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Does Trade Drive Global Growth?

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Abstract

Conventional econometric analysis using VEC suggests that there is a long-term relationship between nominal world GDP and nominal world exports. The analysis cannot say anything about the causal relationships between the levels of GDP and exports. But it says a lot about the rules governing the short-term adjustments in GDP and exports. When considering such short-term adjustments, GDP plays the first fiddle. Short-term GDP changes have driven short-term changes in world exports, at least over the years 1987-2008. But the short-term changes in world exports did not 'cause' positive short-term changes in GDP.

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JEL classification: *F43, F15, O41, O49*

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Does trade drive global growth?

1 Introduction

For many decades now international trade has been gaining in importance. The share of global exports of goods and non-factor services in global GDP, which stood at 11.6% in 1960, climbed to over 32% in 2008 (before falling – during the 2009 global crisis – slightly below the 30% mark[1]). Many explanations have been put forward to explain the tendency for the trade share to rise (see for example Krugman, 1995; Frankel and Romer, 1999; and Baier and Bergstrand, 2001). The phenomenon of world trade growing faster than world GDP could be seen as reflecting progressing liberalisation of international trade (and of international flows of capital and ideas generally) as well as continuing advances in transport and communication technologies. In particular, technological progress combined with the tendencies to liberalise internationally (as well as internally, in major trading nations) are certainly jointly responsible for the development of the new internationalised forms of production organisation, as signified by the importance of offshoring, fragmentation of production, outsourcing of the manufacture of intermediate inputs to low-wage emerging markets etc. (Feenstra, 1998). The ongoing internationalisation of production naturally inflates the values of international trade relative to final output.[2]

Under the standard assumptions of the neoclassical trade theory liberalisation of trade and the reduction in trade costs should be conducive not only to 'more trade', but in the first place to more gains from trade – that is to more additional net output (due to increased efficiency of input allocation). Moreover, those gains should accrue (even if not necessarily equitably) to all countries participating in trade. In any case, cheaper and less restricted international trade is not, according to the conventional trade theory, hurting any trading country.

The 'new' theories of international trade (initiated by Krugman, 1979) and the new 'new' trade theories (as reviewed for example by Helpman, 2006) may not unequivocally support the view that more trade necessarily generates more output to all participating parties. Also, opinions openly doubting the benefits to individual nations of freer trade (often hinting at the advantages of some levels of protectionism) are not quite rare, especially among students of the developing countries (starting for instance from Bhagwati, 1958, to Stiglitz,

* Helpful comments by Neil Foster are gratefully acknowledged.

1 All numbers quoted come from the World Bank's World Development Indicators, 2012 Edition.

2 However, trade seems to have risen much faster relative to GDP *before* the present era of intensified liberalisation and *before* the major technology breakthroughs. This is evidenced by the early studies such as by Houthakker and Magee (1969) who sought to estimate the income (that is GDP) elasticties of demand for exports and imports of developed nations. Their estimates (with averages ranging between 1.6 and 1.7) indicated that trade had been growing much faster than GDP over the relatively illiberal 15-year period (1951-1966). It may be added that the world's exports-to-GDP ratio (calculated from the WDI data) had nearly doubled, to 20.9%, from 1960 through 1980. Thereafter, during the progressing liberalisation, it took almost 30 years for that ratio to reach 32.2%.

2001 or Thirlwall and Pacheco-Lopez, 2008 more recently). Interestingly, the Pope of the neoclassical trade theory himself expressed some heretical doubts about the doctrine he had long preached (Samuelson, 2004).

The reservations about the possibly undesirable consequences (including higher income inequality and depressed wages/employment in developed industrial countries) of growing trade notwithstanding, it is only fair to say that the hypothesis stipulating that 'trade growth drives GDP growth' has assumed the status of a dogma – if not quite uniformly among the academic economists, then at least within genuinely influential institutions. Without the dogma status of that hypothesis it would be rather hard to account for the persistent dogged efforts at global (and internal) liberalisation (GATT/WTO, IMF). Also, such integrative efforts as those on which the European Union (or NAFTA) is founded would lack economic rationale should the hypothesis be rejected.

However, is there compelling empirical evidence supporting that hypothesis when applied to the aggregate *global* economy? Quite surprisingly, the research does not seem to have addressed itself to testing that hypothesis. Naturally, there are numerous studies concerned with the evaluation of the role of trade for individual countries (or 'panels' of countries). However, the rich empirical literature on growth accounting, concerned with the quantification of sources of the long-term income (or/and productivity) growth across time and space, is not quite supportive of the hypothesis endowing rising foreign trade with growth enhancing abilities. As recently documented by Hillebrand et al. (2010), *'... there is a troubling disconnect between the economic growth literature and the trade literature ...'*. Classical studies such as Denison (1985) dismiss trade as the source of the US longer-term economic growth, or fail to mention it altogether (see also Jorgenson, 2005). Econometric studies, of which there is no shortage, attempting to quantify the impacts of various factors on GDP growth rates (or on total factor productivity growth) across larger samples of countries typically do not support the hypothesis on the productive role of trade. Studies attempting to find a positive association between GDP levels and various trade-related indicators (such as measures of openness) do not produce generally accepted conclusions. For example, Rodrik et al. (2004) find out that *'... once institutions are controlled for, trade is almost always insignificant, and often enters the income equations with the "wrong" (i.e. negative) sign ....'.* Given the fact that the longer-term growth performances of most individual countries cannot really be explained by the foreign trade developments, one may not claim that the long-term growth of global (world) income has been meaningfully driven by the rising volume of global trade.

It is rather obvious that in the shorter run the growth of output of some *individual* countries may heavily rely on expansion of their exports. Moreover, the growth of productivity (and of *potential* output) in many cases may depend upon rising imports of capital goods and intermediate inputs. Rising *net* exports may contribute substantially to overall GDP growth in *some* nations.[3] Examples of countries successfully following 'export-led' growth paths abound. But it must be remembered that for each country relying for GDP growth on the improvement of net exports there must be some other countries whose net exports deteriorate – thus depressing their GDP growth. The existence of a club of countries following the 'export-led' growth paths implies the existence of a club of 'import-fed' countries whose GDP growth must sooner or later be held back by contracting net exports. The global economy – being an autarchic system – cannot follow the export-led growth path.[4]

This paper sets out to analyse econometrically the dynamic relationships between world GDP and world trade (which is identified using world exports). Section 2 discusses the data used in the analysis. Section 3 conducts cointegration analysis, presenting the resulting estimates of a Vector Error Correction (VEC) model. The analysis shows that movements in GDP drive movements in exports while movements in exports are not really followed by meaningful movements in GDP. In this sense trade does *not* cause growth – while growth *does* cause trade. Section 4 presents further results derived from a conventional Vector Autoregression (VAR) model. These results support the results derived from the VEC model. Section 5 concludes.

2 The data

The analysis that follows works with two time series taken from the World Bank's World Development Indicators (WDI) data set: world GDP and world exports of goods and non-factor services (as reported by the Balance of Payments). Both items are expressed in current US dollars. Of course it would be desirable to work with the real volumes of GDP and exports. But the WDI do not provide data on the volumes of world exports, though they do provide data on volumes of world GDP. The calculation of export volumes would require deep studies on meaningful price indices for world trade, still a task for the future.[5] (Even the calculation of price deflators for world GDP – which the WDI report starting from 1965 – is a risky business. Deflating the WDI world GDP values with the WDI world GDP deflators one arrives at a world GDP volume series that does not resemble the world GDP volume series reported by WDI at all.)

The world trade and GDP series currently available from WDI extend from 1960 through 2010. Figure 1 shows the development of the trade/GDP ratio over the whole period. As

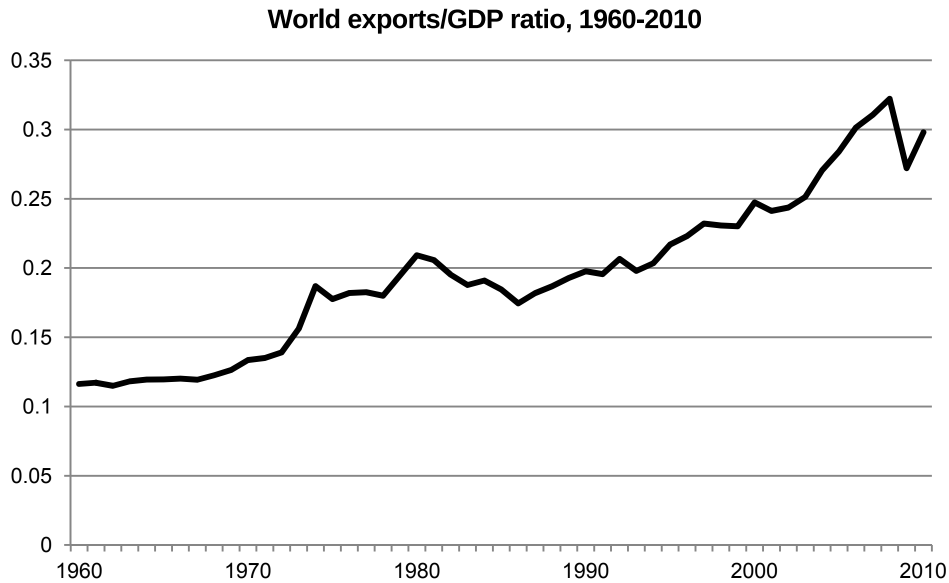
[3] Observe that rising net exports may well be achieved at the cost of overall GDP growth stagnation. This is the case in Germany where high trade surpluses (achieved through the sustained repression of wages and domestic demand) have been associated with secularly anaemic GDP growth (Laski and Podkaminer, 2012).

[4] In any period the change in GDP can be uniquely decomposed into contributions from concurrent exports, imports and domestic demand. Contributions of trade (exports and imports taken together) to any trading country's GDP change can be large or small, negative or positive. But the contributions of trade to *global* GDP growth can only be zero. Seen from the demand side, world GDP can only be driven by consumption and gross capital formation – not by trade.

[5] Feenstra (1994) illustrates some of the difficulties involved in the measurement of price indices for US trade. Measuring price indices for *world* trade must be incomparably more difficult.

can be seen, the ratio followed a quite smoothly accelerating growth trajectory until 1973. A period of instability ensued. By 1987 the ratio seemed to have returned to the pre-1973 trajectory which then abruptly terminated in 2009.

Figure 1

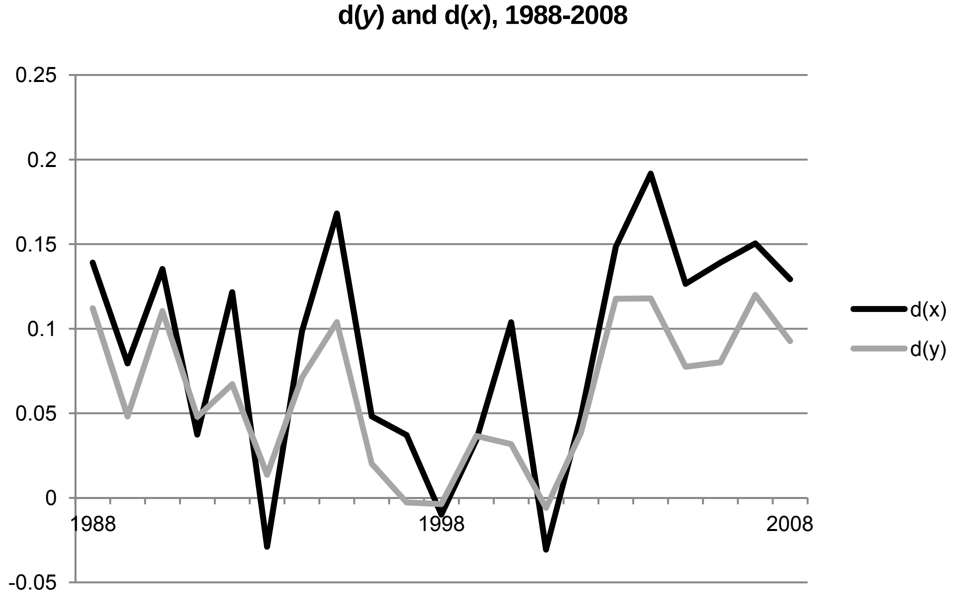


Source: Own calculations based on WDI, 2012 Edition (August).

The analysis to follow is limited to developments from 1987 through 2008. The instability period (1973-1987) differs from both the preceding and succeeding ones on too many essential counts. Two major oil price shocks hit the world economy during that period – fits of very high inflation followed in their wakes, probably additionally inflating the values of trade relative to the values of (depressed) GDP. Moreover, that was a period of great instability in exchange rates which started with the final demise of the Bretton Woods system in 1973 and effectively ended only in 1987 (following the Plaza Accord of 1985 and the Louvre Accord of 1987). Wild longer-term fluctuations in the US dollar exchange rates during that period may have disturbed the underlying relationship between changing trade and changing GDP. Throughout the period the creeping liberalisation of capital flows was followed by a series of sovereign debt crises (e.g. in Latin America) with consequences for global growth and trade. Finally, the exclusion of 2009 (and 2010) also seems to make sense. The great recession of 2009 constituted a true shock to world GDP and to world trade. (For many reasons studied extensively by numerous researchers, the 2009 recession in trade was much deeper than in GDP.)

The following analysis works with the natural logarithms of world GDP and world exports, denoted as y and x respectively. x and y trend together (their simple correlation coefficient equals 0.9950). Figure 2 shows the differenced series d(y) and d(x). As can be seen, d(y) and d(x) are also strongly correlated (the simple correlation coefficient equals 0.915) but do not show much of a common trend.

Figure 2



Source: Own calculations.

Both items (*x* and *y*) are of course non-stationary while their first differences d(*y*) and d(*x*) are stationary (Table 1).

Table 1

ADF tests for the order of integration of log(exports) and log(GDP), years 1987-2008

series	Lag length*	ADF test statistics	Probability**	Conclusion
x	0	0.2823	0.9713	Non-stationary
y	1	-0.2880	0.9720	Non-stationary
d(*x*)	0	-3.4231	0.0211	Stationary
d(*y*)	0	-3.059	0.0450	Stationary

The ADF testing equations assumed an intercept. Nonstationarity of *x* and *y* is not rejected also assuming intercepts and linear trends.

*) Selected automatically based on the Schwartz Information Criterion (max lag = 8).

**) MacKinnon (1996) one-sided p-values.

The results of other unit root tests are in agreement with the conclusions in Table 1.

The next step is to check whether there is Granger causality between d(*x*) and d(*y*).[6] Table 2 strongly suggests that d(*y*) Granger-causes d(*x*) while d(*x*) does *not* seem to Granger-cause d(*y*).

6 There are a large number of empirical studies looking for causality (including Granger causality) between exports and growth in *individual* countries, or 'panels' of countries. The reported results of these studies tend to be mixed and/or subject to various criticisms (e.g. over the measurement of openness, endogeneity of trade policies etc.). The fact that exports can advance growth in some countries – at the expense of retarding growth in importing countries – has yet to be properly accounted for in the panel or cross-country studies.

Table 2

Pairwise Granger causality tests between d(*x*) and d(*y*), years 1987-2008

No of lags*	Null hypothesis	F-statistics (p-value)	Conclusion
1	d(*x*) does not Granger-cause d(*y*)	2.7488 (0.1137)	
1	d(*y*) does not Granger-cause d(*x*)	8.1633 (0.0100)	d(*y*)→d(*x*)
2	d(*x*) does not Granger-cause d(*y*)	1.5522 (0.2404)	
2	d(*y*) does not Granger-cause d(*x*)	3.8703 (0.0412)	d(*y*)→d(*x*)
3	d(*x*) does not Granger-cause d(*y*)	3.070 (0.0634)	d(*x*) →d(*y*)**
3	d(*y*) does not Granger-cause d(*x*)	5.002 (0.0134)	d(*y*)→d(*x*)

*) Number of lags in the testing equations. At longer lags the case for non-rejection of the hypothesis on d(*x*) not Granger-causing d(*y*) gets progressively stronger (the respective p-values become much larger), while the hypothesis on d(*y*) not Granger-causing d(*x*) are rejected at the 0.04 level. The arrow (in the 'Conclusion' column) stands for the direction of Granger causality.

**) At three lags one can reject Granger causality not running from d(*x*) to d(*y*), though at the relatively large p-value (0.0634).

3 Trade and GDP appear to be cointegrated

Further statistical inferences on the links between *x* and *y* require checking for the presence of so-called cointegration: although *x* and *y* appear to be non-stationary, some specific linear combination of the two series (with intercepts or deterministic trends eventually added) may be stationary. In such a case this cointegrating linear combination of *x* and *y* (denoted here as E) would represent a long-run ('equilibrium') relationship between *x* and *y*. In the long run (and in the absence of external disturbances), E is assumed to equal zero. E taking on a value different from zero indicates the occurrence of an imbalance (or error) which the short-term movements in *x* and *y* would gradually reduce. Should *x* and *y* be cointegrated, the short-term changes in *x* and *y* would be captured by two equations:

$$d(x) = \alpha_1 E(-1) + \beta_1 d(x(-1)) + \beta_2 d(x(-2)) + \beta_3 d(x(-3)) + \ldots + \gamma_1 d(y(-1)) + \gamma_2 d(y(-2)) + \gamma_3 d(y(-3)) + \ldots \eta_x$$

and (1)

$$d(y) = \alpha_2 E(-1) + \delta_1 d(x(-1)) + \delta_2 d(x(-2)) + \delta_3 d(x(-3)) + \ldots + \varepsilon_1 d(y(-1)) + \varepsilon_2 d(y(-2)) + \varepsilon_3 d(y(-3)) + \ldots \eta_y$$

where E(-1) is the imbalance recorded in the previous year; *x*(-1), *y*(-1), *x*(-2), *y*(-2), ... represent past values of *x* and *y* respectively (lagged 1, 2, ...); α, β, γ, δ, ε are parameters to be estimated jointly with the parameters characterising the long-run relationship E; ηx and ηy are unobservable random disturbances. The terms $\alpha 1$E(-1) and $\alpha 2$E(-1) represent momentary adjustments towards the long-run 'equilibrium', with $\alpha 1$ and $\alpha 2$ representing speeds of adjustments. The remaining items on the right-hand sides of the above equations represent the effects of past (lagged) changes in *x* and *y* respectively. Of course, the number of lags taken into account should be reasonably small. (In our case the sample, consisting of values for 22 years, permits the estimation of the parameters of equations (1) with a maximum of 7 lags.)

Commonly used Johansen tests suggest the existence of cointegration of *x* and *y*. However, there are many specific cointegrating relationships E, depending on the assumed maximum lag for d(*x*) and d(*y*) and the specifics concerning intercepts/trends. The statistical qualities of the alternative systems of equations (1) are different.

The following system with three lags, passing both Johansen cointegration tests (Trace Test and Maximum Eigenvalue Test, both at the 0.05 level), appears to have quite good statistical properties. Consequently, it is legitimate to apply the Vector Error Correction (VEC) estimation approach. The long-run equilibrium relationship E is estimated as:

$$E(\tau) = (x(\tau) - 1.051059 \cdot y(\tau) - 0.025274 \cdot (\tau - 1960) + 4.02061)$$

(0.100)	(0.0045)
[-10.5]	[-5.58]

(τ denotes the date (year). The standard error of the estimate is in round brackets, the t-statistics in the square brackets).

Equations (1) are specified as follows:

$$d(x) = -1.39 \cdot E(-1) - 0.76 \cdot d(x(-1)) - 0.37 \cdot d(x(-2)) + 0.85 \cdot d(x(-3)) + 1.66 \cdot d(y(-1)) + 1.27 \cdot d(y(-2)) - 0.61 \cdot d(y(-3)) - 0.033$$

(0.312)	(0.233)	(0.26)	(0.26)	(0.32)	(0.022)	(0.425)	(0.24)
[-4.47]	[-3.26]	[-1.42]	[2.23]	[5.23]	[3.17]	[-1.43]	[-1.39]

(2)

$$d(y) = -0.68 \cdot E(-1) - 0.49 \cdot d(x(-1)) - 0.33 \cdot d(x(-2)) + 0.56 \cdot d(x(-3)) + 1.07 \cdot d(y(-1)) + 0.80 \cdot d(y(-2)) - 0.35 \cdot d(y(-3)) - 0.011$$

(0.264)	(0.197)	(0.22)	(0.22)	(0.269)	(0.34)	(0.36)	(0.2)
[-2.59]	[-2.48]	[-1.52]	[2.52]	[3.99]	[2.37]	[-0.98]	[-0.53]

The R-squared equals 0.808 for the d(*x*) equation and 0.697 for the d(*y*) equation (for details see Appendix Table A1). The system better tracks changes in d(*x*) than in d(*y*). Equations pass the usual diagnostic tests with flying colours[7].

It must be observed that the estimates of δ_1 and δ_2 are *negative*. Only δ_3 is positive. But the sum $\delta_1 + \delta_2 + \delta_3$ is negative all the same (-0.26). In contrast, the estimates of γ_1 and γ_2 are *positive*. While the estimate of γ_3 is negative (though essentially not different from zero), the sum $\gamma_1 + \gamma_2 + \gamma_3$ is positive and much larger (2.32) than the absolute value of ($\delta_1 + \delta_2 + \delta_3$).

7 The system is stable (all inverse AR roots lie inside the unit circle while one root is equal 1, as it should). The pairwise Granger causality/Block Exogeneity Wald Tests reject the exogeneity of d(*x*) and d(*y*). The Lag Length Test does not exclude lags 1, 2 and 3. Residual autocorrelation tests do indicate the absence of autocorrelation and serial correlation. Residuals are normally distributed, with the joint p-value for the Jacque-Bera statistics exceeding 0.6. The p-values for White VEC Heteroskedasticity Test are very high – residuals appear to be homoskedastic. Other cointegrated system specifications – even if passing the Johansen cointegration tests at the 0.01 level and exhibiting better fit (and/or other information indicators) – tend to fail some (or many) of the above diagnostic tests.

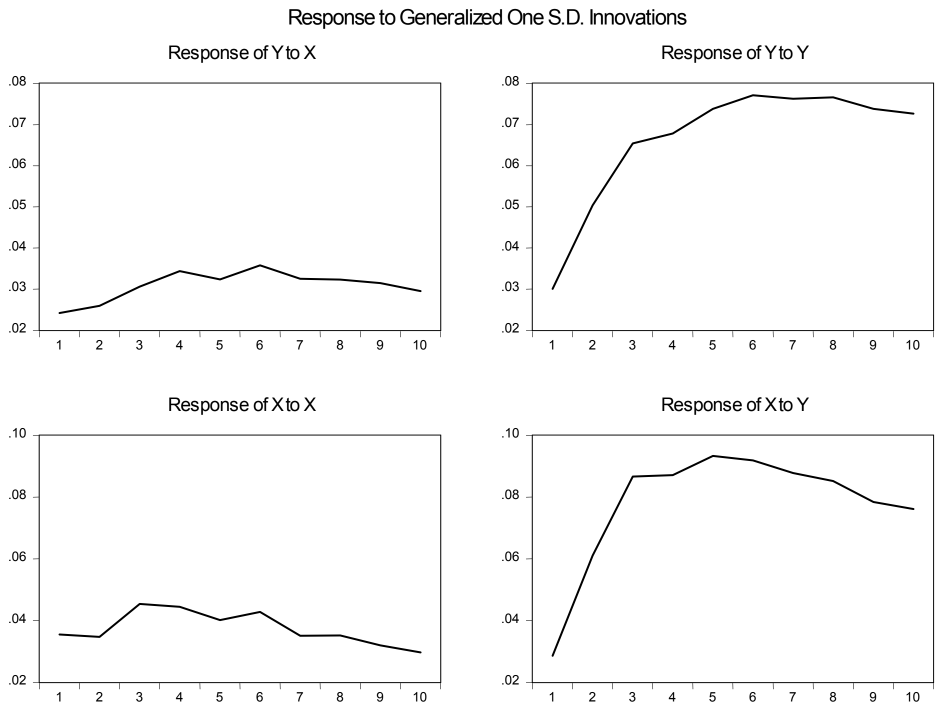
Figuratively speaking, all else being equal, a positive unit rise in world exports would likely be followed by weak response (possibly even a fall) in GDP (and also in exports). But, all else being equal, a positive rise in the world GDP is likely to be followed by a strong response (possibly a rise) of both exports and GDP. Figure 3 shows the generalised impulse responses for the VAR given by (2).

Figure 3 shows *generalised* (as defined by Pesaran and Shin, 1998) impulses for equations (2). The impulse responses in Figure 3 (and in Figures 4-5) allow for the correlations between shocks (including contemporaneous shocks), as implied by the data. Moreover, impulse responses derived here are independent of the ordering of variables. This is why such impulses are termed 'generalised'. Impulse responses ignoring correlations between shocks show strong *absolute declines* in *y* and *x* following a positive shock to *x* and strong *absolute increases* in *x* and *y* following a positive shock to *y* (see Appendix Figure A1).

Of particular interest are the responses of *y* to *x* (the upper left-hand panel) and responses of *x* to *y* (the bottom right-hand panel). The former panel shows that a momentary (one-off) 'positive shock' (or 'innovation') to *x* is followed by a weak and delayed response of *y*. There are no additional effects beyond the fourth year. In contrast, the effects on *x* of a momentary (one-off) positive shock to *y* are not only immediate and incomparably stronger. In addition these effects additionally increase over a longer horizon.

Figure 3

Responses to generalised one standard deviation innovations to *y* and *x*, the VEC equations (2)



The impulse responses in Figure 3 are derived from one concrete cointegrated system, given by equations (2). But there are many cointegrated systems[8] for the same set of data. Is it possible that for some of the other systems the impulse response schedules would be qualitatively different? The answer is no. Without reproducing 21 additional figures on impulse responses derived from these systems, it should suffice to state that the estimates for δ and γ for *all other* cointegrated systems exhibit characteristics similar to those shown in equations (2). Namely, almost each estimated δ is *negative*, and each estimated γ is *positive*. (In addition, estimates for β are negative and estimates for ε are positive.) The sums of γs are positive and the sums δs are negative (and much smaller in absolute terms than the former). Under such a constellation of parameters the impulse response of *y* to *x* is at best weak (in some cases negative) and rather short-lived while the impulse response of *x* to *y* is strong and lasting longer, similarly as in Figure 3. (Moreover, in most cases estimates of γs are much more *significant* [have much lower standard errors] than the corresponding δs. The t-statistics for γ_i tends to be much bigger than for the corresponding δ_i. Of course, this finding is consistent with the findings of the Granger causality testing reported in Section 2.)

4 Evidence from Vector Autoregressions

Impulse responses that can be derived from the Vector Error Correction (VEC) models such as the one represented by Figure 3 do not show the likely *ranges* of the responses following specific one-off shocks hitting the system. Thus the responses shown in Figure 3 represent point estimates of the effects of such shocks.

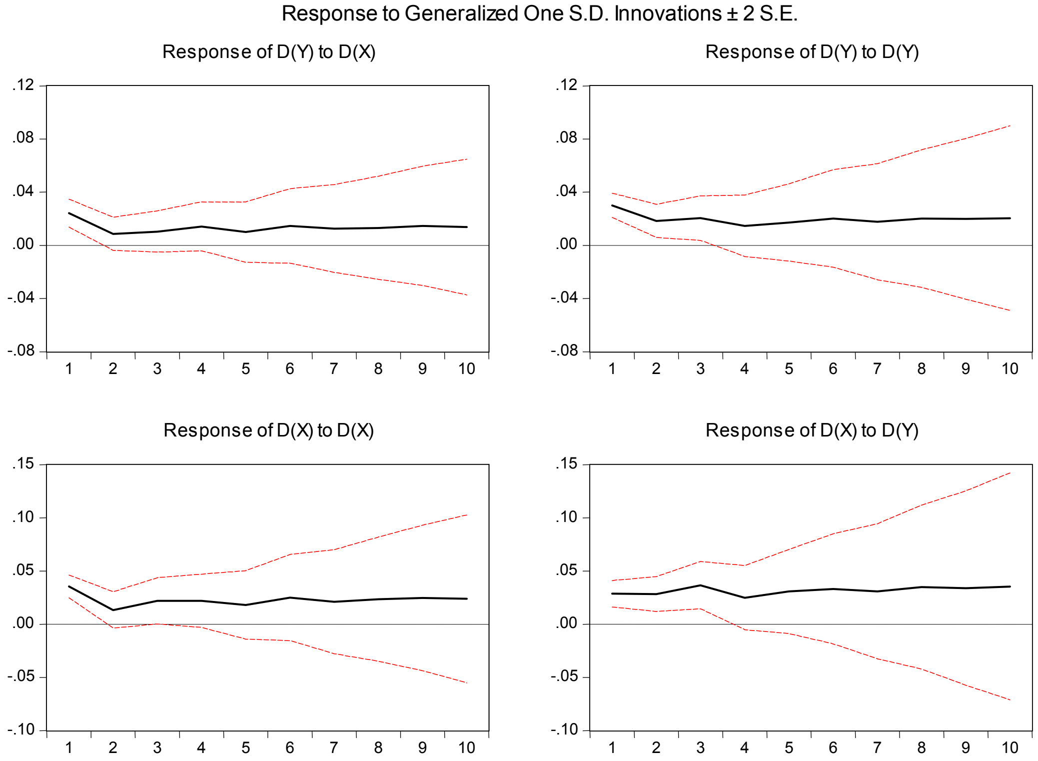
However, one may easily construct a Vector Autoregression (VAR) model making full use of information provided by the data and the E estimated using VEC. Such a VAR assumes two endogenous variables (d(*x*) and d(*y*)), three lags, and two exogenous variables (intercept and the long-term relationship E, the latter numerically specified as in the VEC estimation). It appears that such a derivative VAR model exactly reproduces all remaining original VEC parameter estimates and all other statistics. The advantage of this transformation is that the impulse responses derived from the VAR model allow the presentation of the likely bands around the 'averaged' responses to d(*x*) and d(*y*). Figure 4 shows such impulse responses with bands for the 'VEC turned VAR' model.

The impulse responses in Figure 4 show the dynamic responses following one-off 'positive shocks' to d(*x*) and d(*y*), equal one standard error of the residuals to the respective equations in (2). Additionally, the panels show the confidence bands around the point estimates in question (equal to two standard errors, either way).

[8] In total 22 systems pass the Johansen cointegration tests. When 2 lags are assumed there are 3 different cointegration systems. At 3, 4, 5, 6, 7 lags assumed, the Johansen tests detect, respectively, 1, 5, 4, 4, 5 different cointegrated systems. (No cointegration is detected for the system allowing for 1 lag only.)

Figure 4

Responses to generalised one standard deviation innovations to d(*y*) and d(*x*), the VAR/VEC equations (2)



Again, the upper left-hand-side and the lower right-hand-side panels are of primary interest. The former panel indicates that the one-off 'positive shock' to d(*x*) is as likely to be followed by muted change in d(*y*). Already within one year the lower confidence band for d(*y*) enters the negative territory, indicating a likelihood of a decline in d(*y*). The lower confidence band in the latter panel enters the negative territory only in the fourth year. A positive one-off shock to d(*y*) is much more likely to be followed by *increased* d(*x*) over the first three years.

The VAR analysis (initiated quite long ago by Sims, 1980) has been commonly applied before the advances in the cointegration (VEC) methodology for the study of the behaviour of systems with nonstationary but possibly cointegrated variables.[9] Ignoring the evidence in favour of the cointegration of variables *x* and *y* (and thus dismissing the existence of a long-run 'equilibrium' relationship between world exports and world GDP) one can estimate and analyse the VAR models with d(*x*) and d(*y*) as dependent variables. Equations (3) show the VAR equations estimated assuming three lags in variables and the exogenous linear trend.

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[9] In their evaluation of the 20 years of VAR as a tool of macroeconomic analysis, Stock and Watson (2001) did not mention the term 'cointegrations' even once.

$$d(x)=-1.36\cdot d(x(-1))-0.98\cdot d(x(-2))+0.26\cdot d(x(-3))+1.87\cdot d(y(-1))+1.76\cdot d(y(-2))-0.08\cdot d(y(-3))-0.25+0.008\cdot(\tau-1960)$$

(0.40)	(0.50)	(0.42)	(0.43)	(0.66)	(0.70)	(0.11)	(0.003)
[-3.40]	[-1.98]	[0.61]	[4.29]	[2.65]	[-0.12]	[-2.23]	[2.56]

(3)

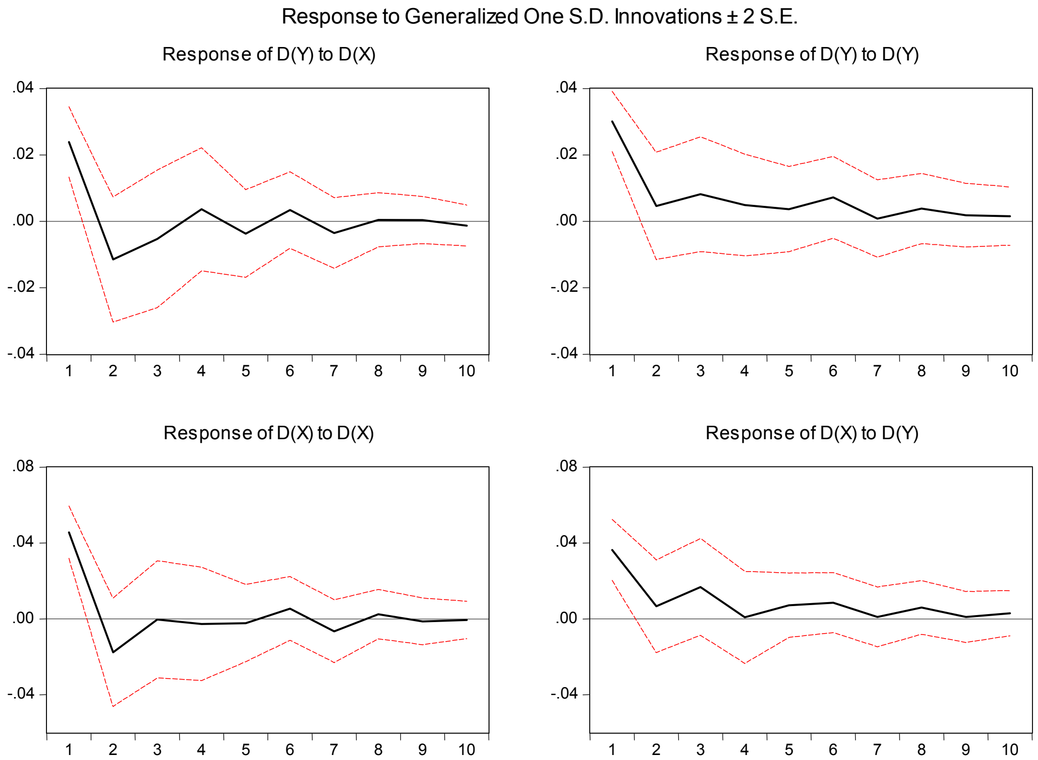
$$d(y)=-0.91\cdot d(x(-1))-0.81\cdot d(x(-2))+0.16\cdot d(x(-3))+1.25\cdot d(y(-1))+1.25\cdot d(y(-2))+0.13\cdot d(y(-3))-0.02+0.005\cdot(\tau-1960)$$

(0.26)	(0.33)	(0.28)	(0.29)	(0.44)	(0.46)	(0.07)	(0.002)
[-3.43]	[-2.47]	[0.57]	[4.35]	[2.86]	[0.27]	[-2.26]	[2.59]

The R squared for the first equation in (3) is 0.6816 and 0.6967 for the second equation (see Appendix Table A2 for details). As can be seen, the standard errors of most estimates are large, many with rather weak t-statistics. Such parameters can be hypothesised to equal zero. The only more significant variables in either equation are d(*x*(-1)) and especially d(*y*(-1)). It is worth noting that the variables d(*x*(-1)) and d(*x*(-2)) enter both equations with *negative* signs. Conversely, the variables d(*y*(-1)), d(*y*(-2)) and d(*y*(-3)) enter both equations with *positive* signs.

Figure 5

Responses to generalised one standard deviation innovations to d(*y*) and d(*x*), the VAR model (3)



Looking at the upper left-hand-side panel in Figure 5 (which shows generalised impulse responses for VAR given by equations (3)) one sees that an impulse to d(*x*) is much more likely to be followed by a decline than a rise in d(*y*) during the next two years. However, an impulse to d(y) is much more likely to be followed by a rise than a decline in d(x) over the next two years. Thus, the simple VAR with three lags[10] generates impulse response schedules carrying similar message as the VEC.

5 Concluding remarks

Econometric analysis suggests that there may have been a long-term ('equilibrium') relationship between the *levels* of nominal world GDP and nominal world exports. The analysis cannot say anything about the *causal* relationships between the *levels* of GDP and exports. But it can say something about the rules governing the short-term adjustments in GDP and exports. It turns out that when considering such short-term adjustments GDP plays the first fiddle. Short-term GDP changes seem to have driven short-term changes in world exports, at least over the years 1987-2008. The evidence strongly suggests that the short-term changes in world exports did *not* 'cause' short-term changes in GDP. In this sense the analysis refutes the popular belief that 'exports cause growth'. The opposite appears to be closer to the truth.

These are tentative conclusions, based on relatively short time series. Further research may still be needed to check whether they hold also with respect to the *volumes* of trade and GDP, not only with respect to their values. Also, it would be useful to repeat the analysis with quarterly data on world trade and GDP (should these be available).

Many questions remain open. What are the 'theoretical' reasons for the empirical patterns of the short-term adjustments revealed? Are these patterns consistent with some specific interpretations of the mechanisms governing *contemporary* global macro-economy? Also, the long-run relationship (E) between the logarithms of GDP and exports suggested by the analysis (E = Log(exports)-1.05106·Log(GDP)-0.0245·(τ-1960)+4.0206) deserves some deeper reflection. Assuming, for example, that exports are a factor of production (on which the supply of output in the importing countries relies) it would appear that the marginal productivity of world imports (world imports in principle must equal world exports) is *diminishing*:

$$GDP(\tau) = A(\tau)\cdot(imports)^{0.9514}$$

where A(τ) = exp[3.8253-0.0245·(τ-1960)] (τ is the date (year); 3.8253 = 4.0206/1.05106; 0.0245 = 0.02527/1.05106 and 0.9514 = 1/1.05106). How should one square the diminishing (long-run) marginal productivity of world trade with the conventional beliefs about its beneficial long-term productivity effects? A heuristic answer could be that world trade *could* have been productive on the global scale should the GDP growth in individual countries engaged in international trade have been approximately balanced most of the time – and

[10] At larger lag lengths the VAR models generate impulse response schedules with very broad confidence bands. Essentially, such VARs cease to be informative.

not only occasionally, in response to the severe payments or exchange rate crises. It is imaginable that reaping the productivity (in terms of output per employee) gains in importing countries has been associated with the total GDP growth slowdowns arising over growing or persistent trade deficits and reflected in high or rising unemployment. GDP growth in the net exporter countries may also have suffered because their high/persistent trade surpluses are often engineered by the policy of wage and domestic demand repression (and/or result from particularly skewed income distributions). Thus, diminishing marginal productivity of trade may have emerged under huge trade *imbalances* that have developed under progressing globalisation. Under a regime enforcing more balanced trade among nations, with major nations not allowed to compensate deficient domestic demand with huge trade surpluses that destabilise their partners, the marginal productivity of global trade need not, perhaps, be diminishing. Of course, for the individual countries to follow the Thirlwallian (McCombie and Thirlwall, 2004) balance-of-payments constrained growth paths, not only would the international economic order need to be overhauled. Also the basic paradigms of the domestic macroeconomic policy making in major nations would have to be radically changed.

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Appendix

Table A1

VEC estimation output (equations 2)

Vector Error Correction Estimates
Date: 08/24/12 Time: 09:56
Sample: 1987 2008
Included observations: 22
Standard errors in ( ) & t-statistics in [ ]

Cointegrating Eq:	CointEq1	
x(-1)	1.000000	
y(-1)	-1.051065	
	(0.10010)	
	[-10.5003]	
@TREND(60)	-0.025274	
	(0.00453)	
	[-5.58075]	
C	4.020799	
Error Correction:	d(x)	d(y)
CointEq1	-1.395477	-0.683622
	(0.31216)	(0.26424)
	[-4.47045]	[-2.58716]
d(x(-1))	-0.759198	-0.488922
	(0.23289)	(0.19714)
	[-3.25985]	[-2.48005]
d(x(-2))	-0.368576	-0.333816
	(0.25965)	(0.21979)
	[-1.41953]	[-1.51881]
d(x(-3))	0.846080	0.559240
	(0.26162)	(0.22146)
	[3.23396]	[2.52523]
d(y(-1))	1.662899	1.073504
	(0.31768)	(0.26891)
	[5.23455]	[3.99205]
d(y(-2))	1.271737	0.804094
	(0.40128)	(0.33968)
	[3.16919]	[2.36721]
D(y(-3))	-0.609401	-0.352984
	(0.42517)	(0.35991)
	[-1.43330]	[-0.98077]
C	-0.032756	-0.010652
	(0.02353)	(0.01992)
	[-1.39215]	[-0.53481]

R-squared	0.807510	0.696661
Adj. R-squared	0.711265	0.544991
Sum sq. resids	0.017637	0.012638
S.E. equation	0.035494	0.030045
F-statistic	8.390156	4.593277
Log likelihood	47.20006	50.86657
Akaike AIC	-3.563642	-3.896961
Schwarz SC	-3.166899	-3.500218
Mean dependent	0.092558	0.064660
S.D. dependent	0.066054	0.044541
Determinant resid covariance (dof adj.)		4.00E-07
Determinant resid covariance		1.62E-07
Log likelihood		109.5487
Akaike information criterion		-8.231702
Schwarz criterion		-7.289438

Table A2

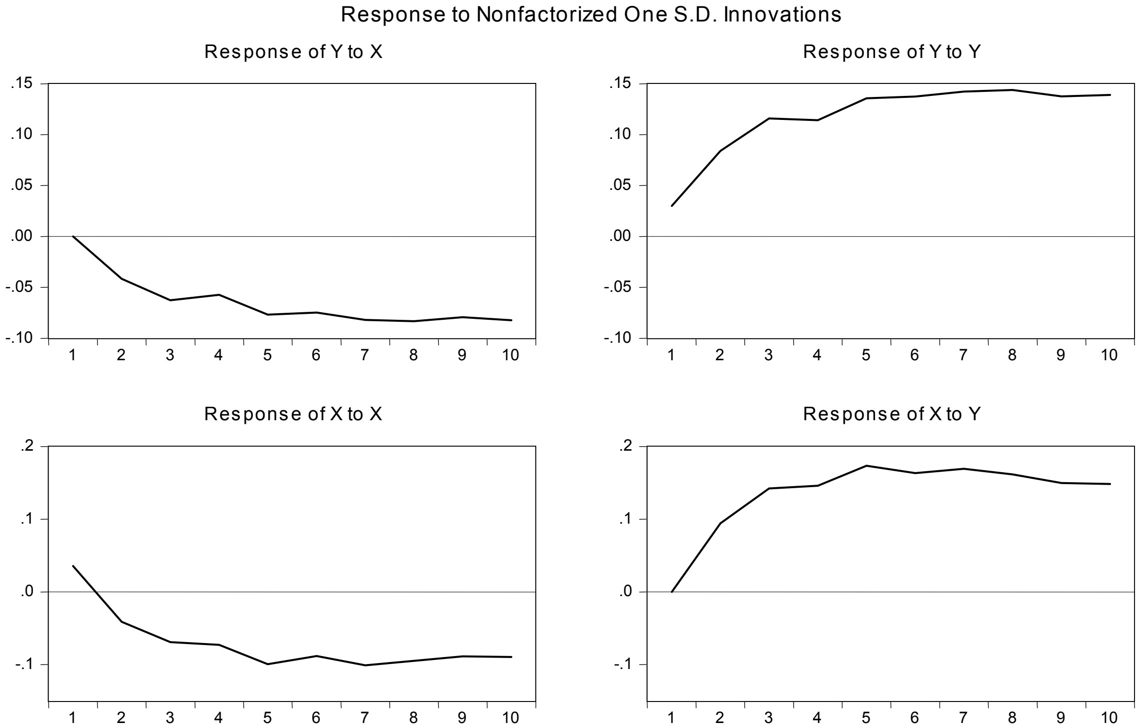
VAR estimation output (equations 3)

Vector Autoregression Estimates
Date: 08/24/12 Time: 10:09
Sample: 1987 2008
Included observations: 22
Standard errors in ( ) & t-statistics in [ ]

	d(x)	d(y)
d(x(-1))	-1.363976	-0.905429
	(0.40093)	(0.26388)
	[-3.40207]	[-3.43122]
d(x(-2))	-0.982839	-0.805910
	(0.49578)	(0.32631)
	[-1.98241]	[-2.46975]
d(x(-3))	0.258568	0.156901
	(0.42165)	(0.27752)
	[0.61323]	[0.56537]
d(y(-1))	1.869904	1.248558
	(0.43593)	(0.28692)
	[4.28949]	[4.35163]
d(y(-2))	1.761849	1.252879
	(0.66452)	(0.43737)
	[2.65133]	[2.86458]
d(y(-3))	-0.083218	0.126715
	(0.70555)	(0.46438)
	[-0.11795]	[0.27287]
C	-0.251384	-0.167168
	(0.11259)	(0.07410)
	[-2.23276]	[-2.25587]
@TREND(60)	0.007887	0.005249
	(0.00308)	(0.00203)
	[2.55885]	[2.58747]
R-squared	0.681631	0.696684
Adj. R-squared	0.522447	0.545027
Sum sq. resids	0.029171	0.012637
S.E. equation	0.045647	0.030044
F-statistic	4.282026	4.593792
Log likelihood	41.66523	50.86743
Akaike AIC	-3.060476	-3.897039
Schwarz SC	-2.663733	-3.500296
Mean dependent	0.092558	0.064660
S.D. dependent	0.066054	0.044541
Determinant resid covariance (dof adj.)		6.91E-07
Determinant resid covariance		2.80E-07
Log likelihood		103.5476
Akaike information criterion		-7.958873
Schwarz criterion		-7.165387

Figure A1

Responses to non-factorised innovations, VEC equations (2)



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