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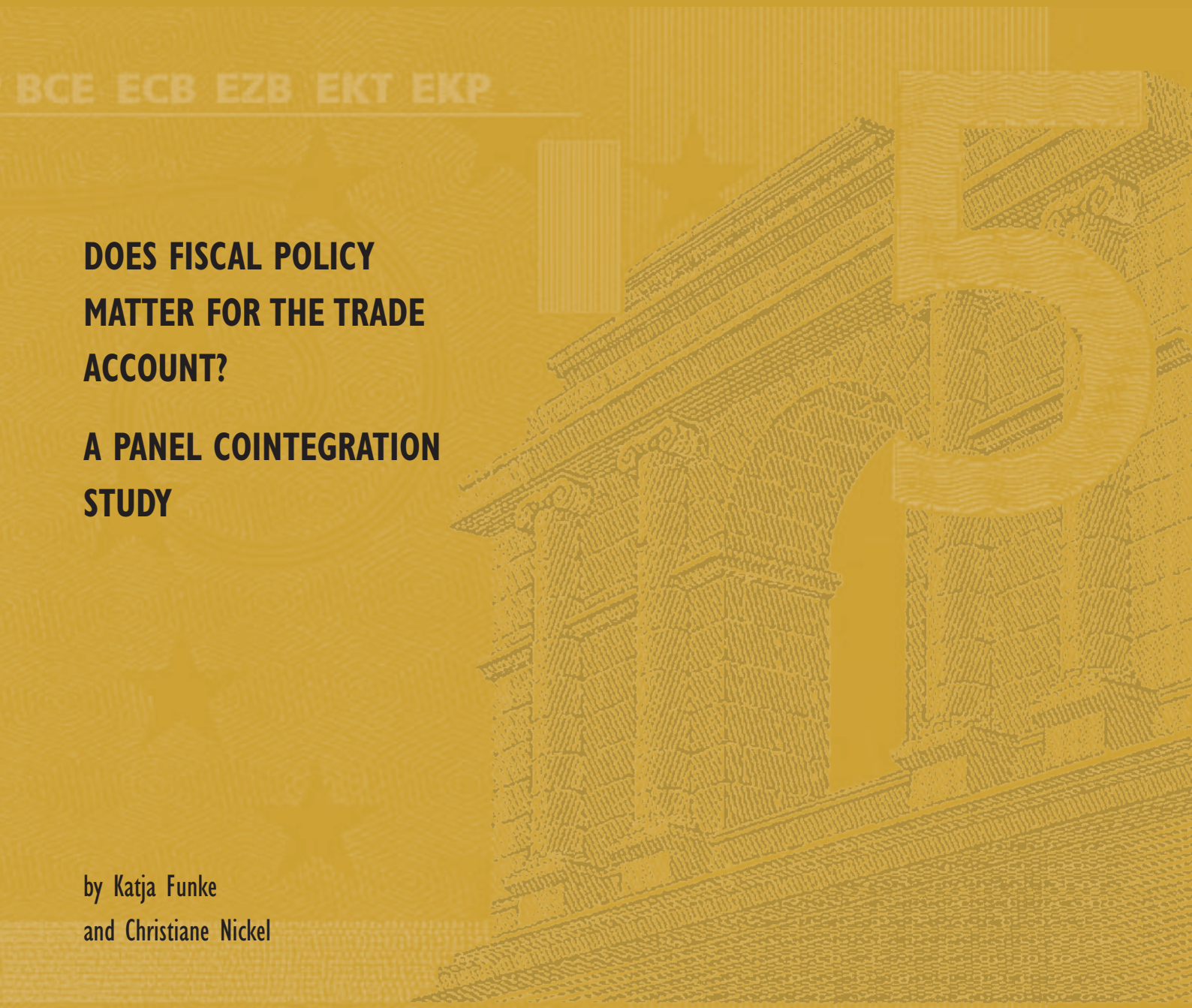

DOES FISCAL POLICY MATTER FOR THE TRADE ACCOUNT?

A PANEL COINTEGRATION STUDY

by Katja Funke
and Christiane Nickel

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Abstract

This paper analyses the empirical relationship between fiscal policy and the trade account. Research prior to this paper did not consider that the components of private and public demand in the import demand equation exhibit different elasticities. Using pooled mean group estimation for annual panel data of the G7 countries for the years 1970 through 2002, we provide empirical evidence that the composition of overall demand – i.e. the distribution among public demand, private demand and export demand – has an impact on the magnitude of the trade account deficit.

Key words: Fiscal policy, trade account, trade elasticities, panel cointegration

JEL: F32, E62, F41

Non-technical summary

Little is known about the effects that a lasting change in government expenditures has on a country's external balance. There seems to be a consensus that lower expenditures and the concomitant improvement in the fiscal balance lead to an improvement in the current account. Empirical research so far, however, has led to ambiguous results: Some empirical studies find that higher budget deficits lead to higher current account deficits; others prove the opposite or show no significant impact at all. A flaw of earlier research could be that reduced-form equations are estimated, wherein different effects might counteract each other without showing the underlying causalities. The latter can only be revealed in a structural model. Furthermore, earlier studies suffer from the fact that econometric techniques that allow studying long-run equilibrium relationships between time series data were not yet developed.

This paper takes a fresh look at the empirical relationship between fiscal policy and the trade account by analysing the relationship between government expenditures and imports. Because trade account deficits are often at the heart of current account problems, a structural model of the trade account is an important step when modelling the impact of fiscal policies on the external balance. Within the trade account, we concentrate on imports because import demand is determined by domestic demand factors, while exports depend on external demand factors. To pin down the effects of fiscal policy we estimate goods and service import equations on the basis of disaggregate demand variables. This implies that – in contrast to the conventional form of trade equations, which take total demand as an explanatory variable – we allow for all components of demand, i.e., private consumption, private sector investment, government expenditure, and exports, to exhibit different elasticities.

Our empirical analysis is based on annual panel data of the G7 countries for the years 1970 through 2002. We determine the cointegration relationship by a pooled mean group estimation. This technique allows the intercepts, short-run coefficients and error variances to differ freely across countries, while the long-run coefficients are constrained to be the same for all cross sections. We are therefore able to account for cross country differences without losing the general message about the long-run relationships between import volumes and the different demand components.

We find that an increase in government expenditures by 1 percent leads to an increase in goods imports by about 0.4 percent and to an increase in service imports by almost 0.5

percent. This implies that, ceteris paribus, an increase in government expenditure would also lead to a deterioration of the trade account. However, the ceteris paribus assumption in our context might lead to wrong policy conclusions if an increase (decrease) in government expenditure were to crowd out (crowd in) the private demand components. If this crowding-in/out effect were to prevail, an increase in government expenditures could bring about the opposite result.

The ambiguity of our results is in line with the findings of the literature; and, against this background, this paper provides an additional explanation for the commonly found ambiguous effect of government expenditures on import demand. We show that the ambiguity is, in part, the outcome of the compositional effect that an increase in government expenditures has on aggregate demand. The nature of this effect is not revealed when using a reduced-form equation. We find that higher government expenditures, ceteris paribus, lead to higher imports simply because the government consumes more from abroad in line with the import content of government consumption. However, when considering the compositional effect that fiscal policy measures have on overall demand – depending on the reaction of private demand – the opposite conclusion can also be derived.

Further research could determine the overall impact, i.e. the direct impact of a change in expenditure and the indirect impact through the reaction of private demand, that a change in government expenditure could have on the trade account of a particular country. For this purpose, a country-specific analysis of the link between fiscal policy measures and private demand would be appropriate.

1 Introduction

Little is known about the effects that a lasting change in government expenditures has on a country's external balance. There seems to be a consensus that lower expenditures and the concomitant improvement in the fiscal balance lead to an improvement in the current account. Empirical research so far, however, has led to ambiguous results:[4] Some empirical studies find that higher budget deficits lead to higher current account deficits; others prove the opposite or show no significant impact at all. A flaw of the models applied in this field of research seems to be that they estimate reduced-form equations, wherein different effects might counteract each other without showing the underlying causalities. The latter can only be revealed in a structural model. Furthermore, earlier studies suffer from the fact that econometric techniques that allow studying long-run equilibrium relationships between time series data were not yet developed.

This paper takes a fresh look at the empirical relationship between fiscal policy and the trade account by analysing the relationship between government expenditures and imports. Because trade account deficits are often at the heart of current account problems, a structural model of the trade account is an important step when modelling the impact of fiscal policies on the external balance. Within the trade account, we concentrate on imports because import demand is determined by domestic demand factors, while exports depend on external demand factors. To pin down the effects of fiscal policy, we estimate goods and service import equations on the basis of disaggregate demand variables. This implies that – in contrast to the conventional form of trade equations, which take total demand as an explanatory variable – we allow for all components of demand, i.e., private consumption, private sector investment, government expenditure, and exports, to exhibit different elasticities. For trade equations, the different elasticities of the aggregate demand components are essential because the import content of government consumption is generally lower than the import content of other demand components. Earlier studies took into account only the effect of different import contents of consumption, investment, and exports, but they do not discriminate between private and public demand.

The empirical analysis is based on annual panel data of the G7 countries for the years 1970 through 2002. We determine the cointegration relationship by a pooled mean group

[4] For a literature review, see Bussière, Fratzscher and Müller (2005) or Cavallo (2005).

estimation. This technique allows the intercepts, short-run coefficients and error variances to differ freely across countries, while the long-run coefficients are constrained to be the same for all cross sections. We are therefore able to account for cross country differences without losing the general message about the long-run relationships between import volumes and the different demand components. Based on this technique, we find that a change in government expenditure has a significant positive impact on both goods and service imports. This implies that an increase in government expenditure would ceteris paribus also lead to a deterioration of the trade account. However, we also show that the ceteris paribus assumption in our context might lead to wrong policy conclusions if an increase (decrease) in government expenditure was to crowd out (crowd in) the private demand components. If this crowding in/out-effect was only strong enough an increase in government expenditures could bring about the opposite result.

The paper is structured as follows. Section 2 presents some stylized facts on government expenditure and imports in our data sample. Section 3 explains the model and the estimation technique. Section 4 presents the empirical analysis and the results. Section 5 concludes.

2 The government sector and the trade account

Notable differences exist with respect to the import content of private consumption, government expenditure, investment and exports, respectively. Table 1 reports the import content of the different demand components for the UK in 2001 and for Germany, France, the UK and Italy in 1980, respectively. Despite some cross-country variation of the general level of import contents, it becomes obvious that compared to the other demand components government expenditure reveals the lowest import content across countries.

Table 1: Total import content of demand components

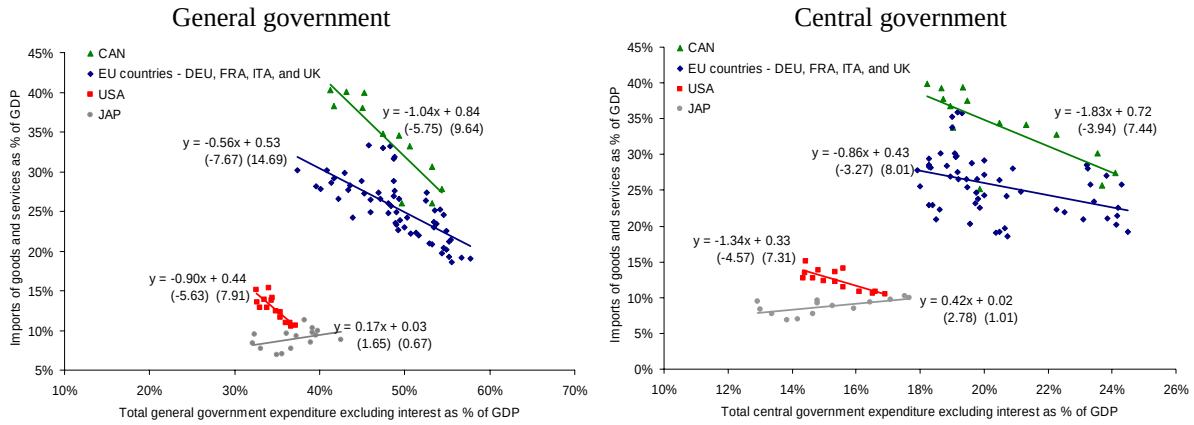
	Germany	France	Italy	United Kingdom	United Kingdom
	1980			1979	2001*
Aggregate expenditure	0.243	0.198	0.216	0.235	0.200
Private consumption	0.264	0.208	0.229	0.249	0.200
Government expenditure	0.134	0.060	0.064	0.097	0.132
Gross investment	0.244	0.267	0.261	0.372	0.318
Exports	0.272	0.201	0.241	0.235	0.224

Source: Giovannetti (1989), * Bank of England (2002).

Because of the different import contents of the demand components one can assume that the smaller the size of the government sector – measured as government expenditures in percent of GDP – the higher the import-to-GDP ratio. When the size of the government decreases the government sector uses less resources of the private sector and the composition of aggregate demand changes in favour of the private sector. Because the government sector has a smaller import content than the private sector, this shift in the composition of aggregate demand has a positive impact on import demand. If the government sector shrinks and the private sector increases, given the relatively low import content of government expenditure in comparison to private consumption, the demand for imports should increase.

To get an idea whether this proposition still holds when looking at G7 countries for recent years, we plotted government expenditure-to-GDP ratios against the import-to-GDP ratios for annual data from 1990-2004 (see Figure 1). The two panels show our findings for AMECO data and the general government (left panel) as well as for the IFS data and the central government (right panel).

Figure 1: Import ratio and government expenditure ratio for G7 countries (1990-2004)[a]



Source: International Financial Statistics, IMF (2005), AMECO, ECB (2005), own calculations.

a The left panel uses the AMECO data base and refers to general government. The right panel shows the IFS data, which was used in our calculations and refers to central government expenditure. T-statistics are given in brackets below the coefficients.

Figure 1 illustrates that an increase in the public expenditure ratio goes hand in hand with a decrease of the import ratio in all countries but Japan. With the exception of Japan, the t-statistics show that all coefficients are significant at the 1% level. The different behaviour of Japan could be related to the exceptional, decade-long stagnation of the economy. Figure 1

also reveals that – by and large – the negative relation between government expenditures and imports holds regardless of the degree of openness.

The negative relationship between general government expenditure and imports also exists for other countries outside the G7. As Table 2 shows, though the correlation coefficients differ across countries, the behavioural relationship seems to be relatively similar.[5] Again, most of the correlation coefficients are significant at the 1%-level. This could be taken as an indication that it might not be too far fetched to apply conclusions drawn from G7 countries to other countries where data problems prevent a more elaborate analysis such as the one conducted in this paper.

Table 2: Correlations of government expenditure ratio and import to GDP ratio (1990 – 2004)

Countries	Correlation
G7 Countries	
Canada	-0.88***
Germany	-0.57***
France	-0.35
United Kingdom	-0.59***
Italy	-0.94***
Japan	0.42****
United States	-0.84***
Other EU Countries	
Austria	-0.81***
Belgium	-0.89***
Denmark	-0.89***
Finland	-0.40*
Ireland	-0.91***
Luxembourg	-0.76***
Portugal	-0.44*
Netherlands	-0.91***
Sweden	-0.96***

Countries	Correlation
EFTA Countries	
Switzerland	0.06
Norway	0.43*
New EU Member States	
Cyprus	-0.98***
Czech Republic	-0.34
Estonia	-0.28
Hungary	-0.73***
Lithuania	-0.47**
Latvia	**0.17**
Malta	-0.68***
Poland	-0.41*
Slovakia	-0.62***
Slovenia	-0.43*

Source: Authors' own calculations.
Notes: *** indicates a 1% significance level, ** indicates a 5% significance level, * indicates a 10% significance level.

[5] Data is taken from the AMECO data base. For some countries, i.e., Canada, Germany, Sweden, Switzerland and the CEECs, the full data range is not available and correlations are calculated by applying the reduced data series. For Cyprus, Hungary, Malta and Slovenia only seven data points are available. The correlations, therefore, only provide an indication for the relation of the variables and must be interpreted cautiously.

Despite these relatively robust results, the correlations do not reveal the potential impact of a change in government expenditure on imports. For a policymaker it is important to know how a change in government expenditure affects imports, the current account and thus the external balance. We assess these effects in the following empirical analysis.

3 The model specification

3.1 Standard formulations of the trade account

Our analysis concentrates on the impact of fiscal policy on the trade account because trade account deficits are often at the heart of current account problems.[6] For all countries in the sample the trade account is quantitatively the most important of the three parts of the current account, though its share has been declining somewhat in recent years.

We base our analysis on an extension of the traditional model of the trade account. The basic trade model consists of an import and an export equation which relate import (*M*) and export (*X*) volumes to domestic (*Y*) and foreign (*Y**) real income and relative prices (*RP*).[7] Equations 1 and 2 show the export and import equations as given in the literature in their general and their log form:

Exports: $X_t = \gamma_0 Y_t^{*\gamma_1} RPX_t^{\gamma_2}$ in logs $x_t = \gamma_0 + \gamma_1 y_t^* + \gamma_2 rpx_t$ (1)

Imports: $M_t = \delta_0 Y_t^{\delta_1} RPM_t^{\delta_2}$ in logs $m_t = \delta_0 + \delta_1 y_t + \delta_2 rpm_t$. (2)

RPX and *RPM* are the relative prices, γ_1 and δ_1 represent the income elasticities and γ_2 and δ_2 the price elasticities of exports and imports, respectively. Domestic real income (*Y*) is equivalent to real GDP, which equals the sum of the demand components, i.e., private consumption, public spending, private investment and net exports. Foreign real income (*Y**) represents the total income of the rest of the world and can not easily be decomposed into demand components. Relating import volumes to total real income implicitly assumes that the import content and the import elasticity are the same for all demand components.

[6] Recent literature also points out a reversed causality between the current account and fiscal policy. In this respect Baker (2004) finds that increased foreign indebtedness may contribute to an erosion of the tax base. We do, however, focus on the impact that fiscal policy has on the current account through the demand side.

3.2 Import equations and expenditure components

Earlier research showed that import demand is not only determined by the level of income and final expenditure but also by the composition of expenditure and the import content of the different components. Abbott and Seddighi (1996), Giovannetti (1989) and Mohammad and Tang (2000), for example, estimate import equations by taking disaggregated demand/expenditure components into account. They divide total demand into consumptive expenditure, investment expenditure and exports. The results show that the elasticities of the different demand components differ significantly.[8]

To our knowledge, so far the existing literature assumed that at least private consumption and government expenditure reveal common elasticities. Hence, the impact of fiscal policy measures on import demand has not been taken into account by the literature. Our model, however, allows us to gauge the impact of a change in public spending on imports because we disaggregate domestic real income into its demand components and separately consider private consumption and government expenditure.

The extended import equation distinguishes between private consumption (C), private investment (I), government expenditure (G) and exports (X):

$$M_t = \partial_0 C_t^{\partial_1} I_t^{\partial_2} G_t^{\partial_3} X_t^{\partial_4} RPM_t^{\partial_5} \quad \text{in logs} \quad m_t = \partial_0 + \partial_1 c_t + \partial_2 i_t + \partial_3 g_t + \partial_4 x_t + \partial_5 rpm_t . \quad (3)$$

Equation (3), thus, permits divergent import elasticities for private consumption and government expenditure[9] because the import content of government expenditure is generally lower than that of private consumption (see Section 2). The major parts of government expenditures are public wages and social expenditures, which have a low or marginal import

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[7] Recent research in this field has been publisher by Hooper et al. (2000) and Marquez (2002). For a more general discussion of the traditional trade model see Goldstein and Khan (1985).

[8] Abbott and Seddighi (1996) apply a likelihood ratio test to see whether the long-run elasticities estimated by a Johansen procedure could be restricted to be the same for all demand components. They had to reject the restriction.

[9] A more detailed analysis could consider public consumption and public investment demand separately. These two components of public expenditure can be expected to reveal major differences in terms of import content. Due to limitations in the availability of consistent data for the empirical analysis disaggregating public expenditure was not possible.

content.[10] Equation (3) shows that the impact of fiscal policy on the trade account depends on the direct effect of government expenditure on imports but also on the indirect effects that fiscal policy measures might have on the other demand components, i.e., private consumption and private sector investment.

3.3 Specification of the empirical model

Our trade volume equations are an extension of the export and import equations 1 and 2 that are separated into trade volume equations for goods (equations 4 and 6) and services (equation 5 and 7). First, the following four conventional trade volume equations are estimated in their log form:[11]

Goods exports: $$xg_t = \gamma_0 + \gamma_1 yg_t^* + \gamma_2 rpxg_t, \quad (4)$$

Service exports: $$xs_t = \theta_0 + \theta_1 ys_t^* + \theta_2 rps_t, \quad (5)$$

Goods imports: $$mg_t = \delta_0 + \delta_1 y_t + \delta_2 rpmg_t, \quad (6)$$

Service imports: $$ms_t = \psi_0 + \psi_1 y_t + \psi_2 rps_t, \quad (7)$$

Then import volume equations for goods (equation 6) and services (equation 7) are extended along the lines described in the previous section:

Extended form of goods imports: $$mg_t = \partial_0 + \partial_1 c_t + \partial_2 i_t + \partial_3 g_t + \partial_4 x_t + \partial_5 rpmg_t, \quad (8)$$

Extended form of service imports: $$ms_t = \vartheta_0 + \vartheta_1 c_t + \vartheta_2 i_t + \vartheta_3 g_t + \vartheta_4 x_t + \vartheta_5 rps_t. \quad (9)$$

The estimation considers annual data for the G7 countries from 1970[12] through 2002. A full description of the variables is given in Appendix 1.

[10] Data from the OECD 1990 input-output table for Germany reveals that about 50% of government expenditure is spent on inputs from producers of government services which in turn have human labour as there only input. Without having a thorough look at the components of government consumption it seems reasonable to conclude that a major portion of government demand is satisfied by domestic output.

[11] In contrast to Driver and Wren-Lewis (1998), a time trend is not included. This, however, does not change the estimation results. In a first step, Driver and Wren-Lewis also estimated the elasticities without considering the time trend and in a second step estimated the time trend while applying the coefficients as derived in the first step (Driver and Wren-Lewis (1998), p. 119).

[12] Due to missing data points, equations (5), (7) and (9) are estimated using data from 1977 through 2002.

For the conventional trade equations 4 to 7, domestic income or demand (y) are expected to have a positive impact on import volumes (ms) or (mg). Likewise, export volumes (xg) or (xm) are expected to increase with foreign income (yg^*) or (ys^*).[13] As discussed by Marquez (2002), economic theory postulates the income elasticity to be equal to one provided it is assumed to be constant. However, various empirical studies show that the estimated coefficient deviates from one but remains close to one.[14] In the present analysis income elasticities are therefore also expected to be close to one across all sample countries. Likewise, it is assumed that the sum of the demand elasticities (i.e. for consumption, investment, government expenditure and exports) should also be equal or close to one in the extended trade equations (8) and (9). Since demand decreases as prices increase, the coefficients of relative prices are expected to be negative in all six equations.

4 Empirical analysis and results

When estimating the trade volume equations the analysis follows the approach by Driver and Wren-Lewis (1998). Panel unit root tests are applied to test for stationarity of the time series. Almost all variables are integrated of order one. Because of this result panel cointegration techniques are applied to a panel of G7 countries to estimate the elasticities of the export and import volume equations in the conventional form as well as in the extended form in the case of import volumes. Furthermore, the Johansen procedure is applied to each country individually to verify whether the common coefficients derived from the panel analysis appropriately reflect the individual country data.[15]

The details of the estimations as well as the results are presented in the following subsections.

[13] The world demand for goods exports (yg^*) is proxied by world merchandise trade, which only includes goods trade. Similar data is not available for services. Hence, the world demand for service imports (ys^*) is proxied by world real GDP.
[14] See, for example, Cline (1989), Caporale and Chui (1999), Hooper et al. (2000) and Marquez (2002).
[15] Comparing the country-by-country estimation with the results of the panel cointegration only provides an eyeball test for the adequacy of the common coefficient from the pooled estimation. The analysis is refined by a pooled mean group estimator and a mean group estimator which allow a quantitative assessment of the relevance of the common coefficient for the individual countries by applying a Hausman test.

4.1 Panel unit root test

Multiple methods for unit root tests as well as cointegration analyses have been developed for panel data in the recent past. These panel unit root tests are mostly based on estimating some version of a standard dynamic model for a panel, such as

$$y_{it} = \rho y_{it-1} + \delta_0 + \delta_1 t + \eta_i + \nu_t + \varepsilon_{it} \tag{10}$$

and testing whether the coefficient ρ is equal to one. The subscript i = (1, 2, ..., N) distinguishes the N countries included in the panel. Examples for such tests are Levin, Lin and Chu (2002) and Breitung (2000). Other procedures, for example, Im, Pesaran and Shin (2003), are based on averages of the individual unit root test statistics. They recommend, for example, to apply the Dickey-Fuller (DF) and the augmented Dickey-Fuller (ADF) tests to the individual time series and to calculate one common test statistic from the individual t-tests.

By determining their test statistics based on the full information contained in the data panel the techniques proposed by Levin, Lin and Chu (2002) (LLC) and Breitung (2000) best offers the most suitable asymptotic properties in the case of medium size panels, i.e., an equivalent extension of the cross section and the time series dimension. We therefore apply both methods to test the relevant time series for stationarity. LLC and Breitung test the null hypothesis that each individual time series in the panel is integrated versus the alternative hypothesis that all individual time series are stationary. Both tests are based on the following pooled ADF equation

$$\Delta y_{it} = \delta y_{it-1} + \sum_{L=1}^{p_i} \theta_{it} \Delta y_{it-L} + \alpha_{0i} + \alpha_{1i} t + \varepsilon_{it},$$

where a common $\delta = \rho - 1$ is assumed. The null of H_0: $\delta = 0$ under the assumption that $\delta_i = \delta$ for all i is tested against the alternative hypothesis, H_a: $\delta < 0$ for $\delta_i = \delta$ for all i. The tests allow for country specific intercepts (α_{0i}) and the trend coefficients (α_{1i}). However, while the LLC is based on a technique which removes autocorrelation as well as the deterministic components, i.e., individual intercept and individual trend, when making the relevant standardisations the test statistic proposed by Breitung is calculated by removing the autoregressive component but not the deterministic portion of the ADF equation. The results of the LLC and the Breitung tests are given in Table 3.

Table 3: Results of the Levin/Lin unit root tests

		LLC H_0: $\delta = 0$ Critical probability	Breitung H_0: $\delta = 0$ Critical probability
Relative price of exported goods	*rpxg*	0.0579	0.0584
Relative price of imported goods	*rpmg*	0.8771	0.4112
Relative price of services	*rps*	0.3403	0.9864
Export goods	*xg*	0.0117	0.9861
Export services	*xs*	0.0059	0.9992
Import goods	*mg*	0.7915	0.7326
Import services	*ms*	0.0473	0.8527
World trade volume	*yg**	0.9836	0.2974
World real GDP	*ys**	0.0001	0.9760
Real GDP	*y*	0.0076	0.6657
Private consumption	*c*	0.0000	0.8933
Government consumption	*g*	0.0001	0.2520
Private investment	*i*	0.3343	0.6828
Export	*x*	0.0113	0.8442

Source: Authors' own calculations.
Note: The ADF specification takes individual intercepts but no trend term into account.

According to the Breitung test statistic the null of nonstationarity can not be rejected for all data series but the relative price for exported goods. The results generated by the LLC are somewhat weaker. Alternative test procedures e.g. the unit root test by Im, Pesaran and Shin (2003) confirm that all but the *rpxg* series possess a unit root and thus support the outcome of the Breitung test. Cointegration techniques are, therefore, the appropriate tool to estimate the trade volume equations.

4.2 Panel cointegration test

The available techniques for panel cointegration tests are Engle/Granger-like residual based tests. Similar to single time series, these approaches test the residuals from the estimation for stationarity. If the estimated residuals are stationary a linear combination of the time series included in the estimation exists so that the resulting time series is a stationary process. The time series are thus cointegrated. As in the case of single time series, this form of cointegration test does not allow to test for the number of cointegrating relationships among the variables. In cases where more than one cointegration relationship exists and/or not all variables are part of the cointegration space, these tests only show that some combination of the included variables reveals stationary residuals. This means that some of the variables but

not necessarily all of them are cointegrated. Therefore, the trace and the maximum eigenvalue statistics suggested by Johansen (1988) are applied on a country by country basis for all G7 countries. Since these tests reveal in almost all cases of the trade volume equations that all relevant variables are part of a single cointegration equation, it is reasonable to apply the available residual based panel cointegration tests.[16]

For the following estimations, residual based panel cointegration tests as suggested by Pedroni (1999) and Kao (1999) are employed. Both assume homogenous slope coefficients across countries. This is in line with the purpose of our analysis, namely deriving a general relationship between government expenditure and import volumes. Pedroni as well as Kao apply the null hypothesis of "no cointegration".

Kao (1999) tests the residuals $\hat{\varepsilon}_{it}$ of the OLS panel estimation by applying DF- (equation 11) and ADF- (equation 12) like tests.

$$\hat{\varepsilon}_{it} = \rho\hat{\varepsilon}_{it-1} + \nu_{it} \tag{11}$$

$$\hat{\varepsilon}_{it} = \rho\hat{\varepsilon}_{it-1} + \sum_{j=1}^{p}\varphi_j \Delta\hat{\varepsilon}_{it-j} + \nu_{itp} \tag{12}$$

The null hypothesis of no cointegration i.e. H_0: $\rho = 1$ is tested against the alternative hypothesis of stationary residuals i.e. H_a: $\rho < 1$. Pedroni (1995) suggest a Phillips-Perron-type test, which implies less strict assumptions with respect to the distribution of the error terms than the DF and ADF tests do. The results of the cointegration tests are given in Table 4. They show that the null hypothesis of no cointegration can be rejected at conventional significance levels in all cases. These results combined with the outcome of the Johansen procedure indicate that the variables included in the different trade volume equations are cointegrated and that one cointegration relationship exists.

[16] The results of the Johansen-Tests can be requested from the author. The fact that the relative price does not appear as separate cointegration relationship might indicate that the time series is in fact not stationary. This supports the decision to apply cointegration analysis despite the panel unit root test does not support the null of a unit root for these variables.

Table 4: Panel cointegration tests

	Goods exports	Service exports	Goods imports	Service imports	Extended goods imports	Extended service imports
Kao (1999)[1]						
DF-roh	-2.14	-3.5994	-0.9768	-2.6577	-3.5242	-5.4910
	(0.0162)	(0.0002)	(0.1643)	(0.0039)	(0.0002)	(0.000)
DF-t	-1.421	-2.2892	-0.6161	-1.7807	-2.1827	-3.4352
	(0.0777)	(0.0110)	(0.2689)	(0.0375)	(0.0145)	(0.0003)
DF-rho*	-6.4188	-8.8235	-4.8909	-7.2725	-8.6555	-10.5543
	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)
DF-t*	-1.9633	-2.6689	-1.5068	-2.3184	-2.6025	-3.6898
	(0.0248)	(0.038)	(0.0659)	(0.0102)	(0.0046)	(0.0001)
Kao (1999)[2]						
ADF	-1.5318	-2.0591	-1.1518	-2.0725	-2.3835	-3.1681
	(0.0628)	(0.0197)	(0.1247)	(0.0191)	(0.0086)	(0.0008)
Pedroni (1995)[3]						
PC_1	-11.0199	-14.1563	-8.7730	-12.4306	-13.1817	-17.6491
	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)
PC_2	-10.8347	-13.8814	-8.6391	-12.1892	-12.9805	-17.3063
	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)

Source: Authors' own calculations.
Notes: p-values are given in parentheses.
[1] The DF test statistics given above are analogous to the parametric Dickey-Fuller test for nonstationary time series. The DF-rho and DF-t statistics assume strict exogeneity of the regressors with respect to errors and no autocorrelation. DF-rho* and DF-t* statistics are based upon endogenous regressors. Note that these tests depend on consistent estimates of the long-run variance-covariance matrix to correct for nuisance parameters once the limiting distribution has been found.
[2] The ADF test is analogous to the parametric Augmented Dickey-Fuller test for nonstationary time series.
[3] PC1 and PC2 are the non-parametric Phillips-Perron tests.

4.3 Estimation of trade volume equations

Trade elasticities are estimated by applying the pooled mean group (PMG) estimator proposed by Pesaran et al. (1999). The long-run relationships are estimated in a pooled as well as in a country-by-country setting. The cross-country average of the coefficients from the latter is the mean group (MG) estimator. A Hausman test allows assessing whether slope homogeneity exists among cross sections and thereby reveals whether the PMG estimator provides a consistent and efficient estimation for the coefficients across all countries.

The estimation is based on the following re-parameterization of the standard autoregressive distributed lag (ARDL) model

$$\Delta y_{it} = \phi_i y_{i,t-1} + \beta_i' x_{it} + \sum_{j=1}^{p-1} \lambda_{ij}^{*} \Delta y_{i,t-j} + \sum_{j=0}^{p-1} \delta_{ij}^{*\prime} \Delta x_{i,t-j} + \mu_i + \gamma_i t + \varepsilon_{it},$$

where y_i and x_i are a vector of observations on the dependent variable, i.e., trade volume, and a vector of explanatory variables, i.e, relative price and income, for country i, respectively. μ_i represents the country specific fixed effect, γ_i is the individual time trend coefficient and ε_i stands for the country specific error term. The long-run relationship between y_i and x_i is given by

$$y_{it} = -\left(\beta_i' / \phi_i\right) x_{it} + \eta_{it},$$

where $-\left(\beta_i' / \phi_i\right)$ is the long-run coefficient, i.e., the respective elasticity, η_{it} is the error term and all other variables are defined as given above.

To address the problem of cross sectional correlation, demeaned data[17] are used in the case of all import equations. In the case of export equations a time trend is considered instead. This is due to the fact that world income is common for all cross sections and can not be demeaned.

Table 5 shows the estimation results. The country sample included in the estimation is adjusted where necessary to include only those countries for which the data allows to determine a long-run relationship.[18] The p-values of the joint Hausman test[19] reveal that for the countries included in the estimations the null of slop homogeneity can not be rejected.

[17] Demeaned data is constructed by subtracting the cross-sectional average of a respective variable from each data point of the respective cross section: $\bar{y}_i = T^{-1} \sum_{t=1}^{T} y_{i,t}$

[18] The fact that no reasonable cointegration relationship can be established for example for France in the case of service export might be a country-specific problem that, for example, forced Driver and Wren-Lewis (1998) to assign values for the elasticities in such cases. Our analysis does not intend to determine country-specific elasticities but general results and the number of cross sections is large enough so that the exclusion of one or two countries from the parts of the analysis does not harm the general propositions drawn from the estimation results.

[19] The joint Hausman test assesses the null hypothesis of slop homogeneity against the alternative hypothesis of heterogeneous slope coefficients across countries

Table 5: Cointegration estimation of conventional trade volume equations[20]

	Goods export (equation (4)) PMGE[1)2)]	Service exports (equation (5)) PMGE[1)2)3)]	Goods imports (equation (6)) PMGE[4)]	Service imports (equation (7)) PMGE[2)4)]
Price elasticity	-0.849**	-0.726**	-0.313**	-1.263**
	(-8.647)	(-3.500)	(-3.076)	(-15.921)
Income elasticity	0.906**	1.018**	1.953**	1.316**
	(36.395)	(3.572)	(9.896)	(56.190)
Joint Hausman test	0.66	0.89	0.31	0.94

Source: Authors' own calculations.
Notes: t-statistics are provided in parentheses.
* and ** denote statistical significance at a 5% and 1% level respectively. t-statistic are given in parentheses.
[1)] Estimation equation includes time trend.
[2)] Japan is excluded from the estimation.
[3)] France is excluded from the estimation.
[4)] The estimation is based on demeaned data.

Comparing the results of the estimation above with those generated by Hooper et al. (2000) and Driver and Wren-Lewis (1998) shows that the estimated coefficients are in the range of those received from single time series analysis. Hooper et al. (2000) estimate long-run trade elasticities for the G7 countries. Their results reflect the fact that income elasticities usually deviate from unity and that price elasticities vary significantly among countries. Driver and Wren-Lewis (1998) use the Johansen approach and vector error correction estimates in order to determine the trade volume elasticities for the G7 countries on a country by country basis. Their results also reflect the fact that the estimates for income elasticities for the G7 countries deviate from unity. This can be inferred from their explanations and from the fact that almost all coefficients that the authors finally use for other estimations were generated through constrained estimations or even imposed without taking the original estimation output into account. The results of the studies of Hooper et al. and Driver and Wren-Lewis are given in Appendix 2 in Table A1 and Table A2, respectively.

In the next step, the extended form of the import volume equations (equation 8 and 9) are estimated to analyze the effects of government expenditure on foreign trade. The results of the PMG estimation are summarized in Table 6.

As in the conventional trade equations, the relative price variable is significant and has the expected sign. All demand variables (but private sector investment in the service import

[20] The dependent variable is the log of the respective trade volume.

equation) are significant. They show a positive effect on goods and service imports. The magnitude of the elasticities differs among the demand components. This confirms that the composition of demand matters for the import equation and that using a single aggregate demand variable might distort the result. In the case of services, private investment does not have a significant impact on import volumes and government expenditure reveals the smallest elasticity among the remaining demand components.

One might argue that these results might be flawed because of multicollinearity, in particular, between government spending and private consumption. The practical consequence of multicollinearity could be that confidence intervals tend to be much wider, leading to the acceptance of the null hypothesis more readily. Hence, the t-ratios might be interpreted as statistically insignificant even though in reality they are significant. Because the t-statistics in Table 6 show that all variables (except for private sector investments in the service import equation) are significant, from a statistical point of view multicollinearity is not a concern.

Table 6: Cointegration estimation of extended import volume equations[21]

	Goods imports (equation 8) PMGE[2) 3) 4)]	Service imports (equation 9) PMGE[2) 4) 5)]
Price elasticity	-0.665**	-1.592**
	(-5.015)	(-6.747)
Private consumption	1.102**	1.433**
(ln C)	(3.481)	(1.916)
Government expenditure	0.392*	0.491**
(ln G)	(1.762)	(2.485)
Private sector investments	0.427**	0.030
(ln I)	(5.152)	(0.076)
Exports	0.435**	0.503**
(ln X)	(4.156)	(1.972)
Joint Hausman test	0.12	0.22

Source: Own estimations.
* and ** denote statistical significance at a 5% and 1% level respectively. t-statistic are given in parentheses.
1) Estimation equation includes a time trend.
2) Japan is excluded from the estimation.
3) France is excluded from the estimation.
4) The estimation is based on demeaned data.
5) The coefficients of private consumption and private sector investment are not restricted to be homogenous across countries.

[21] The dependent variable is the log of the respective trade volume.

Our empirical results show that an increase in government expenditures has a positive impact on total import demand. A lasting increase in government expenditure of one percent will lead to an increase of demand for goods and service imports of 0.4 and 0.5 percent, respectively. An increase in public spending will, thus, ceteris paribus lead to a deterioration of the trade account simply because the government consumes more from abroad in line with its import content. Because of the relative weight of the trade account in the current account the current account would improve if government expenditure were reduced.

However, our results need to be interpreted with caution because the ceteris-paribus interpretation of the coefficients is problematic in our context as an increase (decrease) in government expenditure is likely to crowd out (crowd in) the private demand components. Other empirical studies have shown that an increase in government expenditure might crowd out private sector investment while private consumption is likely to increase as public expenditure rises.[22] If an increase in government expenditure crowds out private investment but positively impacts private consumption, the impact on import volumes becomes less predictable. If public expenditure and private consumption replace private investment – due to the combination of a high elasticity of private consumption and the low elasticity of public expenditure – the decline in import demand due to the slowdown in private investment might or might not be compensated by the surge in import demand caused by the increase in public expenditure and private consumption. The overall effect of such a demand shift on goods imports depends on the relative size of the change in public expenditure and private consumption. In the case of service imports the effects are more predictable. According to our results the increase in government expenditure and the related rise in private consumption cause an increase in service imports while the decrease in private investment does not impact the service account. An increase in government expenditure would thus lead to a deterioration of the service account.

Because the goods account is more sizeable than the service account,[23] it can be expected that the effect coming from the goods account overrides the effect stemming from the service account. If this is the case, the overall impact of an increase in government expenditure of the trade account depends on the reaction of private consumption and private investment on the expansion of the government sector.

[22] See, for example, Karras (1994) and Blanchard and Perotti (2002).
[23] In the case of the G-7 countries, service imports are less than one third of the size of goods imports.

Overall the results of our estimation provide insights regarding the direct effect of a change in government expenditure on import demand but its indirect effects are less clear. Government expenditure reveals a positive elasticity with respect to goods imports and service imports. An increase in government expenditure, ceteris paribus, causes an increase in import volumes. However, the indirect effects of fiscal policy measures caused by the reaction of private consumption and private investment to a change in public expenditure are less clear-cut. Since the empirical literature does not provide unanimous evidence regarding the impact that fiscal policy measures have on private demand,[24] the interpretation of our results depends on the interaction between the public and the private sector.

5 Summary and conclusion

This paper analyzes the empirical relationship between fiscal policy and the trade account. It shows that fiscal policy matters for the trade account and sheds light on how fiscal policy affects the trade account. Research prior to this paper did not take into account the fact that the components of private and public demand in the import equation exhibit different elasticities. Using pooled mean group estimation for annual panel data of the G7 countries for the years 1970 through 2002, we find that an increase in government expenditures has a significant positive impact on both goods and service imports. An increase in government expenditures by 1 percent leads to an increase in goods imports by about 0.4 percent and to an increase in service imports by almost 0.5 percent. This implies that, ceteris paribus, an increase in government expenditure would also lead to a deterioration of the trade account. However, the ceteris paribus assumption in our context might lead to wrong policy conclusions if an increase (decrease) in government expenditure was to crowd out (crowd in) the private demand components. If this crowding-in/out effect was only strong enough an increase in government expenditures could bring about the opposite result.

[24] Considering the impact of government expenditure on consumption and investment separately, Blanchard and Perotti (2002) reveal that fiscal expansion has a positive impact on consumption and a negative impact on investment. Fatás and Mihov (2001), however, find that consumption increases as a response to a positive expenditure shock while investment is not affected significantly. Karras (1994) finds evidence that private consumption and government spending are complementary: private consumption decreases as government expenditures are cut.

The ambiguity of our results is in line with the findings of the literature;[25] and against this background, this paper provides an additional explanation for the commonly found ambiguous effect of government expenditures on import demand. We showed that they are, in part, the outcome of the compositional effect that an increase in government expenditures has on aggregate demand. The nature of this effect would not have been revealed when using a reduced-form equation. We saw that higher government expenditures, ceteris paribus, lead to higher imports simply because the government consumes more from abroad in line with the import content of government consumption. However, when considering the compositional effect that fiscal policy measures have on overall demand – depending on the reaction of private demand – the opposite conclusion can also be derived.

This study reveals that a difference between the trade elasticities of private and public demand exists. Further research could determine the overall impact, i.e. the direct impact of a change in expenditure and the indirect impact through the reaction of private demand, that a change in government expenditure could have on the trade account of a particular country. For this purpose, a country-specific analysis of the link between fiscal policy measures and private demand would be appropriate.

[25] For example Erceg, Guerrieri and Gust (2005), Lane and Perotti (1998) or Baxter (1995) who analyse the impact of fiscal policies on the trade account find divergent effects. Analysing the relation between fiscal deficit and current account deficit, studies by Berenheim (1988), Bussière, Müller and Fratzscher (2004), Normandin (1999), Piersanti (2000), Enders and Lee (1990), Dewald and Ulan (1990) as well as Kim and Roubini (2004) reveal contradicting results. Some of the studies find a positive, some a negative and some find no significant relation between the two deficits at all.

Appendix 1: Data description and sources

All estimations are carried out with annual data for the G7 countries (Japan, the United States, Canada, the United Kingdom, France, Italy and Germany). Data series are taken from the IMF's international financial statistic (IFS), the OECD's main economic indicators (MEI) data bases and the IMF's Direction of Trade Statistics (DOTS).

Data for the estimation of trade equations

For this part of the analysis data series for the G7 countries and world aggregates or OECD data for world variables covering the period from 1970 through 2002 are considered. The trade equations (equations 4 to 9) include the following variables:[26]

Variable	Explanation	Data source and transformation
XG	Goods export volumes	Export volumes (IFS line 72) are turned from an index into constant price series using the 1995 average for merchandise exports in US$ (IFS line 78aa) converted into domestic currency using the 1995 average for the exchange rate (r). The series are then turned into a volume series by deflating by PC.
XS	Service export volumes	Service credits in US$ (IFS line 78ad) are converted into domestic currency using the actual exchange rate (r). The series are then turned into a volume series by deflating by PC.
MG	Domestic goods import volumes	The import volume FOB series (IFS line 73) is turned from an index into a constant price series using 1995 average by multiplication with merchandise exports in US$ (IFS line 78ab) and is converted into domestic currency using the 1995 average for the US$ exchange rate (r).
MS	Domestic service import volumes	Service debits in US$ (IFS line 78ae) are converted into domestic currency using the actual US$ exchange rate (*r*) and into a volume series by deflating by *PCW* after converting *PCW* into domestic currency terms using *EFEX*.
*YG**	World income relevant for goods export demand (equivalent to world trade volume)	OECD. *YG** as world trade volume is proxied by total world exports in US$ at current prices (IFS line 70), deflated using *WPXG*.
*YS**	World income relevant for service export demand (equivalent to world real GDP)	OECD. Total OECD GDP at constant market prices in US$.
Y	Domestic real GDP	IFS line 99b and deflated by *PY*.
C	real private consumption	IFS line 96f and deflated by *CP*
I	Real private sector investment	IFS line 93i plus IFS line 93e and deflated by *PY*.
G	real government expenditure	IFS line 91f and deflated by *PY*.
X	real exports	IFS line 90c and deflated by *PY*.
PC	domestic consumer price index in domestic currency	IFS line 64
PCW	world consumer price index	MEI of the OECD
PXG	domestic export prices	Export prices index (IFS line 76) are used in the case of Japan, UK and the US. *PD* as the domestic prices index in domestic currency is given by wholesale prices (IFS line 63).
WPXG	world export prices in US$	unit value of world exports in US$ (IFS line 74). For Canada, France, Germany and Italy this is an export unit value index (IFS line 74).

[26] The upper case abbreviations for the variables correspond to the lower case equivalents in the equations as given in the text. However the upper case stands for absolute values while the equations are given in logs.

Variable	Explanation	Data source and transformation
PY	domestic GDP deflator	IFS line 99bi
r	nominal US$ exchange rate	IFS line rf
EFEX	nominal effective exchange rate	Calculated from the exchange rates (*r*) and the bilateral trade weights (exports plus imports (lines 70 and 71 of the direction of trade statistics)) of the G7 countries and their 39 largest trading partners (including the G7 countries themselves).
RPXG	relative price for goods exports	(WPXG*r)/PXG.
RMPG	relative price for goods imports	(WPXG*r)/PD
RPS	relative price for service exports and service imports	PCW/ (PC*EFEX)

Appendix 2: Single time series estimations of trade elasticities for the G7 countries

Table A1: Long-run income and price elasticities estimated by Hooper et al. (2000)

	Income elasticities		Price elasticities	
	Export	Import	Export	Import
Canada	1.1*	1.4*	-0.9*	-0.9*
France	1.5*	1.6*	-0.2	-0.4*
Germany	1.4*	1.5*	-0.3	-0.06*
Italy	1.6*	1.4*	-0.9*	-0.4*
Japan	1.1*	0.9*	-1.0*	-0.3*
United Kingdom	1.1*	2.2*	-1.6*	-0.6
United States	0.8*	1.8*	-1.5*	-0.3*

Source: Hooper et al. (2000), p. 8.
* Statistically significant at a 5% level.

Table A2: Income and price elasticities estimated by Driver and Wren-Lewis (1998)

	Income elasticities		Price elasticities	
	Export	Import	Export	Import
Canada	1.00^{++}	0.62	-0.83^{++}	-0.68
France	1.00^{+}	1.00^{++}	-0.67^{+}	-0.50^{++}
Germany	1.00^{+}	1.00^{+}	-1.15^{+}	-0.82^{+}
Italy	1.01^{+}	1.00^{+}	-0.44^{+}	-0.71^{+}
Japan	0.91	1.00^{+}	-1.36	-0.33^{+}
United Kingdom	0.91	1.00^{+}	-1.26	-0.72^{+}
United States	1.12	1.50^{+}	-0.96	-0.40^{+}

Source: Driver and Wren-Lewis (1998), pp. 41, 43.
$^{+}$ indicates that the coefficient comes from a constrained ECM or a constrained Johansen estimation.
$^{++}$ indicates that the coefficient was imposed by the authors.

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