on export prices in the supply relationship is not straight-forward. A temptative explanation might consist in viewing y^* as a proxy for the transformation curve of the rest of the world and growth in foreign production as being driven by cost-saving innovations and technical progress. In the aggregate demand and supply scheme of the rest of the world, y^* would therefore represent outward shifts of the aggregate supply curve whereas p_{NCU} summarizes all monetary and demand-pull influences of producer price inflation and therefore represents shifts in the aggregate demand curve of the world. Both together determine the international price level of tradeables and thus the level of U.S. export prices. Yet, for the sake of stability of the export demand relationship, one has to assume that this technical progress represented by y^* is neutral in the sense that it does not save one of the production factors (U.S. exports or the bundle of all other factors) more than the other one. The hugest difference with respect to the VECM for Germany (Table 24, block “8 GER”) is that the price-depressing role is not played by the trade integration variable f, but by the foreign production variable. However, the two approaches exclude each other as they require two different underlying economic models, which makes international comparability more difficult. The results are only justified for the pragmatic reason that they are much better accepted by the data.

5.7.2. Canada

Is this approach available for the Canadian VECM, as well? The answer is “yes, but…”. To see this, the model is obtained by the following steps. Imposing the just-identifying restriction for demand on the first vector while leaving the second vector unrestricted gives:

N°	x	q_{NCU}	y^*	p^*_{NCU}	p_{NCU}	f
1	1	0.43	–1.06	–0.55	0	–0.59
2	0.03	1	0.61	0.04	–1.14	–0.37

Supply still does not slope upwards but the coefficient is close to zero. Completing the over-identification of demand by the CRS and REER hypotheses and adding $\hat{\beta}_{62} = 0$ as the alternative just-identifying restriction for the supply vector leads to (LR test: $\chi^2(2) = 2.66, [0.26]$):

N°	x	q_{NCU}	y^*	p^*_{NCU}	p_{NCU}	f
1	1	0.83	-1	-0.83	0.000	-0.58
2	-11.30	1	1.65	14.75	-19.42	**0**

The most "costly" over-identifying restriction in terms of the likelihood ratio test is the REER assumption. After withdrawing it and imposing Dornbusch's (1987) simple markup approach on the second vector by restrictions on p_{NCU} and p^*_{NCU}, one possible over-identified specification is (LR test: $\chi^2(3) = 0.42, [0.94]$)[64]:

N°	x	q_{NCU}	y^*	p^*_{NCU}	p_{NCU}	f
1	1	0.27	–1	–0.43	0	–0.69
2	–0.26 (–2.55)	1	0.74	0	–1	0

After setting export and foreign producer prices weakly exogenous[65], the specification reported in Table 24, block "6 CAN" is obtained, "4 CAN" and "5 CAN" being the just-identified and over-identified versions of the VECM

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[64] Export supply may also be assumed to be horizontal (LR test: $\chi^2(4) = 3.01, [0.56]$) although the *t*-value of the volume coefficient is quite high (–2.55).

[65] Setting q_{NCU} exogenous is again quite costly in terms of the LR test.

based on the identification $\beta_{32} = 0$ already known from sub-section 5.6. However, there exists one major problem with this specification besides the theoretical reservations already mentioned in the context of the U.S. VECM: the negative sign of the loading coefficient associated with the supply vector in equation Δx_t. Assume a supply disequilibrium in the form of excessive export prices, i.e. a point northwest of the equilibrium in the long-run export demand-supply scheme. Then as long as the demand curve is falling (which is the case), the only plausible adjustment along the demand curve involves rising export quantities. Therefore the loading $\hat{\alpha}_{21} < 0$ in block "6 CAN" is destabilizing.

Here is where the idea of a slightly different specification of the whole VECM for Canada kicks in. The change is so minor that all the theoretical considerations concerning export demand, export supply and the role of globalization remain in place. All I do is to replace the globalization variable f by a linear deterministic trend (t), which by assumption is restricted to the cointegration space ("model 4" in the notation by Hansen and Juselius (1995)) to rule out a quadratic trend in the logarithms of the time series. The measure should be costly in terms of the fit of the model as we know that globalization has not arrived at a constant speed. Yet this cost might be limited in this case as Canada's total exports are dominated by the relationship with one major industrialized partner country (the United States) so that the recent arrival of new trading partners in (South) East Asia and Central and Eastern Europe is not as relevant as in the models for the United States and Germany. Moreover, removing the one variable for which the assumption of difference-stationarity is difficult to establish can be a statistical advantage as it raises the chances for all time series used of having the same degree of integration. To obtain well-behaved residuals, two additional impulse dummies are needed which take the

Table 24: Synopsis of the Results of Tests of Long-Run Economic Hypotheses: United States, Canada, and Germany[a]

No.[b]	Identification	Cointegration coefficients (and *t*-values)							Loading coefficients (and *t*-values) in equation …					Test for Restrictions[c]	Residual Tests[d]
		x	q_{NCU}	y^*	p^*_{NCU}	p_{NCU}	f	t	$\ldots\Delta x$	$\ldots\Delta q_{NCU}$	$\ldots\Delta y^*$	$\ldots\Delta p^*_{NCU}$	$\ldots\Delta p_{NCU}$	χ^2-statistic (DF);[SL]	L-B (25) LM(1)(4) Normality
1	$\hat{\beta}'_{15}=0$	1	1.47	–1.85	–1.16	0	0.00	—	–0.19	0.01	0.05	–0.08	–0.01	—	[0.40]
USA		—	(9.40)	(–6.11)	(–9.23)	—	(0.02)	—	(–4.70)	(0.78)	(3.04)	(–1.52)	(–0.87)	—	[0.09]
	$\hat{\beta}'_{23}=0$	0.18	1	0	–0.16	–0.76	–0.01	—	0.01	–0.06	–0.18	–0.52	0.18	—	[0.46]
		(3.49)	—	—	(–2.47)	(–17.34)	(–0.09)	—	(0.07)	(–1.17)	(–3.42)	(–2.79)	(3.17)		[0.16]
2	$\hat{\beta}'_{15}=0$	1	1.80	–1	–1.80	0	–0.43	—	–0.15	0	0.06	0	–0.06	11.33	[0.40]
USA		—	(15.06)	—	(–15.06)	—	(–7.89)	—	(–3.49)	—	(3.51)	—	(–3.49)	7	[0.06]
	$\hat{\beta}'_{23}=0$	0.19	1	0	–0.35	–0.62	0	—	0.17	0	–0.17	0	0.35	[0.12]	[0.37]
		(14.58)	—	—	(–4.83)	(–23.53)	—	—	(0.92)	—	(–2.39)	—	(4.77)		[0.34]
3	$\hat{\beta}'_{15}=0$	1	1.53	–1	–1.53	0	–0.59	—	–0.17	0	0.02	0	0.02	7.55	[0.38]
USA		—	(13.56)	—	(–13.56)	—	(–6.13)	—	(–4.39)	—	(1.33)	—	(1.22)	9	[0.08]
	$\hat{\beta}'_{26}=0$	**0**	1	0.46	0	–1	0	—	0.05	0	–0.12	0	0.20	[0.58]	[0.29]
		—	—	(28.50)	—	—	—	—	(0.48)	—	(–3.02)	—	(4.68)		[0.23]
4	$\hat{\beta}'_{15}=0$	1	0.43	–0.59	–0.55	0	–1.06	—	–0.29	–0.06	—	–0.11	–0.06	—	[0.11]
CAN		—	(2.34)	(–6.13)	(–3.89)	—	(–3.50)	—	(–4.29)	(–1.64)	—	(–1.76)	(–2.75)	—	[0.67]
	$\hat{\beta}'_{23}=0$	0.48	1	0	–0.23	–0.91	–0.57	—	–0.22	0.06	—	0.16	0.11	—	[0.51]
		(4.69)	—	—	(–2.47)	(–8.00)	(–3.33)	—	(–2.44)	(1.11)	—	(2.02)	(3.43)		[0.09]
5	$\hat{\beta}'_{15}=0$	1	1	–1	–1	0	–0.52	—	–0.19	–0.05	—	0	–0.03	10.35	[0.12]
CAN		—	—	—	—	—	(–6.78)	—	(–3.37)	(–1.63)	—	—	(–1.67)	7	[0.27]
	$\hat{\beta}'_{23}=0$	0.35	1	0	–0.09	–1	–0.35	—	–0.10	–0.03	—	0	0.08	[0.17]	[0.38]
		(4.51)	—	—	(–1.83)	—	(–4.51)	—	(–1.09)	(–0.53)	—	—	(2.57)		[0.08]
6	$\hat{\beta}'_{15}=0$	1	0.25	–1	–0.42	—	–0.72	—	–0.49	0	—	0	0.01	9.41	[0.13]
CAN		—	(2.17)	—	(–4.64)	—	(–11.38)	—	(–5.50)	—	—	—	(0.23)	7	[0.21]
	$\hat{\beta}'_{26}=0$	–0.18	1	0.63	0	–1	0	—	–0.23	0	—	0	0.07	[0.22]	[0.15]
		(–1.71)	—	(3.14)	—	—	—	—	(–2.98)	—	—	—	(2.82)		[0.04]

Table 24 continued

No.[b]	Identifi-cation	Cointegration coefficients (and *t*-values)							Loading coefficients (and *t*-values) in equation …					Test for Restrictions[c]	Residual Tests[d]
		x	q_{NCU}	y^*	p^*_{NCU}	p_{NCU}	f	t	$\ldots\Delta x$	$\ldots\Delta q_{NCU}$	$\ldots\Delta y^*$	$\ldots\Delta p^*_{NCU}$	$\ldots\Delta p_{NCU}$	χ^2-statistic (DF);[SL]	L-B (25) LM(1)(4) Normality
7	$\hat{\beta}'_{15} = 0$	1	0.63	–1	–0.63	0	—	–0.007	–0.36	0	—	–0.02	–0.06	6.42	[0.38]
CAN		—	(3.86)	—	(–3.86)	—	—	(–7.00)	(–5.52)	—	—	(–0.29)	(–2.74)	5	[0.57]
	$\hat{\beta}'_{23} = 0$	**0**	1	0	–0.21	–0.87	—	0.003	0.36	0	—	0.18	0.15	[0.27]	[0.89]
		—	—	—	(–2.52)	(–9.40)	—	(>6.12)	(4.14)	—	—	(2.23)	(5.32)		[0.16]
8	$\hat{\beta}'_{15} = 0$	1	0.37	–1.15	–0.78	0	–0.43	—	–0.31	–0.03	0.15	–0.04	0.00	—	[0.01]
GER		—	(2.13)	(–9.92)	(–7.37)	—	(–6.00)	—	(–3.18)	(–1.23)	(3.34)	(–0.35)	(0.09)	—	[0.09]
	$\hat{\beta}'_{23} = 0$	–0.31	1	0	0.34	–0.99	0.29	—	0.29	–0.06	0.36	0.04	0.01	—	[0.52]
		(–8.46)	—	—	(5.46)	(–12.38)	(6.21)	—	(1.58)	(–1.49)	(4.37)	(0.21)	(0.09)		[0.15]
9	$\hat{\beta}'_{15} = 0$	1	0.32	–1	–0.78	0	–0.53	—	–0.31	–0.03	0.17	0	0	2.72	[0.02]
GER		—	(3.41)	—	(–11.10)	—	(–22.13)	—	(–2.82)	(–1.24)	(3.39)	—	—	7	[0.08]
	$\hat{\beta}'_{23} = 0$	–0.40	1	0	0.42	–1	0.40	—	0.20	–0.06	0.35	0	0	[0.91]	[0.55]
		(–12.87)	—	—	(14.31)	—	(12.87)	—	(1.14)	(–1.56)	(4.36)	—	—		[0.15]

[a]Table is organized as Tables 21 to 23. – [b]The blocks correspond to the following specifications:
1 USA – Just identified U.S. model (cf. Table 21, block 1);
2 USA – All not rejected restrictions on U.S. model (cf. Table 21, block 16);
3 USA – Overidentified U.S. model derived from the alternatively identified supply relationship ($\hat{\beta}'_{26} = 0$);
4 CAN – Just identified Canadian model (cf. Table 22, block 1);
5 CAN – All not rejected restriction on Canadian model (cf. Table 22, block 16);
6 CAN – Overidentified Canadian model derived from alternative supply identification ($\hat{\beta}'_{26} = 0$);
7 CAN – All not rejected restrictions on Canadian model with *t* instead of *f* and $\hat{\beta}'_{15} = \hat{\beta}'_{23} = 0$ as identifying restrictions;
8 GER – Just identified German model (cf. Table 23, block 1);
9 GER – All not rejected restrictions on German model (cf. Table 23, block 14). – [c]Cf. Table 21, footnote c. – [d]Cf. Table 21, footnote d.

values of zero in all quarters except for respectively 1980:2 (d802) and 1995:1 (d951), where they are 1. The rank test again yields $r=2$, the reduced rank regression is restricted to this outcome and demand and supply are identified as initially discussed, i.e. by the respective assumptions of no long-run influence of domestic producer prices on demand and no long-run influence of foreign industrial output on supply. In the initial model, the coefficient of export volumes in the supply relationship has still not the correct slope, but in this setting, the over-identifying restriction getting it horizontal ($\hat{\beta}_{21}=0$) is fairly well supported by the data (LR test: $\chi^2(1)=1.63,[0.20]$) and the cointegration vectors then read:

N°	x	q_{NCU}	y^*	p^*_{NCU}	p_{NCU}	t
1	1	0.82	–0.93	–0.66	0	–0.008
2	0	1	0	–0.20	–0.82	0.003

Adding both over-identifying restrictions of the demand vector known from block “5 CAN” (i.e.: CRS, REER) and assuming weak exogeneity of export prices (rather than foreign prices or exchange rates) gives the final model reported in block “7 CAN” of Table 24, which has not one single implausible sign and is altogether better accepted by the data than the version one block above. This version with the deterministic trend will serve as the point of reference in my further discussion.

A natural question is then to ask whether the linear trend should not replace f in the models for the other two countries, as well. However, apart from the argument that it is difficult to see why a a variable with a clear economic meaning (such as the trade intensity of world production) should be replaced by a purely statistical one (t), it works neither for the United States nor for Germany. This statement is independent of which one of the two supply identification strategies is followed ($\hat{\beta}_{32}=0$ or $\hat{\beta}_{62}=0$). In the U.S. VECM, using the linear trend combined with the first kind of supply identification (the irrelevance of foreign production) makes the long-run supply curve fall even more steeply, as does the second kind of identification (no influence of the time-trend on export prices), which in addition to that even produces irrealistic producer price elasticities. In the VECM for Germany, first the conventional identifying restrictions are implemented (i.e. $\hat{\beta}_{32} = 0$ for the supply vector). Then the supply curve is upward-sloping but the elasticities of export prices with respect to domestic and foreign producer prices amount to irrealistic numbers (5 and –3, respectively) putting the whole interpretation as an export supply relationship at risk. If, in turn, the alternative identification restriction is used (no influence of the time-trend on export prices), the producer price coefficients are plausible but the supply curve is falling and the over-identifying restriction of it being horizontal is clearly rejected by the data (LR test: $\chi^2(1) = 9.01, [0.00]$).

To sum up, the final specifications are those shown in blocks “3 USA”, “7 CAN” and “9 GER”. While the model already found in the previous section remains best for Germany, the final result for the United States departs with respect to the identifying restriction of the long-run supply relationship as a zero restriction on the globalization variable (suggested by the unrestricted vector) is used. For Canada, the results improve substantially when the globalization variable is replaced by a deterministic trend confined to the cointegration space while leaving the set of identifying restrictions unchanged ($\hat{\beta}_{51} = 0$ for demand, $\hat{\beta}_{32} = 0$ for supply).

6. Estimation of the Partial Models

Following the results of the tests for weak exogeneity obtained in the previous section, this section presents the partial models for the three countries. The partial model differs from the restricted VECM insofar as the weakly exogenous variables are not modelled any more. The number of equations in the VECM is reduced by the number of variables found weakly exogenous and the adjustment processes become less complex. As a consequence, the contemporaneous first differences of these variables can be integrated into the partial VECM (PVECM), which very often raises the explanatory power of the model for the remaining endogenous variables and improves the statistical properties of the VECM. This may stem from two reasons which are themselves desirable: first, problematic data features (e.g. outliers) in the suppressed error-correction equations do not weight on the system any more; second, the number of parameters in the short-run VECM is reduced substantially (Harris 1995: 98–99). In general there are $(n-p)\cdot(l-1)$ fewer coefficients to be determined (the –1 stems from the fact that $(n-p)$ contemporaneous first differences are added to the model). In my example, each PVECM contains three equations because $n=6$ and $p=3$. As the lag length is $l=2$ for Canada and Germany, $l=3$ for the United States, the reduction in parameters amounts to 6 for the United States and to 3 for the other two countries.

The estimates of the matrix of cointegration vectors β as well as the ones of the loading coefficients α remain unchanged compared to those of the respective over-identified final models of Table 24 (blocs 3 USA, 7 CAN, and 9 GER). Without presenting a thoroughly identified PVECM (as the short run will remain unidentified), it is also asked in this section which short-run coefficients are significant and which ones are not in order to prepare the elimination of insignificant short-run coefficients. This measure will make the model still more parsimonious and statistically more reliable.

In the PVECMs for the three countries, the principle of the most harmonized specification possible cannot be sustained any more. The three countries have only the number of equations and the endogeneity of export volumes in common, the latter feature being expected in a system designed to model exports. But as far as the remaining pairs of endogenous variables are concerned, they are different for each one of the three country pairs that can be formed from my sample of three countries. Specifically, export prices are endogenous only in the German case (Table 27). This fits well the idea of an upward-sloping supply curve with suppliers adjusting prices actively according to existing demand and supply pressures. For the United States (Table 25) and for Canada (Table 26), the weak exogeneity of export prices match the previously found horizontal long-run export supply curves and stress the idea that export quantities and two of the remaining endogenous variables react to supply and demand imbalances.

Table 25 presents the PVECM for the United States. Besides export quantities, foreign economic activity and American producer prices react to deviations from long-run export demand and/or supply whereas the effective dollar exchange rate and foreign producer prices (jointly expressed by p^*_{NCU}) do not react. First look at the loading coefficients. The aggregate volume of exports is the sole variable to adjust in case of a demand disequilibrium, the sign of the significant loading coefficient is negative as expected, i.e. exports shrink (rise) when exports have exceeded (remained below) the long-run level of demand in the quarter before. What about the loading coefficients belonging to the supply relationship that has been fixed at $q_{NCU,t-1} = p_{NCU,t-1} - 0.46 y^*_{t-1} + \hat{u}_{2,t-1}$ in the previous section? Imagine the disequilibrium takes the form of “too high” an export price level, i.e. the error correction term is positive. Then equilibrium can be restored either by an increase in domestic producer prices or by a decline in foreign activity. As the loading coefficient of the latter actually is significantly negative (–0.12) and the loading coefficient of the former significantly positive (0.20), the empirical finding is that usually both reactions take place and both help stabilizing the system.

Table 25: Results of the Partial Model: United States[a]

Equation	Δx	Δy^*	Δp_{NCU}
Loading Coefficients			
Loading coefficients of demand relation[b]	–0.17***	0.02	0.02
Loading coefficients supply relation[c]	0.05	–0.12***	0.20***
Coefficients of lagged endogenous variables			
Δx_{t-1}	–0.22***	0.08***	–0.04
Δx_{t-2}	–0.09	0.10***	–0.05*
Δy^*_{t-1}	1.10***	0.42***	0.06
Δy^*_{t-2}	0.20	–0.05	–0.12
$\Delta p_{NCU,t-1}$	–0.08	0.13	0.19**
$\Delta p_{NCU,t-2}$	0.22	0.19**	–0.05
Coefficients of exogenous variables			
$\Delta q_{NCU,t}$	–0.33	0.22**	0.53***
$\Delta q_{NCU,t-1}$	–0.45	0.15	–0.23**
$\Delta q_{NCU,t-2}$	–0.11	–0.11	0.09
$\Delta p^*_{NCU,t}$	0.11	–0.01	0.05**
$\Delta p^*_{NCU,t-1}$	0.02	0.02	–0.01
$\Delta p^*_{NCU,t-2}$	0.07	0.00	0.02
Δf_t	0.80	0.70*	0.34
Δf_{t-1}	0.35	–0.29	–0.32
Δf_{t-2}	–1.17	0.25	–0.04
Coefficients of dummy variables			
$d782$	0.08***	0.02**	0.01
$d802$	–0.02	–0.04***	0.00
Constant	–0.39*	0.28***	–0.38***
Residual analysis			
Standard deviation (in percent)	1.60	0.59	0.54
ARCH (3)[d]	1.46	0.90	1.77
Normality[e]	3.80	2.47	3.75
R^2	0.51	0.62	0.73
Multivariate tests[f]	LB[0.36], LM(1)[0.07], LM(4)[0.19], N[0.05]		

*(**,***) denote significance at the 10 (5, 1) percent level. – [a]According to the results from Table 24, block 3 USA, export prices, foreign producer prices and globalization are set exogenous and the partial VECM is estimated under the same restrictions on the cointegrating vectors. The latter turn out to be numerically the same. The restrictions are not rejected by the data according to the LR test ($\chi^2(5) = 2.05,[0.84]$) – [b]The loading coefficient applies to the lagged "long-run" disequilibrium $\hat{u}_{1,t-1}$ in $x_{t-1} = y^*_{t-1} - 1.53\left(q_{NCU,t-1} - p^*_{NCU,t-1}\right) + 0.59 f_{t-1} + \hat{u}_{1,t-1}$. – [c]The loading coefficient applies to the lagged "long-run" disequilibrium $\hat{u}_{2,t-1}$ in $q_{NCU,t-1} = p_{NCU,t-1} - 0.46 y^*_{t-1} + \hat{u}_{2,t-1}$. – [d]Critical values of the $\chi^2(3)$-distribution apply (*:6.25; **:7.82; ***:9.35). – [e]Critical values of the $\chi^2(2)$-distribution apply (*:4.61; **:5.99; ***:7.38). – [f]Ljung-Box, LM (1) and LM(4) tests for autocorrelation and test for normality by Doornik and Hansen. Probabilities in square brackets.

As to the short-run coefficients, many of them are insignificant. In the first PVECM equation which determines the quarter-on-quarter rate of change in real exports (second column of Table 25), apart from the constant and the dummy only the once-lagged coefficients of the first difference of exports themselves and of foreign industrial production are significant. The 1.10 coefficient of the latter suggests an about one-to-one reaction of exports. That means that the long-run reaction to a permanent 1 percent increase in foreign production is basically achieved after one quarter. However, the system will not be motionless after one quarter as each change in exports triggers a slight counter-reaction one quarter later due to the negative coefficient of the lagged left-hand side variable (Δx_{t-1}). In the equation of foreign industrial production growth, four significant parameters of once-differenced I(1) variables show up (column 3): increases in American exports are followed by a less pronounced movement in the same direction of foreign industrial output.[66] Next two positive contemporary movements of y_t^* are noteworthy: the one with a change in Δf_t and the one with a change in $\Delta q_{NCU,t}$. These lend support to the familiar procyclical nature of international trade volumes and prices. Furthermore, note the positive AR(1) parameter (0.42).

The positive AR(1) term in the foreign production equation compares well to the one in the equation of domestic producer price inflation (0.19 in column 4). Unlike the negative parameter for Δx_{t-1} in the export growth equation (–0.22), both these positive AR(1) terms *per se* accelerate the return into equilibrium after an export supply disturbance. To see this, again refer to the above example of “too high” export prices in period $(t-1)$. The pressure on foreign production and domestic prices declenched by the stabilizing loading coefficients in period t is reinforced in $(t+1)$ by the action of the positive AR(1) terms. Last but not

66 All these short-run reactions are discussed in terms of contemporary versus lagged movements rather than in terms of causality as the short-run part of the model remains unidentified.

least, there are positive contemporaneous price coefficients in the equation for domestic producer price inflation suggesting a co-movement between American prices set on the U.S. market and those set abroad and a weak but positive relationship between foreign prices in U.S. dollar and the domestic PPI.

The PVECM for Canada is presented in Table 26. It reveals both similarities and differences compared with the U.S. model. As in the U.S. case, export quantities and domestic producer prices are endogenous but the third endogenous variable is foreign producer prices in Canadian dollars pointing to endogenous adjustments in the exchange rate and/or in foreign producer prices. This variable (p^*_{NCU}) endogenously adjusts after a deviation of the long-run supply relationship from equilibrium as the level of foreign (i.e.: U.S.) prices is part of the supply vector. Taking the already familiar example of "too high" an export price level, i.e. a positive residual in the structural supply relationship $q_{NCU,t-1} = 0.88 p_{NCU,t-1} + 0.21 p^*_{NCU,t-1} - 0.003t + \hat{u}_{2,t-1}$, the supply equilibrium may be restored by increases in both Canadian and U.S. producer prices and this is what usually seems to happen according to the significantly positive loading coefficients of p^*_{NCU} and Δp_{NCU} (0.18 and 0.15), respectively. A more striking difference from the U.S. PVECM is that export quantities react to supply disequilibria although the supply curve is horizontal. The loading coefficient is positive and quite substantial (0.36). Whether it is stabilizing or destabilizing depends on the specific situation.[67]

Turning to a demand disequilibrium, an excess level of exports is corrected by a decrease in exports themselves, which is about twice as strong as in the U.S. case given the higher loading coefficient (–0.36 versus –0.17), and by a slight decrease in Canadian producer prices. The latter reaction neither widens nor

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[67] In the example of "too high" an export price level, the resulting increase in export volumes is stabilizing as long as the price disequilibrium is matched by a lack in export quantities (points northwest of the intersection between the supply and demand curve). For instance, this holds for all "excessive-price" points *on* the demand curve.

narrows the export gap but implies an export price gap (a rise in $\hat{u}_{2,t-1}$) tending to bring Δp_{NCU} back to the initial level.

The short-run part is again quite parsimonious when only the significant parameters are looked at. In the export growth equation (column 2), two contemporaneous effects from exogenous variables stand out: the negative export price coefficient (–0.48) and the positive coefficient of U.S. industrial production (0.98). Both quicken very much the necessary adjustment to a new equilibrium. When U.S. production grows by 1 percent, so do Canadian exports in the long run but this shift happens at the same time. When export prices rise by 1 percent, export demand shrinks, *ceteris paribus*, by 0.63 percent in the long run but a 0.48 percent fall in export volumes is already observed contemporaneously. The domestic producer price equation (column 4) has more significant coefficients. As in the U.S. case, there is a co-movement between export prices and domestic producer prices, furthermore current increases in U.S. production seem to boost prices but this effect is reversed a quarter later. Unlike the export growth equation, the short-run determination of Δp_{NCU} also depends on lagged endogenous variables. The AR(1) term (0.19) is the same as in the U.S. PVECM and helps restoring the supply equilibrium after a disturbance. The coefficient of lagged exports (0.06) suggests that an export boom implies higher domestic capacity utilization and higher producer prices, whereas the negative coefficient on foreign prices (–0.06) is not immediately seized by intuition.

The counterpart of this puzzle is the –0.76 coefficient of lagged domestic PPI inflation on foreign prices in Canadian dollars in the Δp^*_{NCU} equation (column 3) suggesting a rise in domestic production prices might be followed by a fall in U.S. prices or a nominal appreciation of the Canadian dollar. Yet the positive AR(1) parameter (0.33) in the Δp^*_{NCU} equation is plausible and fits into

Table 26: Results of the Partial Model: Canada[a]

Equation	Δx	Δp^*_{NCU}	Δp_{NCU}
Loading Coefficients			
Loading coefficients of demand relation[b]	–0.36***	–0.02	–0.06***
Loading coefficients of supply relation[c]	0.36***	0.18***	0.15***
Coefficients of lagged endogenous variables			
Δx_{t-1}	0.01	0.04	0.06***
$\Delta p^*_{NCU,t-1}$	–0.14	0.33***	–0.06*
$\Delta p_{NCU,t-1}$	–0.37	–0.76***	0.19**
Coefficients of exogenous variables			
$\Delta q_{NCU,t}$	–0.48***	0.95***	0.36***
$\Delta q_{NCU,t-1}$	0.25	–0.19	0.09
Δy^*_t	0.98***	0.03	0.11**
Δy^*_{t-1}	0.16	–0.20	–0.14***
Coefficients of dummy variables			
$d802$	–0.02	0.02	–0.02***
$d824$	–0.07***	–0.02*	–0.00*
$d951$	0.04**	0.02	0.02***
Constant	–0.10***	–0.03	0.00
Residual analysis			
Standard deviation (in percent)	1.82	1.32	0.46
ARCH (2)[d]	1.86	1.13	4.89*
Normality[d]	1.45	2.81	3.54
R^2	0.54	0.56	0.84
Multivariate tests[e]	LB[0.47], LM(1)[0.68], LM(4)[0.83], N[0.23]		

*(**,***) denote significance at the 10 (5, 1) percent level. – [a]According to the results from Table 24, block 7 CAN, export prices and foreign industrial production are set exogenous and the partial VECM is estimated under the same restrictions on the cointegrating vectors. The latter turn out to be numerically the same. The restrictions are not rejected by the data according to the LR test ($\chi^2(3) = 5.77, [0.12]$). – [b]The loading coefficient applies to the lagged "long-run" disequilibrium $\hat{u}_{1,t-1}$ in $x_{t-1} = y^*_{t-1} - 0.63\left(q_{NCU,t-1} - p^*_{NCU,t-1}\right) + 0.007t + \hat{u}_{1,t-1}$. – [c]The loading coefficient applies to the lagged "long-run" disequilibrium $\hat{u}_{2,t-1}$ in $q_{NCU,t-1} = 0.88 p_{NCU,t-1} + 0.21 p^*_{NCU,t-1} - 0.003t + \hat{u}_{2,t-1}$. – [d]Critical values of the $\chi^2(2)$-distribution apply (*:4.61; **:5.99; ***:7.38). – [e]Ljung-Box, LM (1), LM(4) tests for autocorrelation and test for normality by Doornik and Hansen. Probabilities in square brackets.

the above considerations on the correction of long-run supply disequilibria. Finally, the contemporaneous coefficient of export prices is interesting as it is very close to one (0.95). This can mean that a 1 percent depreciation of the Canadian dollar or a producer price increase in the United States (both implying a 1 percent rise in p^*_{NCU}) coincides with an increase in Canadian export prices of the same size. This is perfectly in line with the idea of Canadian export prices being determined on the U.S. market (perhaps even invoiced in U.S. dollars) and that the domestic-currency level of the export deflator adjusts almost automatically to changes in this predominant foreign market.

Finally the PVECM for Germany remains to be analyzed (Table 27). The long-run demand curve was found to be downward-sloping in the previous section, the long-run supply curve upward-sloping, foreign industrial production acts as a shift variable of the demand curve only and changes in domestic producer prices exclusively shift the supply curve. To study what happens in case of a deviation from long-run levels of export volumes and prices, the reader is well advised to refer to Figure 5 in the previous section. First take a demand disequilibrium. Let it be characterized by an excess volume of exports, i.e. by a point off the demand but on the supply curve northeast of the intersection of the two curves ($\hat{u}_{1,t-1} > 0$). The loading coefficients of the demand relationship read –0.31, –0.03, and 0.17. These values are stabilizing the system as they imply a strong reduction in export quantities (movement to the west) as well as a slight but significant fall in export prices (movement to the south). At the same time, the rise in y^*_t triggers an outward shift of the demand curve moving northeast the equilibrium itself and bringing it closer to the initial point.

Next assume the disequilibrium happens to be on the supply side and the initial export price-quantity data point is situated on the demand curve northwest of the intersection, which implies too high an export price level and too low exports quantities ($\hat{u}_{2,t-1} > 0$). Now the relevant loading coefficients are 0.20, –0.06, and 0.35. The first of them has the correct sign as it helps increasing real exports but is insignificant. But the negative loading coefficient in the export price equation clearly puts downward pressure on the export price level (movement to the south). Moreover, foreign production rises provoking an outward shift of the demand curve. The latter has two desirable implications: first, a given quantity of exports can now be sold at a higher price thereby partly justifying the initially “too high” export price level; second, the increase in y_t^* brings demand into disequilibrium opening up a positive export gap followed by higher exports in period $(t+1)$, i.e. the quantity movement (to the east) finally starts with a delay of one quarter.

What about the short-run coefficients? As to the export growth equation (column 2), two of the three lagged endogenous coefficients are significant. Like for the other countries, there is an about one-to-one short-run response to changes in foreign production. The AR(1) term, in turn, is insignificant (like for the Canadian and unlike the U.S. model). Curiously, there is a sharp negative response to a change in export prices which is striking not by its sign but by its strength (–2.74) compared to the rather weak long-run response (–0.32). Although this effect might be somewhat counterbalanced by the opposite-signed coefficient on the lagged domestic producer price change (1.42) in situations when export prices and producer prices have moved into the same direction in $(t-1)$, it suggests that the adjustment path after exogenous domestic price shocks or export demand and supply shocks is far more complicated than can be understood with a *ceteris paribus* discussion of coefficients. However, the result is indicate that a cost push (resulting, for example, from an oil price like or from a strong increase in negotiated wages per hour worked) produce an overshooting negative impact on Germany’s export volumes in the short run.

As to the other two equations, the positive AR (1) terms (0.18 in column 3, 0.42 in column 4) *per se* quicken export demand and supply adjustments and remind very much the results for the United States and Canada. Moreover, the four highly significant positive coefficients (0.11, 0.08, 0.06, 0.27) in the export price equation (column 3) point to a positive but far from perfect co-movement between domestic and export prices and back the assumption of foreign trade prices being procyclical.

Table 27: Results of the Partial Model: Germany[a]

Equation	Δx	Δq_{NCU}	Δy^*
Loading Coefficients			
Loading coefficients of demand relation[b]	–0.31***	–0.03*	0.17***
Loading coefficients of supply relation[c]	0.20	–0.06**	0.35***
Coefficients of lagged endogenous variables			
Δx_{t-1}	–0.12	0.00	–0.01
$\Delta q_{NCU,t-1}$	–2.74***	0.18*	–0.06
Δy^*_{t-1}	1.11***	0.11***	0.42***
Coefficients of exogenous variables			
$\Delta p^*_{NCU,t}$	0.03	0.08***	–0.03
$\Delta p^*_{NCU,t-1}$	0.18	0.06***	0.08
$\Delta p_{NCU,t}$	0.49	0.27***	0.08
$\Delta p_{NCU,t-1}$	1.42***	–0.00	–0.06
Δf_t	–0.24	–0.05	1.18***
Δf_{t-1}	1.14	–0.15	–1.05***
Coefficients of dummy variables			
$d801$	0.07***	0.00	0.01
$d842$	–0.05***	–0.00	–0.00
$d903$	0.09***	–0.01***	–0.00
$d904$	0.08***	0.00	–0.00
Constant	–1.35***	0.03	0.20*
Residual analysis			
Standard deviation (in percent)	1.40	0.22	0.65
ARCH (2)[d]	0.92	4.37	4.41
Normality[d]	0.71	6.83**	10.97***
R^2	0.73	0.86	0.46
Multivariate tests[e]	LB[0.32], LM(1)[0.07], LM(4)[0.39], N[0.31]		

*(**,***) denote significance at the 10 (5, 1) percent level. – [a]According to the results from Table 24, block 9 GER, foreign and domestic producer prices as well as globalization are set exogenous and the partial VECM is estimated under the same restrictions on the cointegrating vectors. The latter turn out to be numerically the same. The restrictions are not rejected by the data according to the LR test ($\chi^2(3)=1.41,[0.70]$). – [b]The loading coefficient applies to the lagged "long-run" disequilibrium $\hat{u}_{1,t-1}$ in $x_{t-1} = y^*_{t-1} - 0.32q_{NCU,t-1} + 0.78p^*_{NCU,t-1} + 0.53f_{t-1} + \hat{u}_{1,t-1}$. – [c]The loading coefficient applies to the lagged "long-run" disequilibrium $\hat{u}_{2,t-1}$ in $q_{NCU,t-1} = 0.40x_{t-1} + p_{NCU,t-1} - 0.42p^*_{NCU,t-1} - 0.40f_{t-1} + \hat{u}_{2,t-1}$. – [d]Critical values of the $\chi^2(2)$-distribution apply (*:4.61; **:5.99; ***:7.38). – [e]Ljung-Box, LM (1) and LM(4) tests for autocorrelation and test for normality by Doornik and Hansen. Probabilities in square brackets.

When comparing the results for the United States, Canada and Germany, one can observe that the adjustment processes after demand disturbances are of similar nature because the reaction in export quantities is the predominant (in case of the U.S. even the only) error correction mechanism. The smaller the country the stronger the initial percentage change in exports seems to be as suggested by the size of the loading coefficients. The correction of export supply disequilibria is much more heterogenous across countries which is certainly related to the differences between the long-run supply relationships identified (horizontal versus upward-sloping supply curves, positive, negative or no correlation between foreign producer prices and export prices, constant markup on domestic production costs versus more complex long-run ties between the export deflator, domestic and foreign producer prices and the exchange rate). A common feature for the two bigger countries (United States and Germany) is the feedback of export disequilibria on foreign industrial production. The smallest country, Canada, is the only one where export volumes also respond to a supply disturbance. All coefficients of lagged left-hand variables except one[68] are either insignificant or positive which *per se* allows for a quicker error correction than without such influences. However, to describe the differences between the three countries more precisely, a thorough analysis of the impulse-response pattern is indispensable and represents the most natural extension to this paper. The most interesting question that will be answered then is the one of how long it takes to absorb a supply disequilibrium, a demand disequilibrium, and shocks to the weakly exogenous variables.

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[68] The exception is the AR(1) term in the U.S. export growth equation (–0.22 in Table 25, column 2).

7. Summary Remarks

This study addressed two weaknesses commonly found in the empirical literature on aggregate exports: the exclusive focus on the demand side and the neglect of the trade-accelerating effects of the globalization process. The supply side was integrated into the analysis by taking into account potential output and the level of producer prices of the exporting countries. The increase in the international division of labor was modeled by introducing a new variable, the "trade intensity of world production", which was expected to have a positive long-run influence on export volumes and a negative impact on export prices. After leaving out potential GDP, which was found to play no role in the long-run determination of export supply and produced inferior results, a vector error correction model (VECM) was set up using logarithms of the following six variables: the volume of aggregate exports, the price level of aggregate exports, a trade-weighted index of industrial production in major trading partner countries, a trade-weighted index of foreign producer prices "translated" into domestic currency units using bilateral exchange rates, the index of producer prices in the domestic economy, and the trade intensity of world production, proxied by the ratio of real world merchandise exports to real world GDP. All steps of the analysis were carried out for three countries: the United States, Canada, and Germany.

Section 4 provided an extensive discussion of the properties of these time series. Both univariate and multivariate models to analyze the degree of integration are presented. While the latter produced mixed results, the univariate unit-root tests backed the assumption that the data are difference-stationarity for an overwhelming majority of the time series, the ADF test even more so than the KPSS test for stationarity. Only the globalization variable is close to I(2).

Although an error margin remains concerning the degree of integration, the VECM appears to be the appropriate framework to analyze the long-run economic equilibrium relationships as well as the short-run adjustments of export demand and supply.

In section 5, the I(1) analysis of the VECM was carried out using the Johansen procedure. Two rank tests as well as an analysis of the roots near the companion matrix came to the conclusion that each six-variable VECM contains two cointegrating vectors which can be identified as long-run relationships of aggregate export demand and supply, respectively. Zero restrictions were used to achieve this identification: for the demand vector, it was assumed that domestic producer prices are irrelevant whereas the supply vector was identified by restricting the long-run influence of foreign industrial production to zero. Then over-identifying restrictions were tested to see whether long-run predictions by economic models are supported by the data.

On the demand side, it was tested whether the foreign-production elasticity of exports is one (CRS), whether demand reacts with the same strength to changes in the price level of exports and to changes of foreign prices and the exchange rate, making the real effective exchange rate the appropriate price measure in models of export demand (REER), and whether a 1 percent increase in world trade intensity leads to a 1 percent increase in U.S., Canadian and German exports, respectively, i.e. whether the world market share of the three countries studied is constant in real terms over the long run (CMS). Whereas the the first two hypotheses were accepted (except for REER in the German case), the coefficient of the globalization variable was found to be significantly smaller than one and quite similar for all three countries (close to 0.5) reflecting declining shares in world exports over time. The advantage of taking a globalization proxy explicitely into account is the possibility of restricting the long-run influence of foreign industrial production to one, a restriction both required by the theoretical model (deriving export demand from a foreign production function with constant returns to scale) and accepted by the data. However, the three countries are very different as far as the price elasticity of export demand is concerned. Germany has the lowest own-price elasticity with around 0.3,[69] Canada's export demand is also quite inelastic (0.6) whereas the price elasticity of export demand amounts to about 1.5 for the United States possibly reflecting the higher share of consumer goods in total exports.

Results for the supply side turned out to be more heterogenous across countries and less easy to model. The identification strategy discussed above yielded sensible results for Germany whose long-run export supply curve is not infinitely price-elastic (as generally assumed in the literature) but upward-sloping. The same identification strategy produced slightly (but significantly) downward-sloping supply curves when applied to the United States and to Canada, an unsatisfying feature that did not disappear whatever over-identifying restriction was added to the model. Therefore the identifying restriction and the model specification itself were questioned and alternatives were searched for. In case of the United States, an alternative identifying restriction, the irrelevance of globalization for the level export prices, was chosen, then the over-identifying restriction of a horizontal long-run supply curve was very well supported by the data. For Canada, the same identifying restrictions as for Germany were chosen but the globalization variable was replaced by a linear deterministic trend confined to the cointegration space, which also allowed for a horizontal export supply curve in the long run. Based on these quite different basic models, the same set of over-identifying restrictions was tested for all three countries. Most importantly, a linear homogeneity restriction on the coefficients of exportprices and domestic producer prices was tested to see whether "relative profitability"

69 However, German exports are more responsive to changes in the effective exchange rate in the long run (0.8).

of the exporting activity (the ratio between export and producer prices) is the relevant price variable on the supply side as argued by Goldstein and Khan (1978). This is the case for Germany and the United States. In the U.S. model it was even possible to restrict the influence of foreign prices and exchange rates to zero thus allowing for an interpretation of export prices as a constant markup on production costs discussed in the Dornbusch (1987) model. In the Canadian export price formation, however, there seems to be a variable markup as the coefficients of domestic and foreign producer costs add up to approximately one. The countries also differ as to the relevance of globalization in the empirical long-run export supply relationship: Whereas the expected negative influence on export prices was found for Germany, the assumption of no influence was best supported by the U.S. data. In the preferred Canadian VECM where the globalization proxy was replaced by the linear trend, the finding is qualitatively the same as for Germany as there is a negative time-trend in the long-run cointegration relationship normalized on export prices and identified as export supply. Interestingly, there is no single testable restriction that was accepted or rejected for all three countries.

The United States, Canada and Germany also differ as to the question which variables are involved in the error correction mechanism. The three VECMs only have in common the number of endogenous variables (3 out of the 6) as well as weak exogeneity of the globalization variable, a feature *imposed* on the VECMs rather than found to hold in the data, however. The countries with horizontal long-run supply curves, i.e. the United States and Canada, are characterized by weak exogeneity of export prices. The third weakly exogenous variable is foreign production in the Canadian model and the foreign price level in U.S. dollar in the U.S. model. For Germany, in turn, the framework presented allowed for a simultaneous determination of export volumes and prices according to the initial motivation of this paper. The third endogenous variable is foreign industrial production. Just as for the United States and probably due to the size of the German economy compared to its most important trading partners, there is a feedback from exports on foreign economic activity.

Finally, in section 6, I exploited the results of the tests for weak exogeneity and presented the partial models (PVECM) which also incorporated the non-rejected over-identifying long-run hypotheses on the cointegrating vectors. The dynamic adjustment behavior of the models was discussed by analyzing the error correction mechanisms with the help of the sign of the loading coefficients. A common feature in all models is that export volumes adjust to demand disequilibria, the stronger so the smaller the country. At the same time, there is only one case (Canada) in which export volumes also react to supply disequilibria and only one case (Germany) in which export prices respond endogenously, but then to both supply and demand imbalances. These findings are in favor of Goldstein and Khan (1978) ("volumes respond to demand disequilibria, prices to supply disequilibria") and rather contradict Browne (1982) who assumed that for small open economies, volumes respond to supply disequilibria, prices to demand disequilibria. In countries for which export prices are exogenous, domestic producer prices and — for Canada — even foreign prices or the exchange rate play an important stabilizing role in the adjustment. For instance, a supply disequilibrium characterized by an "excess" export price leads to an increase in domestic producer prices thereby reducing the imbalance. The short-run part of the model in most cases seems to quicken the adjustments towards a new equilibrium, only in the equation of German export growth is there an indication that the negative impact of increasing export prices on the amount of goods and services sold abroad is stronger in the short run than in the long run.

The discussion of the reaction of the endogenous variables to innovations, changes in exogenous variables and imbalances is not yet complete. The indications given so far cannot replace the analysis of the impulse-response functions. This analysis should be carried out with the most parsimonious PVECM, i.e. after elimination of the insignificant short-run coefficients remaining in the system. A further reduction in the number of coefficients as well as a thorough impulse-response analysis represent the natural extension of the analysis presented in section 6. Another extension consists of using my results for model-based forecasts, e.g. a simultaneous forecast of export volumes and prices for Germany. Last but not least, it would be useful to see how sensitive the central finding of this paper, the existence of two cointegrating vectors, identifiable as export supply and demand, is with respect to the choice of variables. For instance, it would be interesting to see whether leaving out the domestic PPI (the typical supply variable) would reduce the rank to one. Likewise, one could call into question the *a priori* restriction set here that changes of equal size in the foreign PPI and the exchange rate have the same effect on export volumes and prices. This restriction made it possible to merge both variables to p^*_{NCU}. When using the trade-weighted foreign PPI and the nominal effective exchange rate separately instead, one would have to handle a VECM of seven I(1) variables and probably end up with three cointegrating vectors.

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