

On the Causal Relationship between Trade Openness and Government Size: Evidence from OECD Countries^{*}

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Abstract

The *compensation hypothesis* predicts a positive causation from international economic openness to the size of the public sector, as governments step in to perform a risk mitigating role to counterbalance the increasing exposure to external risk and the economic dislocations caused by growing international openness. We use time series data from 22 OECD countries over the period 1955-2003 and examine the statistical significance of both long-run and short-run causality channels in each country separately. Our findings fail to provide an overwhelming support for this hypothesis, with only 5 countries showing some evidence in its favour.

Keywords: globalisation; compensation hypothesis; trade openness; government size; Granger causality; cointegration; vector error correction

JEL Classification: F15, H11, H5

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1. Introduction

In recent years, economists and political scientists have increasingly focused on understanding how the increase in the degree of international economic openness in a country may influence the size of its government. One of the dominant views that emerges from this literature — particularly amongst economists, see for instance Alesina and Perotti (1997) — is reflected in the so called *efficiency hypothesis* which suggests that economic globalisation inevitably reinforces the need to roll back governments' spending programmes on account of: (i) the fact that the distortionary effects of public expenditure and the taxation necessary to finance it harm the international competitiveness of national firms and industries, and (ii) the threat of international relocation of increasingly mobile capital, firms and jobs that undermine the revenue raising ability of governments.

This conventional wisdom is, however, somewhat at odds with the concomitant occurrence of two major trends that characterise the post World War II period, namely: (i) the process of international economic integration, which has resulted in rapid and progressive increases in cross border flows of goods, services, capital and technology, and (ii) the expansion of government sectors. In his seminal contributions, Rodrik (1997a, 1998) uses cross-country data to investigate the nature of the relationship between 'trade openness' and 'government size'. He measures these respectively by $(\text{exports} + \text{imports})/\text{GDP}$ averaged over the period 1980-1989 and $(\text{government consumption})/\text{GDP}$ averaged over the period 1990-1992 and finds that there is a strong positive causation from the former to the latter. Rodrik then argues that, in contrast to the view that regards markets and governments as substitutes, this evidence suggests that there may be some degree of complementarity between them. In particular, he suggests that the causal relationship between trade-openness and government-size can be explained by what has become known as the '*compensation hypothesis*' — according to which the increased volatility brought about by growing exposure to (and dependence on) developments in the rest of the world creates incentives for government to provide social insurance (against internationally generated risk and economic dislocations).

There is now a large body of empirical literature on this topic. Nevertheless, there is still a great deal of ambiguity about the effect of economic globalization on the size of the public sector.¹ According to Rodrik, if the compensation hypothesis holds — that is, provided that

¹ Brady et al. (2005) put this point forcefully. Cameron (1978) was amongst the first to point out the existence of a positive relationship between openness and government-size, suggesting that more open economies, due to higher rates of industrial concentration, were more likely to develop labour movements which exert stronger pressure on governments to provide social transfers. The burst of research on this issue mainly occurred after Rodrik revisited the issue, and since then many studies have empirically investigated the relationship between various measures of economic globalization (of which trade-openness is only one dimension) and the size of government or the structure and composition of the public sector. An exhaustive survey of this literature is beyond the scope of this paper (see footnote 4 below for some examples). Studies focussing on different facets of the size and composition of the public

openness does increase exposure to external risk and governments do fulfil a risk mitigating role — we ought to find a positive causality from trade-openness to government-size over time. This causality is reinforced by the argument recently put forward by Epifani and Gancia (2009a) who propose an additional channel that works through a country's terms-of-trade effect. They too use Rodrik's definition of trade-openness and government-size and provide supportive evidence from a sample of both cross section and panel time-series data that includes 143 countries.

Most existing empirical studies that focus on this relationship are based on data that pool many countries together. The regression analysis of pooled cross-section time-series data will only shed light on the *average* behaviour. However, even a casual observation of individual country performances suggests that the relationship between government-size and openness might be highly country specific.² In order to gain deeper insights as to whether a specific country exhibits a systematic pattern that is consistent with one of these rival hypotheses we need robust time series evidence for that country. Since evidence based on data that pools a large number of countries can only tell us what is likely to have happened on average, it renders the strength of the policy conclusions that can be drawn from such evidence somewhat ambiguous.³

Our aim in this paper is to re-examine a narrower aspect of this relationship empirically, using Rodrik's definition of trade-openness and government-size but within a more robust time-series approach. More specifically, we focus on the existence, in an individual country, of a causal effect — in a purely time series context following an updated version of Granger's original methodology — from trade-openness to government-size. We apply the causality test to 22 OECD countries individually using annual time series data over the last five decades. Our results are not as encouraging as those of Rodrik and Epifani and Gancia; indeed, we find that individual country time series data tell a different story: only for few countries in our sample do we find some evidence of the existence of a positive and significant causation from trade-

sector are reviewed in Persson and Tabellini (1999). Gemmell et al. (2008) offer an excellent summary of the existing findings in the literature and Brady et al. (2007) present a general discussion of the impact of globalisation.

² For instance, since 1987 the Republic of Ireland has pursued international competitiveness policies consistent with the efficiency hypothesis. These policies, which led the pre 2009 Irish economy to be labelled 'Celtic Tiger', resulted in substantial growth and rise in trade-openness while the welfare provision over the same period deteriorated relative to the EU and OECD averages — see Kirby (2008) for details. On the other hand, Germany (which we had to exclude from our sample given the short span of data due to the Unification break) has maintained its strong tradition of social protection policies which are highly consistent with the compensation hypothesis. The country's exports (and hence trade-openness) have in recent months reached their all time high while it currently has one of the most comprehensive welfare systems in Europe with 26.7% percent of its GDP being channelled into public welfare spending — in comparison with 20.7% for all OECD average and 15.9% for the USA; see Giehle (2011) for further details.

³ For instance, Epifani and Gancia (2009b) use their results in Epifani and Gancia (2009a) to make strong policy recommendations. They state that "*In a world of integrated product markets, sustaining demand via public spending is considered more effective than a tax cut*" and speculate on the benefits of extending "...the WTO rules, currently limiting terms-of-trade manipulations arising from non-cooperative trade policy, to correct the externalities arising from fiscal policy as well." While we acknowledge that these might be sound general conclusions, we are not able to deduce which country would actually comply with and benefit from such policies.

openness to government-size. These results question the universality of the compensation hypothesis — or, indeed, that of its rival efficiency hypothesis — and beg for a more careful scrutiny of both the theoretical processes underlying the link between the extent of openness to trade in a country and the relative size of its public sector, as well as the appropriateness of measurements which approximate economic globalisation and government-size.⁴

The rest of the paper is organised as follows. Section 2 explains the data and illustrates the relevant patterns. Section 3 outlines the main aspects of our econometric methodology whose details are provided in an appendix. Section 4 reports the results of the causality tests. Section 5 concludes the paper.

2. Data and preliminary observations

We use data with annual frequency over the post war period from the following 22 OECD countries⁵: Australia (1), Austria (2), Belgium (3), Canada (4), Denmark (5), Finland (6), France (7), Greece (8), Iceland (9), Ireland (10), Italy (11), Japan (12), Luxembourg (13), Netherlands (14), New Zealand (15), Norway (16), Portugal (17), Spain (18), Sweden (19), Switzerland (20), United Kingdom (21) and United States (22), where the numbers in parentheses are the reference numbers for these countries which we use in Tables 1 and 2 below. We have used the same measures of *trade-openness* and *government-size* as those in Rodrik (1998) and Epifani and Gancia (2009), namely (exports+imports)/GDP and (government consumption)/GDP. In the rest of the paper we shall refer to these variables by Y and G , respectively

⁴ In recent studies, the extent of economic globalisation has been approximated by different measures. For instance, Liberati (2007) uses inward and outward FDI, portfolio investment as well as trade-openness. He constructs various (unbalanced) panels by pooling data from the main industrial countries (excluding Japan). Estimating various specifications, he finds only a negative significant impact from financial-openness on both government-size and the allocation of the public sector between central and local governments. Gemmell et al. (2008) use the inward stock of FDI as well as trade-openness. Using data from 25 OECD countries over the period 1980–1997, they estimate panel error correction specifications. They find these variables not to affect government-size, but report that FDI significantly shifts government expenditure towards social spending. Kimakova (2009) uses dynamic panel methods to estimate regression equations which generalise the analysis of Rodrik (1998) in capturing the impact of trade- and financial-openness on government-size. In particular, she controls for demographic, institutional and political factors and measures financial openness by private capital flows. She uses data from 87 developing and developed countries between 1976 and 2003 (but only utilises 6 data points during this period) and finds that while there is a positive relationship between international capital flows and government-size, richer open economies tend to have relatively smaller governments. Some studies — e.g., see Iversen and Cusack (2000) and Iversen (2001) — argue that a further channel, in addition to trade-openness, stems from the fact that the risk mitigating expansion of government spending could be due to responding to the needs for social protection that are felt as a country's production activities change, e.g., as new efficient sectors displace the more traditional ones hence requiring a change in the skill composition of the country's labour force.

⁵ Data are from International Finance Statistic and Government Finance Statistic (IMF publications). We have chosen to focus on all the industrialised countries for which data for longest common period exists (and hence Germany is excluded due to the unification break). For all countries, data exists for the 1955-2003 period, but for some we can go back to 1948. While data can be extended beyond 2003, this sample period has the advantage of being roughly compatible with that used in recent studies.

To form a basic view of how these countries compare and whether the time dimension is an important factor, in Table 1 we provide scatter-diagrams across these countries that plot G against Y using average data similar to that used in Rodrik's study — i.e. the average of G over a number of years is plotted against the average of Y over the previous decade — for the following four decades: 1960-1969, 1970-1979, 1980-1989 and 1990-1999. As it is clear from these diagrams, an individual country's position over time is not immutable and some countries have changed their position remarkably from one decade to the next. Also, as the broken lines (which represent the best fit based on a fractional polynomial specification) show, the nature of the relationship seems to change over the four decades. In order to eliminate any bias in the fit due to Luxemburg's position (country number 13) which may be considered as an outlier, Table 2 repeats the exercise excluding Luxemburg from the sample. The result of this exercise indicates that a mild positive relationship between G and Y seems to hold during the 1960s and 1970s, but is not present in the following two decades. This simple examination of the data clearly indicates that the nature of the relationship between trade-openness and government-size across the countries in the sample has altered over the four decades under consideration and thus it will matter which years one uses for the exercise.

Garrett (2001) explains this phenomenon by noting that, insofar as the relationship between trade-openness and government-size is an effect of globalisation, it ought to be considered as an *evolving process* rather than a *steady state position*. He then stresses the need to distinguish between the short-run and long-run nature of the relationship between these two variables. Measuring trade-openness and government-size as in Rodrik (1997a,b), Garrett compares the results of cross-country regressions based on levels (averaged over the 1985-1995 period) with those based on changes (measured as the difference between 1970-1984 averages and 1985-1995 averages). His results confirm the importance of this distinction: whilst the regressions based on levels support Rodrik's finding that more open countries tend to have larger governments, those based on changes indicate that government-size tends to grow slower in those countries whose trade-openness has expanded faster. This evidence casts doubt on the robustness of Rodrik's findings and points to the need for a more thorough and robust examination of the interaction between the short-run and long-run behaviour.

We conclude this section by a simple graphical inspection of the way the relationship between our series G and Y has evolved over time in each one of the countries within our sample. Table 3 plots government-size against trade-openness for each country separately. A glance at these plots confirms that the time-series nature of the relationship is not so straightforward. In fact, only 7 countries — Australia, Belgium, France, Italy, Portugal, Spain and Sweden — show

some support for a positive relationship between these variables; Norway reveals a very clear negative relationship and the rest of the countries do not seem to exhibit any clear pattern.⁶

3. Econometric methodology

Our main purpose is to examine the existence of short-run and long-run causal effects from trade-openness (Y) to government-size (G) for each of the above mentioned countries over the sample period 1955-2003. The usual approach would be to follow Granger's procedure: estimate an adequately specified bivariate vector autoregressive (VAR) model and use the standard F test (or its asymptotic equivalents) to check whether the past history of Y contributes significantly to explaining the current level of G — see Lütkepohl (1982) on block exogeneity tests within the VAR framework. There are two shortcomings with this approach. The first concerns the omission of the so called 'other relevant variables'. Although Granger's approach is thought to perform well as long as the residuals of the bivariate VAR model are well-behaved, any statistics based on the estimates could lack precision if the model omitted relevant variables on which Y and G should have been conditioned *a priori*. In the interest of robustness, therefore, one is required to add at least the most relevant variables that are thought to belong to the information set on which Y and G are conditioned. We propose to deal with this problem by including the real GDP and the inflation rate as our 'other relevant variables' in the VAR system.⁷

The second shortcoming stems from the specification of the VAR model. As Engle and Granger (1987) point out, the conventional VAR will be misspecified, hence the estimates could be inconsistent and inefficient, if the existence of a long-run, cointegrating, relationship between the variables could not be ruled out. To take account of this, we use the version of 'Granger non-causality' test that is applicable within the vector error correction (VECM) framework. This procedure, detailed in the Appendix, allows for explicitly distinguishing between the long-run

⁶ Rewriting the national accounts identity $y=c+i+g+x-m$ as $[(c+i)/y-2m/y]+g/y+(x+m)/y=1$, where small letters are used to avoid confusion, the accounting links between trade-openness $(x+m)/y$ and government-size g/y can be highlighted. For instance, for both of these to rise we need $[(c+i)/y-2m/y]$ to fall, which requires a significant reduction in the size of private sector relative to that of import sector. This can also be used to explain the situation in Norway where $(x+m)/y$ is rising whilst g/y falls (see Tables 3 and 4). As a referee has kindly pointed out, in Norway x is likely to have grown steadily and faster than the other variables due to oil exports. This is likely to have been reflected in both y and m growing such that $(x+m)/y$ rose, but g/y could fall simply because g did not need to grow as fast.

⁷ To maintain consistency with the existing literature, Y and G are respectively defined as the ratio of general government consumption expenditure and (exports+imports) to GDP, all measured at current prices. Also, in the interest of parsimony, we have limited the analysis to these four variables as we think that their history provides adequate information for our analysis within the VAR context. Real income is expected to be linked to both openness and government-size for a number of reasons; in particular, theoretical priors suggest that trade and income ought to be positively correlated and Wagner's law anticipates a positive causation from income to government expenditure, while an opposite causality is predicted via the typical Keynesian fiscal policy effects. Inflation is usually expected to dampen the effects of income and could also be argued to capture any reallocation effects.

and short-run links between the variables in question.⁸ Let us denote by $X'_t = (Y_t, G_t, Q_t, I_t)$ the vector of observations at time t on our variables where Q and I are our ‘other relevant variables’ and stand for the logarithm of real GDP and the inflation rate, respectively. We therefore base our causality tests on the estimates of the following VECM system of order p ,

$$\Delta X_t = \sum_{s=1}^{p-1} \Gamma_s \Delta X_{t-s} + \Pi X_{t-1} + U_t, \quad (1)$$

where the coefficients matrices can be conformably expanded as

$$\Gamma_s = \begin{bmatrix} \Gamma_{yy,s} & \Gamma_{yg,s} & \Gamma_{yq,s} & \Gamma_{yi,s} \\ \Gamma_{gy,s} & \Gamma_{gg,s} & \Gamma_{gq,s} & \Gamma_{gi,s} \\ \Gamma_{qy,s} & \Gamma_{qg,s} & \Gamma_{qq,s} & \Gamma_{qi,s} \\ \Gamma_{iy,s} & \Gamma_{ig,s} & \Gamma_{iq,s} & \Gamma_{ii,s} \end{bmatrix}, \quad \Pi = \begin{bmatrix} \Pi_y \\ \Pi_g \\ \Pi_q \\ \Pi_i \end{bmatrix}.$$

For each country, two possible scenarios exist:

- (a) We may find that the relevant unit root tests do not hold and/or that Johansen’s non-cointegration tests cannot reject the restriction $r \equiv \text{Rank}(\Pi) = 0$. In this case, there is no long-run causation and we simply apply the conventional block exogeneity test based on the null hypothesis $\Gamma_{gy,s} = 0$ for $s=1, \dots, p-1$. Rejecting these restrictions would then support the existence of (short-run) *Granger-type causality* from Y to G , in that the past values of Y significantly contribute to predicting the current value of G . In this situation, we approximate the impact of ΔY on ΔG by the estimate of

$$m_{gy} = \frac{\sum_{s=1}^{p-1} \Gamma_{gy,s}}{1 - \sum_{s=1}^{p-1} \Gamma_{gy,s}}, \quad (2)$$

and use the corresponding asymptotic standard error to test the statistical significance of the estimate, \hat{m}_{gy} .

- (b) Alternatively, we may find that the relevant unit root tests hold and that Johansen’s non-cointegration tests reject $r \equiv \text{Rank}(\Pi) = 0$. In this case, a long-run causation might be present and we need to check if it is statistically significant. Johansen (1991) shows that, when $r \equiv \text{Rank}(\Pi) \geq 1$, parameter matrices α and β exist such that $\Pi = \alpha \beta'$. Recalling that Y is the first element of vector X , the existence of a long-run causality then involves testing the hypothesis $\alpha_{g,1}^* = 0$ when $r=1$ and sequentially testing hypotheses $\beta_{11}^* = 0$ and $\alpha_{g,1}^* = 0$ when $r>1$ where $\alpha_{g,1}^*$ and β_{11}^* are the relevant elements of the normalised versions of α and β (see the Appendix for details). The short-run causality remains as

⁸ Recent theoretical studies of this approach can be found in Toda and Phillips (1993), Giannini and Mosconi (1992), Pesaran et al. (2000), Rault (2000) and Pradel and Rault (2003), amongst others.

before and involves testing the null hypothesis $\Gamma_{gy,s} = 0$ for $s=1, \dots, p-1$ and calculating the corresponding \hat{m}_{gy} .

4. Results

As a first step, we used standard statistical techniques to test for the existence of unit roots in the variables in order to determine their nonstationarity properties. We found that, with the exception of Iceland, the hypothesis that the series are first difference stationary could not be rejected in any of the countries in the sample.⁹ This implies that, in general, in each of the remaining 21 countries we may find up to three cointegrating relationships between the four variables in question (Iceland was excluded from the cointegration analysis and the causality test was carried out using a VAR system based on the stationary transformations of the variables). We then used the likelihood ratio test, information criteria, and the residual autocorrelation tests to establish p , the order of the VAR, for each country and went on to apply Johansen's 'Trace' and 'Maximum Eigenvalue' tests to determine $r \equiv \text{Rank}(\Pi)$. Finally, depending on the value of r , we proceeded to test for causality as described above.

The results are reported in Table 4. For each country, a graph illustrates the behaviour of G and Y over the sample period, shown by the solid and broken lines measured on the left and right axes, respectively. Below the graph of each country, we give the values of p and r and, depending on the value of r , go on to provide the values of the following test statistics where the associated probabilities (p -values) are given in square brackets:

$S_{(\beta)}$: likelihood ratio statistic for $\beta_{11}^* = 0$ against $\beta_{11}^* = 1$, distributed asymptotically as $\chi^2_{(r)}$

$S_{(\alpha)}$: likelihood ratio statistics for $\alpha_{g,1}^* = 0$ against $\alpha_{g,1}^* \neq 0$, distributed asymptotically as $\chi^2_{(1)}$

$S_{(\Gamma)}$: Wald statistics for $\Gamma_{gy,s} = 0$, for all $s = 1, \dots, p-1$, distributed asymptotically as $\chi^2_{(p-1)}$

We also provide estimates of $\alpha_{g,1}^*$ and m_{gy} , indicated by a '^', where the number in parentheses following each estimate is the corresponding asymptotic t -ratio.

On the whole, the results are not very supportive of the compensation hypothesis. Only 5 out of 22 countries — Austria, Australia, Canada, France and Greece — show a significant (and positive) causation from past Y to current G . Moreover, only for Australia are both the short-run and long-run channels significantly present; in Canada, France and Greece the causation is due to the long-run channel while in Austria the nature of causation is merely short-run. While

⁹ We used both Augmented Dickey-Fuller and Ng and Perron — ADF and NP — statistics to examine the existence of a unit root in the series. See MacKinnon (1996) for the critical values of the ADF statistic and Ng and Perron (2001) for the NP statistic. As mentioned above, the exception was the trade-openness series in Iceland which we found to be stationary at 5% significance level. We also found the inflation series to be stationary for Australia and therefore used the logarithm of price level instead of inflation for that country. The unit root and the subsequent test results are not reported in the paper but are available on request from the authors.

significant short-run effects are also present in Belgium and Portugal, and to some extent in Finland, these seem to cancel out in the aggregate rendering the estimated total effect captured by m_{gy} statistically insignificant. For the rest of the countries, there is no evidence of a significant causation from Y to G .

Given the relative robustness of these results — in that they are obtained (i) without imposing any restrictions on data across different countries, and (ii) using a more vigorous time series approach — our findings cast some doubt on the theoretical presumption that the public sectors of industrialised countries tend to grow as the extent of openness to trade in these economies is enhanced. In fact, these results strongly indicate that the relationship between government-size and trade-openness is, to a large extent, country specific and hint that caution should be exercised when pooled data from a number of countries are used to study the likely impact of globalisation on the size and composition of the public sector without at least establishing a priori that the pooled sample satisfies the most essential pooling restrictions.

5. Summary and conclusion

Conventional wisdom holds that the authorities in more open economies are under pressure to reduce taxes and to trim down public expenditure so as to maintain competitiveness in goods markets and in attracting foreign capital and/or preventing the flight of increasingly footloose firms. The observed positive relationship between globalisation and government-size challenges this view. The *compensation hypothesis*, formalised by Rodrik (1997a), casts doubts on the predicted negative effect of globalisation on the size of governments and contends that the public sector of a more open economy is likely to be larger in order to meet higher demands by economic agents for public insurance, induced by bigger exposure to external risks. Recently, Epifani and Gancia (2009a) have strengthened the reason underlying the existence of a positive causality by adding a terms of trade effect. They propose that the reason why more open countries have a relatively larger government is because a rise in trade reduces the relative price of public goods which in turn facilitates its growth. Both Rodrik and Epifani and Gancia obtain results that supports a positive causality running from a country's international openness to its public sector size.

The evidence provided by the above studies is based on data that pool many countries together and both studies find that the positive correlation between openness and government-size holds strongly for specific subset of countries. The main purpose of this paper has been to see whether there is any statistical causation at the *individual country* level that is consistent with the compensation hypothesis. We examine the same measure of openness and government-size used by Rodrik and Epifani and Gancia but focus on 22 OECD countries *individually* over the period 1955-2003. We carry out the time series causality tests in the context of general VAR

analysis to examine the existence of long-run and short-run causality. Our empirical investigation fails to find a unanimous support for such causality; instead there is strong evidence that the short-run and long-run links between trade-openness and government-size are country specific and are on the whole rather weak. This evidence can be interpreted as casting doubt on the validity of the compensation hypothesis insofar as it concerns the relationship between specific measures of economic globalisation and the size of the public sector used in these studies — namely, $(\text{exports}+\text{imports})/\text{GDP}$ and $(\text{government consumption})/\text{GDP}$. Of course, our findings could also be taken as an indication that increasing openness does not necessarily invoke a net expansion of governments. In fact, as Rodrik (1997b) suggests, increasing openness in capital and financial markets, by constraining the revenue raising ability of governments, may severely undermine the potential for public expansion.¹⁰ It is also possible that government-size in general does not represent the most relevant indicator that responds to openness. For instance, it could be argued — as Rodrik and others do — that a more suitable measure for mature industrial economies in this context is in fact the welfare spending component of government budget. Moreover, as Rodrik (1998) points out, one may complement the direct causality tests described above by an alternative, indirect, approach that allows examining if openness does, in the first instance, raise risk (captured by, for instance, an increase in consumption volatility and/or unevenness of income distribution), and whether this rise in risk is significantly hampered by an increase in the relevant components of government spending.

One way to proceed, therefore, is to expand the set of variables to include those mentioned above, namely, capital mobility, balance of payments, inward and outward FDI, welfare components of government spending, and income or consumption volatility indices, and break the causality chain into two stages as described above. However, time series data on these additional variables do not exist for a sufficiently long period to enable analysis for individual countries to be carried out within the VAR-VECM context. Thus, further applied research in this connection would require the use pooled time series data. Given that our empirical analysis in this paper shows clearly that the dynamics of the relationship between openness and government-size are country specific, obtaining robust evidence from pooled data involving different countries requires appropriate econometric techniques which allow for this type of heterogeneity in the sample.¹¹

¹⁰ However, this still remains a controversial issue since some studies have convincingly argued that capital mobility is likely to be associated with even more public spending — see, for instance, Quinn (1997) and Garret (1995), or Bretschger and Hettich (2002) and Devereux et al. (2008) for an opposite view.

¹¹ See Pesaran and Zhao (1999), Hsiao et al. (2002), Binder et al. (2005) and Coakley et al. (2004) for a discussion of the theoretical issues. Smith and Fuertes (2003) review the relevant econometric issues and Pesaran et al. (2000) provide an application.

APPENDIX: Econometric methodology

The testing procedure used in this paper is based on the generalised version of the original ‘Granger non-causality’ test, which allows for the existence of a cointegration relationship between the variables involved. To explain the general principle briefly, let X_t be the vector of observations at time t on n variables (in our analysis above $n=4$). Consider the situation in which all variables in vector X_t are $I(1)$ and the VAR(p) representation,

$$X_t = \sum_{s=1}^p A_s X_{t-s} + U_t, \quad (A1)$$

exists¹² where A_s satisfy the required conditions and $U_t \sim N(0, \Omega)$. Johansen (1991) shows that the VAR system in (A1) can be reparameterised as the following VECM system,

$$\Delta X_t = \sum_{s=1}^{p-1} \Gamma_s \Delta X_{t-s} + \Pi X_{t-1} + U_t, \quad (A2)$$

where the coefficients in Π inform us of the existence of (maximum of $n-1$) long-run relationships between the n variables and the way deviations from them feed back into the evolution of the variables. It has also been shown that when $\text{Rank}(\Pi) \geq 1$ — i.e., when one or more long-run relationships exist — omitting X_{t-1} from (A2) can result in serious misspecification problems.¹³

Now, let us partition the n variables in X into two groups consisting of n_1 and n_2 variables denoted by vectors Y to Z respectively where $n_1+n_2=n$, and suppose that we wish to examine the significance of causation from Y to Z . Hence, let $X'_t = (Y'_t, Z'_t)$ and partition U_t , Ω , Γ_s and Π conformably as,

$$U'_t = (U'_{y,t}, U'_{z,t}), \quad \Omega = \begin{bmatrix} \Omega_{yy} & \Omega_{yz} \\ \Omega_{zy} & \Omega_{zz} \end{bmatrix}, \quad \Gamma_s = \begin{bmatrix} \Gamma_{yy,s} & \Gamma_{yz,s} \\ \Gamma_{zy,s} & \Gamma_{zz,s} \end{bmatrix}, \quad \Pi = \begin{bmatrix} \Pi_y \\ \Pi_z \end{bmatrix},$$

and use these to rewrite (A2) as follows

$$\Delta Y_t = \sum_{s=1}^{p-1} (\Gamma_{yy,s} \Delta Y_{t-s} + \Gamma_{yz,s} \Delta Z_{t-s}) + \Pi_y X_{t-1} + U_{y,t}, \quad (A2.1)$$

and

$$\Delta Z_t = \sum_{s=1}^{p-1} (\Gamma_{zy,s} \Delta Y_{t-s} + \Gamma_{zz,s} \Delta Z_{t-s}) + \Pi_z X_{t-1} + U_{z,t}. \quad (A2.2)$$

The sub-system in (A2.1) can then be rewritten such that it defines the ‘conditional model’ for Y_t , namely,

¹² For simplicity, we have excluded the intercept and time trend from our analytical discussion here, but have taken them into account in our empirical analysis.

¹³ See Engle and Granger (1987) for the original discussion of the properties of VECMs, and Pesaran et al. (2000) for more recent developments.

$$\Delta Y_t = \sum_{s=1}^{p-1} \Gamma_{yy,s}^* \Delta Y_{t-s} + \sum_{s=0}^{p-1} \Gamma_{yz,s}^* \Delta Z_{t-s} + \Pi_y^* X_{t-1} + U_{y,t}^*, \quad (\text{A2.1}')$$

so that (A2.2) constitutes the corresponding ‘marginal model’ for Z_t . (A2.1') is obtained by pre-multiplying (A2.2) by $\Theta = \Omega_{yz} \Omega_{zz}^{-1}$ and subtracting the resultant from (A2.1) and rearranging such that

$$U_{y,t}^* = U_{y,t} - \Theta U_{z,t}; \Pi_y^* = \Pi_y - \Theta \Pi_z; \Gamma_{yz,0}^* = \Theta; \Gamma_{yy,s}^* = \Gamma_{yy,s} - \Theta \Gamma_{zy,s}; \Gamma_{yz,s}^* = \Gamma_{yz,s} - \Theta \Gamma_{zz,s}.$$

Thus, the covariance matrix of the new full system, consisting of (A2.1') and (A2.2), satisfies the required independence conditions, that is

$$\text{Cov} \left(\begin{bmatrix} U_{y,t}^* \\ U_{z,t}^* \end{bmatrix} \right) = \begin{bmatrix} \Omega_{yy}^* & 0 \\ 0 & \Omega_{zz} \end{bmatrix}, \quad (\text{A3})$$

where $\Omega_{yy}^* = \Omega_{yy} - \Theta \Omega_{zy}$ — for further details see Pesaran et al. (2000).

The above framework enables us to construct our causality test more rigorously. Let Y be the ‘trade openness’ variable — thus $n_1=1$ and $n_2=3$ — and partition Z such that its first variable is ‘government size’ G and the rest, denoted by W , contain the other $n-2=2$ ‘other relevant variables’. Thus, $Z_t' = (G_t, W_t')$ and the marginal model in (A2.2) is conformably partitioned as

$$\Delta G_t = \sum_{s=1}^{p-1} (\Gamma_{gy,s} \Delta Y_{t-s} + \Gamma_{gg,s} \Delta G_{t-s} + \Gamma_{gw,s} \Delta W_{t-s}) + \Pi_g X_{t-1} + U_{g,t}, \quad (\text{A2.2.1})$$

and

$$\Delta W_t = \sum_{s=1}^{p-1} (\Gamma_{wy,s} \Delta Y_{t-s} + \Gamma_{wg,s} \Delta G_{t-s} + \Gamma_{ww,s} \Delta W_{t-s}) + \Pi_w X_{t-1} + U_{w,t}, \quad (\text{A2.2.2})$$

and the process of testing whether Y ‘does not Granger cause’ G — in accordance with Granger (1969) — can focus on (A2.2.1) which shows explicitly the channels through which past Y can have an impact on G ; the effect is captured by $\Gamma_{gy,s}$ as well as the first element of Π_g which is associated with Y_{t-1} . While it is straightforward to test restrictions concerning $\Gamma_{gy,s}$ using the standard nested methods — e.g., Wald, LR or LM statistics — within the *seemingly unrelated regressions* framework, testing restrictions regarding the elements of Π_g turns out to be rather complicated. This is because the elements of Π_g are non-linear combinations of the so-called long-run and adjustment coefficients, and the asymptotic properties of the above test statistics are not clearly defined under joint restrictions involving these elements especially when $\text{Rank}(\Pi) > 1$ — see Toda and Phillips (1993). However, a number of studies — e.g., Hunter (1992), Giannini and Mosconi (1992), Rault (2000), and Pradel and Rault (2003) — have suggested indirect ways of testing for exogeneity in these circumstances and the approach we propose below is similar in spirit to that proposed by Rault (2000). An added advantage in our case is the fact that our tests involve only one variable and Y_t is a scalar and hence, as we shall

see below, the rank of the relevant long-run sub-matrix cannot exceed unity even when $\text{Rank}(\Pi) > 1$.¹⁴

Johansen (1991) has shown that matrices β and α exist such that $\Pi = \alpha \beta'$. β is the $(n \times r)$ matrix of long-run coefficients whose columns constitute the coefficients of cointegrating vectors. For $j=1, \dots, r$, let β_j denote the j th column of β . Then each $\beta_j' X_t$ is a stationary residual which reflects the deviation at time t from a long-run equilibrium relationship between elements of X_t . α is the $(n \times r)$ matrix of adjustment coefficients; each α_{ij} element, $i=1, \dots, n$ and $j=1, \dots, r$, captures the impact of $\beta_j' X_{t-1}$ on the evolution of the i th element of X_t . Thus, denoting by $\alpha_g = (\alpha_{g,1}, \dots, \alpha_{g,r})$ the row of α corresponding to equation (A2.2.1) above which explains G , we have $\Pi_g X_{t-1} = \sum_{j=1}^r \alpha_{g,j} \beta_j' X_{t-1}$. Recalling that Y is the first element of vector X and denoting the first element of β_j by β_{1j} , it becomes clear that each $\alpha_{g,j} \beta_{1j}$ embodies a part of the impact that Y_{t-1} exerts on ΔG_t , whilst each $\Gamma_{gy,s}$ captures the effect of ΔY_{t-s} on ΔG_t . The causality test therefore ought to take account of both these channels. However, the test can be conducted in two stages, first checking the significance of the long-run channel through testing $\alpha_{g,j} \beta_{1j} = 0$, $j=1, \dots, r$ and then testing the hypothesis $\Gamma_{gy,s} = 0$, for all $s=1, \dots, p-1$ in order to check the short-run effects.

As mentioned above, a standard test based on the conventional statistics for the significance of the long-run causality channel has not yet been developed. However, the underlying hypothesis to be tested is shown to be simplified by exploiting the fact that the elements of α and β are not uniquely determined. This is because for any $(r \times r)$ non-singular matrix D , α and β can be replaced with $\alpha^* = \alpha D^{-1}$ and $\beta^* = D\beta$, respectively, since $\Pi = \alpha \beta' = \alpha^* \beta^{*'}$ and this pivoting does not affect the estimated value of the likelihood function of the underlying VECM. This property is used to uniquely identify β by choosing D so as to impose $(r \times r)$ identifying (normalisation) restrictions on β . For our purpose, we shall let D be such that the top $(r \times r)$ sub-matrix of β^* is the unit matrix. More explicitly, partition β as $\beta' = [\beta'_{(r \times r)} \quad \beta'_{(r \times n-r)}]$ where $\beta'_{(r \times r)}$ and $\beta'_{(r \times n-r)}$ are the corresponding $(r \times r)$ and $(r \times n-r)$ sub-matrices. Then, given that β is full column rank, the inverse of $\beta'_{(r \times r)}$ also exists and we can choose $D = (\beta'_{(r \times r)})^{-1}$. Hence, $\beta^{*'} = [\beta^{*'}_{(r \times r)} \quad \beta^{*'}_{(r \times n-r)}]$, where $\beta^{*'}_{(r \times r)} = I$ and $\beta^{*'}_{(r \times n-r)} = (\beta'_{(r \times r)})^{-1} \beta'_{(r \times n-r)}$. It

¹⁴ When we cannot reject the hypothesis $\text{Rank}(\Pi) = 0$, $\Pi_g = 0$ can be safely imposed on (A2.2.1) and the causality test reduces to that originally proposed by Granger (1969).

also follows that $\alpha^* = \alpha \beta'_{(r \times r)}$. The test concerning the hypotheses $\alpha_{g,j} \beta_{1j} = 0$, $j=1, \dots, r$, then reduces to testing the single restriction

$$\alpha_{g,1}^* \beta_{11}^* = 0. \quad (\text{A4})$$

To see this, write $\beta^* = [\beta_1^*, \beta_2^*, \dots, \beta_r^*]$, denote the first element of each β_j^* vector by β_{1j}^* , and recall that now by normalisation $\beta_{11}^* = 1$ and $\beta_{1j}^* = 0$ for $j > 1$. Clearly, the fact that $\beta_{11}^* = 1$ may lead one to further simplify the nonlinear restriction in (A4) and only consider testing the single homogenous restriction, $\alpha_{g,1}^* = 0$. But, in principle, we ought to allow for the possibility that Y does not feature significantly in any of the cointegrating vectors and hence first test $\beta_{11}^* = 0$ against $\beta_{11}^* = 1$ when $r \leq n-2$.¹⁵ If this restriction cannot be rejected, the ECM components, $\beta_j^{*'} X_{t-1}$, will not be a source of concern in connection with the causality problem. We then proceed by disregarding the long run causality (as if $\Pi_g = 0$, as explained above) and examine m_{gy} as explained above —see equation (2) in the main text. However, if $r = n-1$ or if $\beta_{11}^* = 0$ is rejected, we need to test whether $\beta_1^{*'} X_{t-1}$ features significantly in (A.2.2.1). This is done by testing $\alpha_{g,1}^* = 0$, whose non-rejection will again indicate that the causality test can safely proceed as if $\Pi_g = 0$ and check $m_{g,y}$. Otherwise, rejecting $\alpha_{g,1}^* = 0$ clearly implies that the long-run causality from Y to G cannot be ruled out. In this case, both $\alpha_{g,1}^*$ and $m_{g,y}$ reflect the nature and extent of causality. The restrictions $\beta_{11}^* = 0$ and $\alpha_{g,1}^* = 0$ can be tested sequentially (against $\beta_{11}^* = 1$ and $\alpha_{g,1}^* \neq 0$, respectively) using the likelihood ratio test — see Rault (2000) and Pesaran et al. (2000) for details.

¹⁵ Note that when $r=n-1$, the normalisation restrictions suggested above will imply that each column of β^* only has two nonzero elements, in which case $\beta_{11}^* = 0$ will be inappropriate when all variables are $I(1)$.

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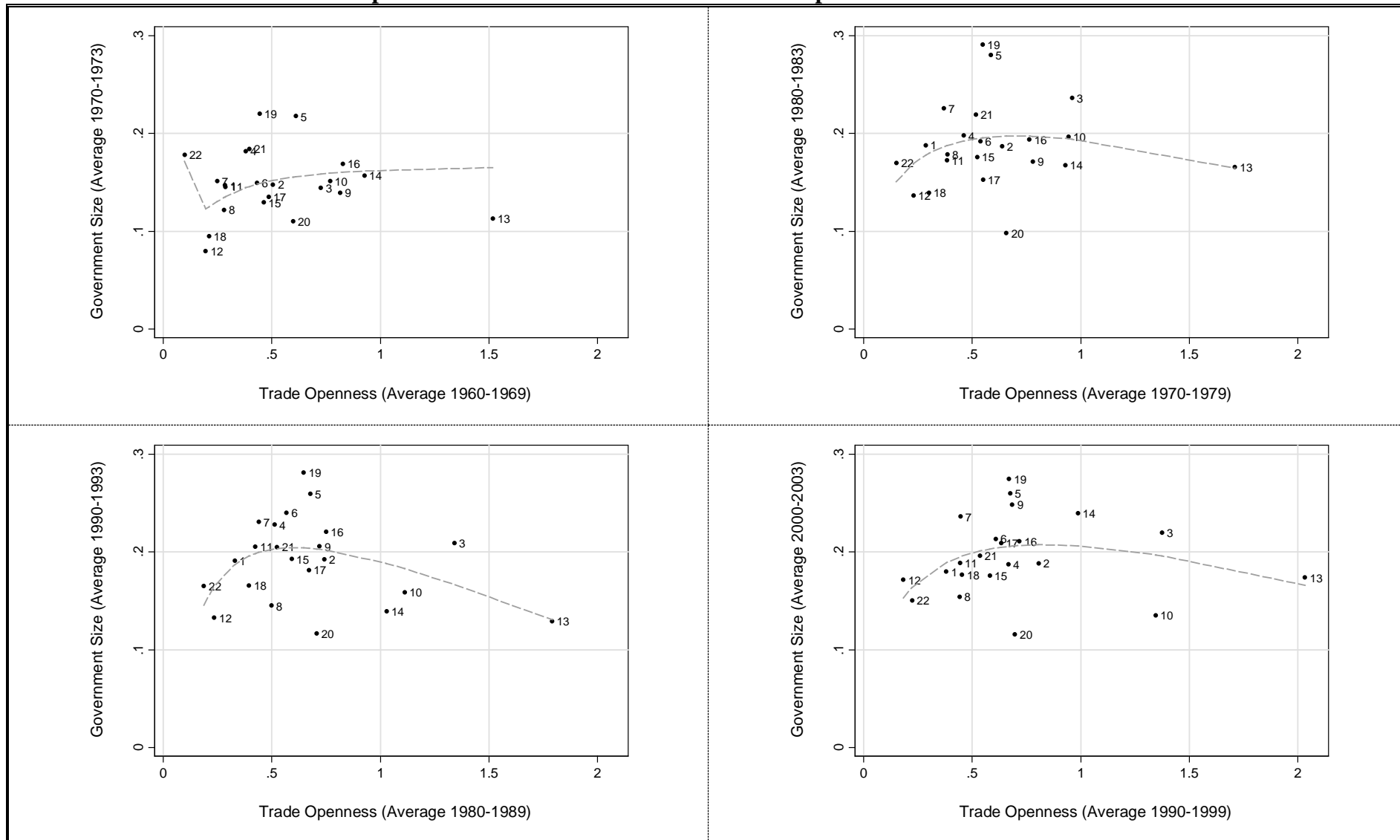
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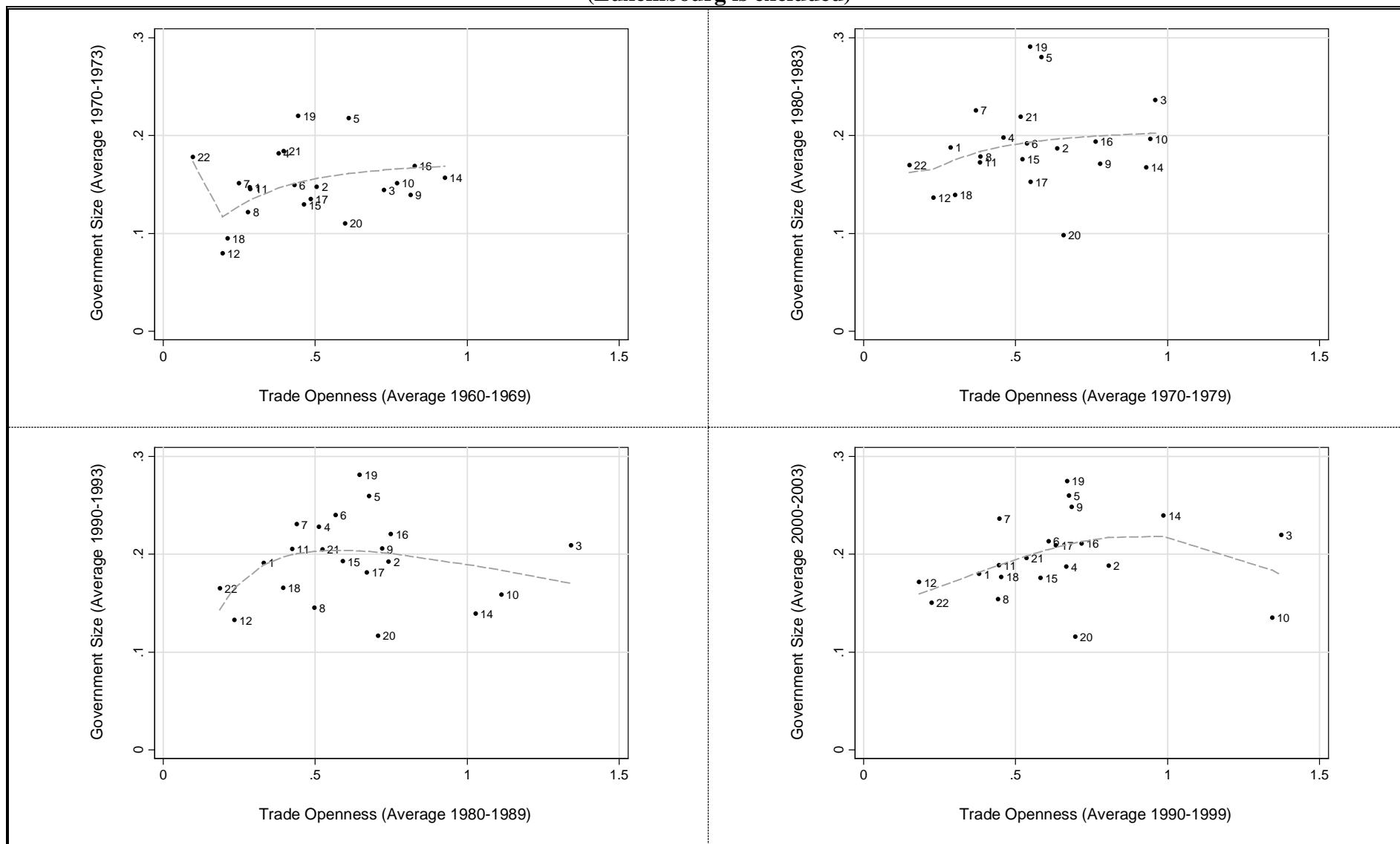
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Table 1. Cross section relationship between ‘Government Size’ and ‘Trade Openness’ across 22 OECD countries over the last 4 decades



- The dashed lines represent the best fit using the fractional polynomial procedure.
- The numbers refer to the countries as listed at the beginning of Section 2 above.

**Table 2. Cross section relationship between ‘Government Size’ and ‘Trade Openness’ across 21 OECD countries over the last 4 decades
(Luxembourg is excluded)**



■ See note to Table 1.

Table 3. Relationship between ‘Government Size’ and ‘Trade Openness’ for 22 OECD countries over the last 5 decades

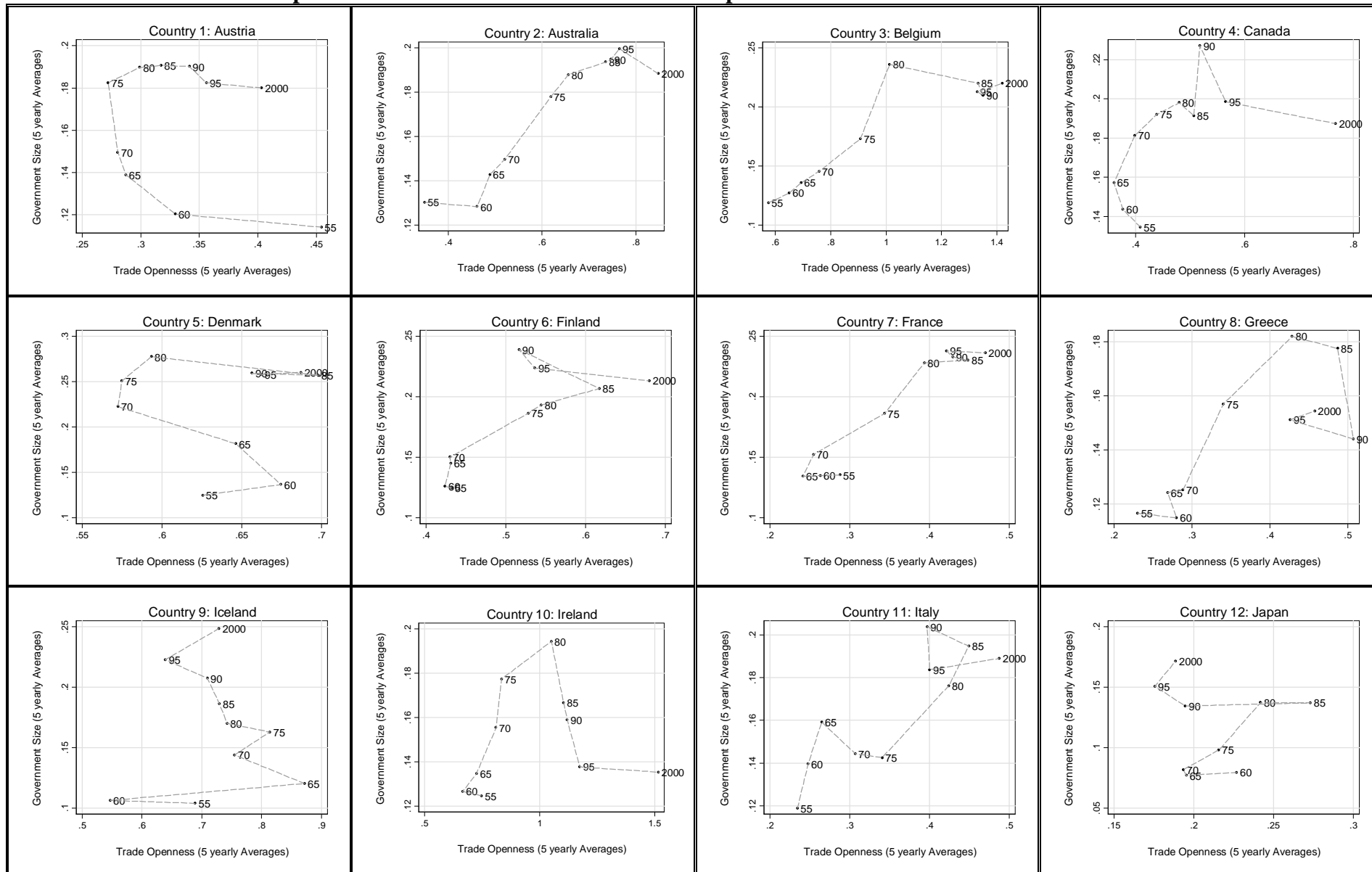


Table 3 continued

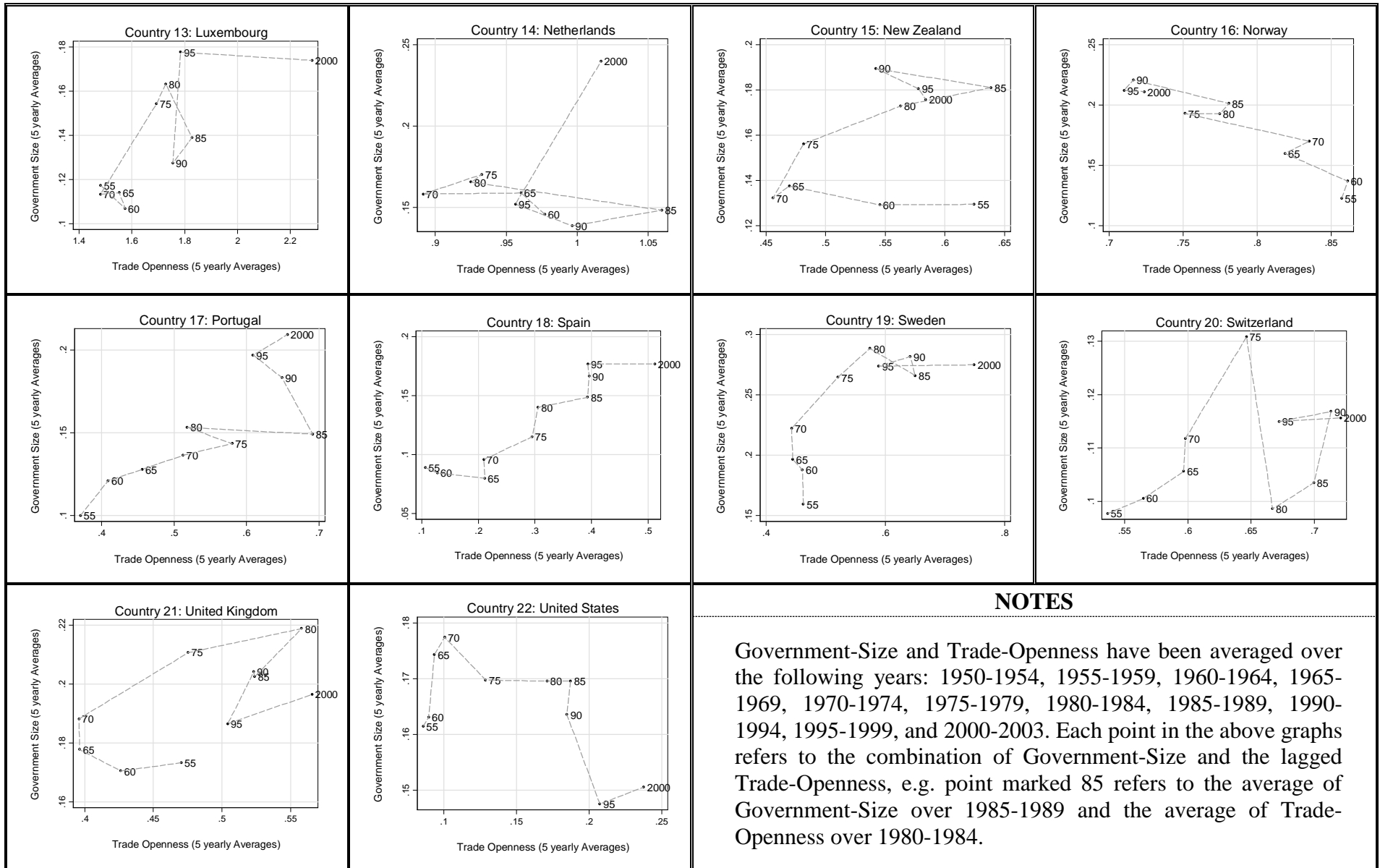


Table 4. Statistical tests for causality from ‘Government Size’ to ‘Trade Openness’ for 22 OECD countries^{1,2}

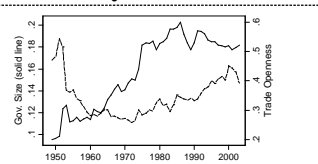
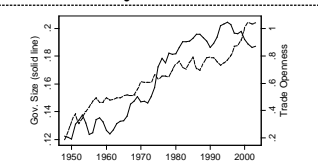
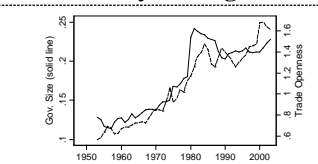
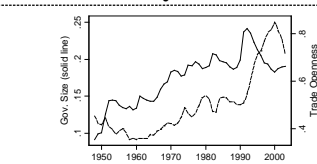
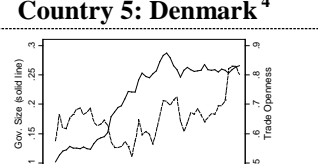

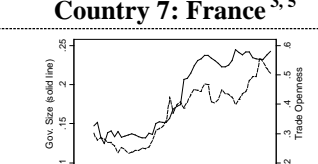
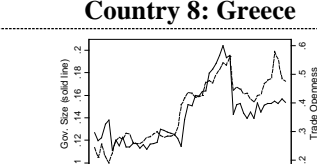

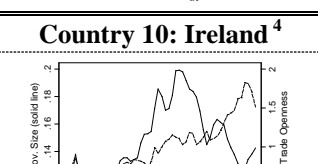

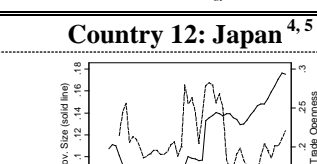
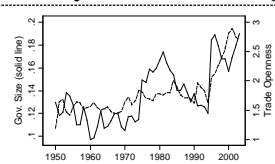
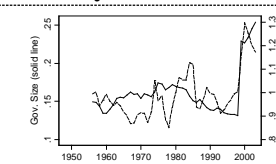
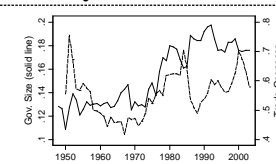
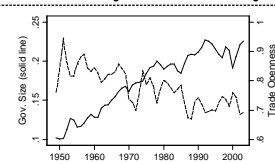
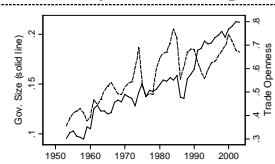
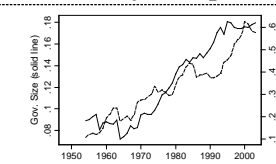
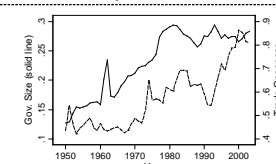
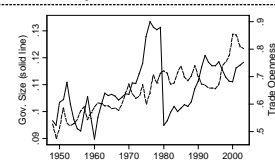
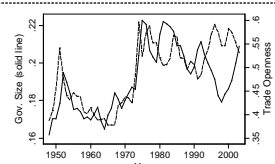
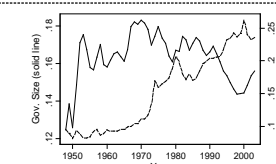
<p>Country 1: Austria³</p>  <p>Sample: 1949 – 2003 $p = 2; r = 2; S_{(\beta)} = 1.64 [0.44]$ $S_{(\alpha)} = NA; \hat{\alpha}_{g,1}^* = NA$ $S_{(\Gamma)} = 16.2 [0.0001]; \hat{m}_{gy} = 0.43 (3.32)$</p>	<p>Country 2: Australia</p>  <p>Sample: 1948 – 2003 $p = 3; r = 2; S_{(\beta)} = 21.96 [0.000017]$ $S_{(\alpha)} = 5.93 [0.015]; \hat{\alpha}_{g,1}^* = -0.133 (5.24)$ $S_{(\Gamma)} = 14.2 [0.0008]; \hat{m}_{gy} = 0.464 (2.3)$</p>	<p>Country 3: Belgium³</p>  <p>Sample: 1953 – 2003 $p = 2; r = 1; S_{(\beta)} = 3.51 [0.0612]$ $S_{(\alpha)} = 3.14 [0.076]; \hat{\alpha}_{g,1}^* = 0.09 (2.48)$ $S_{(\Gamma)} = 3.91 [0.048]; \hat{m}_{gy} = 0.367 (1.72)$</p>	<p>Country 4: Canada</p>  <p>Sample: 1948 – 2003 $p = 2; r = 1; S_{(\beta)} = 10.9 [0.00096]$ $S_{(\alpha)} = 11.32 [0.00077]; \hat{\alpha}_{g,1}^* = 0.134 (3.76)$ $S_{(\Gamma)} = 3.0 [0.083]; \hat{m}_{gy} = 0.53 (1.55)$</p>
<p>Country 5: Denmark⁴</p>  <p>Sample: 1950 – 2003 $p = 1; r = 2; S_{(\beta)} = 6.57 [0.037]$ $S_{(\alpha)} = 0.246 [0.62]; \hat{\alpha}_{g,1}^* = 0$ $S_{(\Gamma)} = NA; \hat{m}_{gy} = NA$</p>	<p>Country 6: Finland⁴</p>  <p>Sample: 1950 – 2003 $p = 2; r = 1; S_{(\beta)} = 10.5 [0.0012]$ $S_{(\alpha)} = 0.0432 [0.84]; \hat{\alpha}_{g,1}^* = 0$ $S_{(\Gamma)} = 3.02 [0.082]; \hat{m}_{gy} = 0.403 (1.13)$</p>	<p>Country 7: France^{3,5}</p>  <p>Sample: 1952 – 2003 $p = 1; r = 3; S_{(\beta)} = NA$ $S_{(\alpha)} = 3.77 [0.052]; \hat{\alpha}_{g,1}^* = 0.124 (4.39)$ $S_{(\Gamma)} = NA; \hat{m}_{gy} = NA$</p>	<p>Country 8: Greece</p>  <p>Sample: 1948 – 2003 $p = 2; r = 1; S_{(\beta)} = 18.5 [0.000017]$ $S_{(\alpha)} = 4.83 [0.028]; \hat{\alpha}_{g,1}^* = 0.088 (2.34)$ $S_{(\Gamma)} = 0.214 [0.64]; \hat{m}_{gy} = 0.051 (0.109)$</p>
<p>Country 9: Iceland⁶</p>  <p>Sample: 1950 – 2003 $p = 1; r = NA; S_{(\beta)} = NA$ $S_{(\alpha)} = NA; \hat{\alpha}_{g,1}^* = NA$ $S_{(\Gamma)} = 0.52 [0.47]; \hat{m}_{gy} = -0.043 (0.74)$</p>	<p>Country 10: Ireland⁴</p>  <p>Sample: 1948 – 2003 $p = 2; r = 1; S_{(\beta)} = 17.96 [0.000023]$ $S_{(\alpha)} = 0.0000 [0.999]; \hat{\alpha}_{g,1}^* = 0$ $S_{(\Gamma)} = 0.26 [0.61]; \hat{m}_{gy} = 0.07 (0.51)$</p>	<p>Country 11: Italy⁴</p>  <p>Sample: 1951 – 2003 $p = 4; r = 2; S_{(\beta)} = 27.33 [0.000001]$ $S_{(\alpha)} = 0.085 [0.77]; \hat{\alpha}_{g,1}^* = 0$ $S_{(\Gamma)} = 5.50 [0.138]; \hat{m}_{gy} = -0.51 (0.48)$</p>	<p>Country 12: Japan^{4,5}</p>  <p>Sample: 1955 – 2003 $p = 1; r = 1; S_{(\beta)} = 5.32 [0.021]$ $S_{(\alpha)} = 1.8 [0.18]; \hat{\alpha}_{g,1}^* = 0$ $S_{(\Gamma)} = NA; \hat{m}_{gy} = NA$</p>

Table 4 continued

<p>Country 13: Luxembourg⁷</p>  <p>Sample: 1950 – 2003 $p = 1$; $r = 0$; $S_{(\beta)} = NA$ $S_{(\alpha)} = NA$; $\hat{\alpha}_{g,1}^* = NA$ $S_{(\Gamma)} = 0.91 [0.34]$; $\hat{m}_{gy} = -0.23 (0.87)$</p>	<p>Country 14: Netherlands⁷</p>  <p>Sample: 1951 – 2003 $p = 2$; $r = 0$; $S_{(\beta)} = NA$ $S_{(\alpha)} = NA$; $\hat{\alpha}_{g,1}^* = NA$ $S_{(\Gamma)} = 0.53 [0.77]$; $\hat{m}_{gy} = 0.23 (0.39)$</p>	<p>Country 15: New Zealand³</p>  <p>Sample: 1950 – 2003 $p = 3$; $r = 1$; $S_{(\beta)} = 0.40 [0.53]$ $S_{(\alpha)} = NA$; $\hat{\alpha}_{g,1}^* = NA$ $S_{(\Gamma)} = 1.11 [0.58]$; $\hat{m}_{gy} = 0.017 (0.13)$</p>	<p>Country 16: Norway</p>  <p>Sample: 1949 – 2003 $p = 4$; $r = 2$; $S_{(\beta)} = 25.8 [0.000003]$ $S_{(\alpha)} = 3.36 [0.067]$; $\hat{\alpha}_{g,1}^* = -0.31 (1.61)$ $S_{(\Gamma)} = 6.22 [0.10]$; $\hat{m}_{gy} = 1.06 (1.44)$</p>
<p>Country 17: Portugal⁷</p>  <p>Sample: 1953 – 2003 $p = 2$; $r = 0$; $S_{(\beta)} = NA$ $S_{(\alpha)} = NA$; $\hat{\alpha}_{g,1}^* = NA$ $S_{(\Gamma)} = 9.42 (0.009)$; $\hat{m}_{gy} = -0.095 (0.64)$</p>	<p>Country 18: Spain³</p>  <p>Sample: 1954 – 2003 $p = 4$; $r = 2$; $S_{(\beta)} = 0.57 [0.75]$ $S_{(\alpha)} = NA$; $\hat{\alpha}_{g,1}^* = NA$ $S_{(\Gamma)} = 2.94 [0.40]$; $\hat{m}_{gy} = -0.02 (0.24)$</p>	<p>Country 19: Sweden⁴</p>  <p>Sample: 1950 – 2003 $p = 1$; $r = 1$; $S_{(\beta)} = 3.31 [0.069]$ $S_{(\alpha)} = 0.432 [0.51]$; $\hat{\alpha}_{g,1}^* = 0$ $S_{(\Gamma)} = NA$; $\hat{m}_{gy} = NA$</p>	<p>Country 20: Switzerland⁴</p>  <p>Sample: 1948 – 2003 $p = 3$; $r = 2$; $S_{(\beta)} = 7.96 [0.019]$ $S_{(\alpha)} = 1.87 [0.17]$; $\hat{\alpha}_{g,1}^* = 0$ $S_{(\Gamma)} = 3.84 [0.17]$; $\hat{m}_{gy} = 0.52 (1.76)$</p>
<p>Country 21: United Kingdom⁷</p>  <p>Sample: 1948 – 2003 $p = 2$; $r = 0$; $S_{(\beta)} = NA$ $S_{(\alpha)} = NA$; $\hat{\alpha}_{g,1}^* = NA$ $S_{(\Gamma)} = 1.93 [0.38]$; $\hat{m}_{gy} = 0.082 (0.53)$</p>	<p>Country 22: United States⁴</p>  <p>Sample: 1948 – 2003 $p = 4$; $r = 1$; $S_{(\beta)} = 6.13 [0.013]$ $S_{(\alpha)} = 0.41 [0.52]$; $\hat{\alpha}_{g,1}^* = 0$ $S_{(\Gamma)} = 1.48 [0.69]$; $\hat{m}_{gy} = -0.11 (0.84)$</p>	<p>Notes:</p> <ol style="list-style-type: none"> 1. The solid and broken lines respectively depict government-size and trade openness series. 2. The numbers in square brackets and in parentheses are the asymptotic p-values and t-ratios, respectively. 3. When $r > 0$ but the value of $S_{(\beta)}$ statistic is below the critical level, $S_{(\alpha)}$ statistic and $\hat{\alpha}_{g,1}^*$ coefficient are not relevant. This is denoted by NA. Also, when $r = 3$, $S_{(\beta)}$ is not relevant since the underlying restriction cannot be imposed. 4. When the value of $S_{(\alpha)}$ statistic is below the critical level, we impose $\hat{\alpha}_{g,1}^* = 0$. 5. Since the order of the VECM is $p-1$, when $p=1$ and $r > 0$ only the ECM terms appear as regressors and hence $S_{(\Gamma)}$ statistic and \hat{m}_{gy} coefficient are not relevant. This is denoted by NA. 6. In the case of Iceland, the trade openness series is found to be stationary and cointegration is not applicable. The analysis in this case is based on the VAR using stationary transformations of $I(1)$ series. 7. When $r=0$, $S_{(\beta)}$ and $S_{(\alpha)}$ statistics and $\hat{\alpha}_{g,1}^*$ coefficient are not relevant. This is denoted by NA. In these cases we have based the test on unrestricted VAR(p) in first differences. 	