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Is there a tax instrument that can be used to reduce public debt in Poland without causing output contraction? A Structural VAR analysis of the dynamic effects of various tax instruments on output.

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Abstract

This paper evaluates the dynamic effects of various tax instrument on real GDP in Poland, as a contribution to the current debate on public debt sustainability, with a particular focus on determining the tax instrument that can be used to finance public debt without significantly reducing real GDP growth. For this purpose, a structural vector autoregression (SVAR) model of fiscal policy is estimated and random discretionary policy shocks in taxes are identified using an extension of identification scheme by Blanchard and Perotti (2002), based on theoretical and institutional aspects of Polish tax system. The study finds that value added tax is the best instrument to finance public debt in Poland, with the possibility of significant positive short-run output response, compared to personal income and corporate taxes, which both yield significant negative output responses, even in the very short run.

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1 Introduction

As a European transition economy, Poland faces two important goals which may be difficult to combine: steady real GDP growth and a healthy state of public finance, involving sustainable levels of public debt. While Poland's public debt is lower compared to many European countries, the debt to GDP ratio of around 54% is growing towards the 60% constitutional limit, in presence of politically motivated rising government expenditure of the current government. This situation has led many economists to discuss various ways to finance the public debt. Most research however, has not discussed the effects of proposed policies on steady real GDP growth, a crucial goal for a transition economy. Since different taxes have distinct transmission channels that differ among economies, this study aims to contribute empirically to this debate by investigating how increases in various tax instruments affect output, in order to determine which tax instrument can be used to finance growing public debt without contracting output.

This study aims to contribute to current literature in two ways. The first one is the contribution to public debate in Poland, where dynamic effects of particular tax rates on output were not analysed, as most researchers have only focused on forecasting public debt levels and analysing its composition. Another contribution involves the extension of structural VAR estimation and identification procedure of Blanchard and Perotti (2002) onto disaggregated tax series. While existent literature provides some examples of analysis of dynamic effects of disaggregated tax revenues on output, in my view, the assumptions and methodology used in those studies can be improved upon. As an extension, I introduce multiple restrictions based on theoretical and institutional aspects of the tax system structure, price and wage setting, in order to account for contemporaneous changes in some tax revenues as responses to changes in other tax rates, which is not considered in existing literature.

2 Literature Review

2.1 Literature on public debt in Poland

The public debt debate in Poland is largely two sided, with some economists claiming that future tax increases are inevitable and others claiming that current debt levels are sustainable. In general, the two sides agree that fiscal policy decisions in Poland are largely politically motivated. Korniluk (2016), who argues that the stabilising expenditure rule introduced by the previous government has the potential to reduce future debt, maintains that it might fail due to discretionary expenditure decisions influenced by political support. Similarly, studies by Redzepagic & Llorca (2007) and Zabinski (2010) both find evidence of fiscal policy in Poland being influenced by short-term social objectives of governments in power. This is also supported by Kazandziska (2015), who identifies strong pro-cyclicality of fiscal policy in Poland and an econometric study by Mackiewicz and Krajewski (2009), who argue for non-stationarity of government expenditure and irresponsiveness to growing debt. Thus, while there is a consensus about politically motivated use of fiscal policy, some studies e.g. Korniluk (2016) present an optimistic approach about the future, while others, e.g.

Golebiowski & Kozlowski (2011) who investigate the composition of actual and implicit public debt, argue for inevitable tax increases and disturbance of future economic growth.

2.2 Literature on disaggregated SVAR fiscal policy analysis

Existing literature on disaggregated SVAR is limited, and in my view, the present approaches do not properly account for the importance of tax transmission channels. A disaggregated SVAR study by Tenhofen et al (2006), who estimates effects of direct and indirect taxes on the German economy, unfortunately does not distinguish between personal and corporate income taxes in the direct tax category. The authors decide to order indirect taxes first, which is a correct assumption on ordering of corporate tax (CIT) vs. VAT, but upon consideration of tax transmission channels, the ordering is incorrect for personal income tax (PIT) vs. VAT, as contemporaneous shocks in VAT revenue could occur due to disposable income shocks resulting from PIT rate changes. In another disaggregated study, Unal (2015) considers all three of PIT, CIT and VAT revenues, but estimates separate SVAR models for each tax. His study, however, does not to account for contemporaneous effects of a policy shock in one tax affecting another tax revenue, e.g. a shock in PIT rate affecting VAT revenue within a quarter, through disposable income effect. Thus, reduced form residuals of tax revenues should not only be functions of output, government expenditure and their own shocks as in Unal (2015), but also of shocks in other tax rates, as considered in this study.

The SVAR studies of fiscal policy dynamics in Poland only involve approaches using aggregate tax revenues. Notably, Haung, et al (2013) find positive response of output to tax shocks, but the authors claim that their results are driven by PIT reductions and VAT increases over the sample period, suggesting the need for disaggregation in future research. Additionally, Mirdala (2009) proposes an aggregate study of fiscal policy on the region, using Blanchard and Perotti (2002) identification scheme. His result, however, that output does not significantly respond to tax shocks is unreliable due to only 30 observations being used in the analysis.

3 Methodology

3.1 Structural VAR and reduced form models

The structural vector autoregression (SVAR) model used for policy evaluation includes five variables: output y_t, government expenditure g_t, personal income tax revenue τ_p, value added tax revenue τ_v and corporate income tax revenue τ_c, all expressed in natural logarithms of real quantities. Structural VAR extends the standard VAR by allowing for contemporaneous relationships between endogenous variables. It is a good way to model simultaneous relationships in the economy, which allows to overcome endogeneity problems and compute impulse responses to show how the economic system responds to tax shocks. The SVAR can be expressed as:

$$C_0 Y_t = c + C(L) Y_t + \varepsilon_t \tag{3.1.1}$$

where $Y_t \equiv [y_t, g_t, \tau_p, \tau_v, \tau_c]'$ is a 5×1 vector of endogenous variables, c is a 5×1 vector of constants, $C(L)$ is a 5×5 matrix lag polynomial and C_0 is a 5×5 matrix of coefficients from contemporaneous relationships between the elements of Y_t, such that:

$$C_0 = \begin{pmatrix} 1 & -\delta_{yg} & -\delta_{y\tau_p} & -\delta_{y\tau_v} & -\delta_{y\tau_c} \\ -\delta_{gy} & 1 & -\delta_{g\tau_p} & -\delta_{g\tau_v} & -\delta_{g\tau_c} \\ -\delta_{\tau_p y} & -\delta_{\tau_p g} & 1 & -\delta_{\tau_p\tau_v} & -\delta_{\tau_p\tau_c} \\ -\delta_{\tau_v y} & -\delta_{\tau_v g} & -\delta_{\tau_v\tau_p} & 1 & -\delta_{\tau_v\tau_c} \\ -\delta_{\tau_c y} & -\delta_{\tau_c g} & -\delta_{\tau_c\tau_p} & -\delta_{\tau_c\tau_v} & 1 \end{pmatrix} \tag{3.1.2}$$

where δ_{ij} is the contemporaneous effect of a change in variable j on variable i. Finally, $\varepsilon_t \equiv [\varepsilon_t^y, \varepsilon_t^g, \varepsilon_t^p, \varepsilon_t^v, \varepsilon_t^c]'$ is a 5×1 vector of structural shocks to each variable, which in the SVAR are assumed to be orthogonal with covariance matrix $\sum_\varepsilon$, such that:

$$\mathbb{E}(\varepsilon_t \varepsilon_t') = \sum\nolimits_\varepsilon = \begin{pmatrix} \sigma_{\varepsilon_t^y} & 0 & 0 & 0 & 0 \\ 0 & \sigma_{\varepsilon_t^g} & 0 & 0 & 0 \\ 0 & 0 & \sigma_{\varepsilon_t^p} & 0 & 0 \\ 0 & 0 & 0 & \sigma_{\varepsilon_t^v} & 0 \\ 0 & 0 & 0 & 0 & \sigma_{\varepsilon_t^c} \end{pmatrix} \tag{3.1.3}$$

The contemporaneous terms in the SVAR cannot be estimated by OLS, as it would be subject to simultaneity bias. For this purpose, an estimatable reduced form representation is derived by pre-multiplying (3.1.1) by C_0^{-1}:

$$Y_t = C_0^{-1}c + C_0^{-1}C(L)Y_t + C_0^{-1}\varepsilon_t \tag{3.1.4}$$

which can be rewritten as:

$$Y_t = d + D(L)Y_t + u_t \tag{3.1.5}$$

where $D(L) = C_0^{-1}C(L)$, $d = C_0^{-1}c$ and $u_t \equiv [u_t^y, u_t^g, u_t^p, u_t^v, u_t^c]' = C_0^{-1}\varepsilon_t$. The vector of reduced form residuals u_t is obtained from estimated model (3.1.5) and may have non-zero covariances between them, with covariance matrix given by:

$$\mathbb{E}(u_t u_t') = \sum\nolimits_u = \begin{pmatrix} \sigma_{u_t^y} & cov(u_t^y, u_t^g) & cov(u_t^y, u_t^p) & cov(u_t^y, u_t^v) & cov(u_t^y, u_t^c) \\ cov(u_t^g, u_t^y) & \sigma_{u_t^g} & cov(u_t^g, u_t^p) & cov(u_t^g, u_t^v) & cov(u_t^g, u_t^c) \\ cov(u_t^p, u_t^y) & cov(u_t^p, u_t^g) & \sigma_{u_t^p} & cov(u_t^p, u_t^v) & cov(u_t^p, u_t^c) \\ cov(u_t^v, u_t^y) & cov(u_t^v, u_t^g) & cov(u_t^v, u_t^p) & \sigma_{u_t^v} & cov(u_t^v, u_t^c) \\ cov(u_t^c, u_t^y) & cov(u_t^c, u_t^g) & cov(u_t^c, u_t^p) & cov(u_t^c, u_t^v) & \sigma_{u_t^c} \end{pmatrix} \tag{3.1.6}$$

The aim of identification procedure is to find a matrix C_0^{-1} such that $u_t = C_0^{-1}\varepsilon_t$ and $\mathbb{E}(u_t u_t') = \sum_u$. The second condition can be rewritten as follows:

$$\mathbb{E}(u_t u_t') = \mathbb{E}(C_0^{-1}\varepsilon_t\varepsilon_t' C_0^{-1'}) = C_0^{-1}\mathbb{E}(\varepsilon_t\varepsilon_t')C_0^{-1'} = C_0^{-1}\sum\nolimits_\varepsilon C_0^{-1'} \tag{3.1.7}$$

Once such matrix C_0^{-1} is found to satisfy these conditions, it is then possible to recover the parameters of structural model (3.1.1) and compute structural impulse response functions, as outlined in Section 5.3.

3.2 Identification procedure

The nature of tax-output relationships, which involves automatic responses of tax revenues to output shocks, as well as contemporaneous responses of output to tax shocks, rules out the use of Choleski decomposition as it is impossible to argue for a respective ordering of taxes vs. output. Instead, I pursue and extended version of Blanchard and Perotti (2002) identification scheme, which exploits tax to output elasticities, in addition to institutional relationships present in the economy. Blanchard and Perotti (2002) note that reduced form residuals u_t^i, for $i = g, p, v, c$ can be understood as linear combinations of three components: i) automatic responses of tax revenues to output shocks, ii) systematic discretionary responses of expenditure and revenues and iii) random discretionary policy shocks. Using quarterly data, it is possible to rule out case ii), as government legislation takes longer to pass than a quarter and thus, there are no within-quarter systematic discretionary government expenditure and tax policy responses to output shocks. In this case, without loss of generality, the residuals from estimated reduced form model (3.1.5) can be expressed as linear combinations of reduced form shocks u_t^i and random discretionary policy shocks ε_t^i:

$$u_t^y = \alpha_{yg} u_t^g + \alpha_{yp} u_t^p + \alpha_{yv} u_t^v + \alpha_{yc} u_t^c + \varepsilon_t^y \tag{3.2.1}$$

$$u_t^g = \alpha_{gy} u_t^y + \beta_{gp} \varepsilon_t^p + \beta_{gv} \varepsilon_t^v + \beta_{gc} \varepsilon_t^c + \varepsilon_t^g \tag{3.2.2}$$

$$u_t^p = \alpha_{py} u_t^y + \beta_{pg} \varepsilon_t^g + \beta_{pv} \varepsilon_t^v + \beta_{pc} \varepsilon_t^c + \varepsilon_t^p \tag{3.2.3}$$

$$u_t^v = \alpha_{vy} u_t^y + \beta_{vg} \varepsilon_t^g + \beta_{vp} \varepsilon_t^p + \beta_{vc} \varepsilon_t^c + \varepsilon_t^v \tag{3.2.4}$$

$$u_t^c = \alpha_{cy} u_t^y + \beta_{cg} \varepsilon_t^g + \beta_{cp} \varepsilon_t^p + \beta_{cv} \varepsilon_t^v + \varepsilon_t^c \tag{3.2.5}$$

which can be rewritten in a matrix form $Au_t = B\varepsilon_t$:

$$\begin{pmatrix} 1 & -\alpha_{yg} & -\alpha_{yp} & -\alpha_{yv} & -\alpha_{yc} \\ -\alpha_{gy} & 1 & 0 & 0 & 0 \\ -\alpha_{py} & 0 & 1 & 0 & 0 \\ -\alpha_{vy} & 0 & 0 & 1 & 0 \\ -\alpha_{cy} & 0 & 0 & 0 & 1 \end{pmatrix} \begin{pmatrix} u_t^y \\ u_t^g \\ u_t^p \\ u_t^v \\ u_t^c \end{pmatrix} = \begin{pmatrix} 1 & 0 & 0 & 0 & 0 \\ 0 & 1 & \beta_{gp} & \beta_{gc} & \beta_{gv} \\ 0 & \beta_{pg} & 1 & \beta_{pc} & \beta_{pv} \\ 0 & \beta_{vg} & \beta_{vp} & 1 & \beta_{vc} \\ 0 & \beta_{cg} & \beta_{cp} & \beta_{cv} & 1 \end{pmatrix} \begin{pmatrix} \varepsilon_t^y \\ \varepsilon_t^g \\ \varepsilon_t^p \\ \varepsilon_t^v \\ \varepsilon_t^c \end{pmatrix} \tag{3.2.6}$$

where the $Au_t = B\varepsilon_t$ representation of equations (3.2.1)-(3.2.5) allows to find the matrix C_0^{-1} using the substitution $C_0^{-1} = A^{-1}B$, such that $u_t = A^{-1}B\varepsilon_t = C_0^{-1}\varepsilon_t$. In matrices forming equation (3.2.6), coefficients α_{ij} represent the response of variable i due to a reduced form shock in variable j, while coefficients β_{ij} represents the contemporaneous response in variable i due to a random discretionary (or "structural") shock in variable j.

Equation (3.2.1) reflects an assumption that output contemporaneously responds to shocks in all tax revenues and government expenditure, while equations (3.2.3)-(3.2.5) allow tax revenues and government expenditure to respond contemporaneously to random policy shocks in various fiscal policy instruments ε_t^i, but not to every contemporaneous shock in tax revenues (given by u_t^i) which can be due to factors other than policy changes. For example, a CIT revenue shock could be due to VAT rate change affecting corporate profits, but not necessarily due to any contemporaneous shock in VAT revenue, which may occur due to output shocks that also affect CIT revenue, and thus, be captured by $\alpha_{cy}u_t^y$ instead of $\alpha_{cv}u_t^v$. Therefore, $\beta_{ij}\varepsilon_t^j$ terms are included in (3.2.2)-(3.2.5) instead of $\alpha_{ij}u_t^j$. The next step requires the restriction of parameters on matrices A and B, where necessary and sufficient conditions require $m(m-1)/2$ restrictions imposed on A and B, where m is the number of model parameters, which gives 10 restrictions required on matrices A and B in total (Canova, 2007).

The first restriction in equation (3.2.2) sets $\alpha_{gy} = 0$, as the government's legislative process to alter expenditure as a response to output shocks is longer than a quarter, in addition to considering the historical evidence that in Poland, tax policy decisions are usually made to finance previously planned government spending decisions, and thus, government expenditure does not contemporaneously respond to tax revenue shocks, implying that $\beta_{gp} = \beta_{gc} = \beta_{gv} = 0$. Considering equation (3.2.3), I assume that PIT revenue does not contemporaneously respond to a random policy shock in CIT revenue, that is, $\beta_{pc} = 0$. Since corporate profits are subject to taxation in Poland, a decision to raise CIT rates would lower corporate profits, which could lead to price increases over time, depending on their flexibility. The PIT revenue, however, depends on individual income and not on corporate profits or prices. While firms could, over time, reduce employees' wages to recover profits, due to contracting agreements in Poland being usually longer than a quarter, this is unlikely to occur, which justifies assuming that $\beta_{pc} = 0$ (Strzelecki and Wyszynski, 2016). It is also plausible to assume that $\beta_{pv} = 0$, because if wages are constant within a quarter, a VAT rate change which affects firm's prices and/or profits, has no effect on individual wages and thus, on PIT revenue. In equation (3.2.4), it is assumed that a policy change in CIT rate has no within-quarter effect on VAT revenue. The only way in which VAT revenue could be affected due to a CIT shock would be through firms raising their prices in order to increase profits on demand-inelastic products. This effect, however, is unlikely within a quarter, as evidence from Jankiewicz and Kolodziejczyk (2008) shows that firms face short-run price rigidities and therefore, it is likely that $\beta_{vc} = 0$ with firms initially absorbing the CIT rate changes. In contrast, a policy shock in PIT rate is likely to have an effect on VAT revenue, because a rise in PIT rate reduces disposable income and consumption, decreasing the VAT revenue which depends on consumption, and thus, β_{vp} will be left unrestricted. In equation (3.2.5), the β_{cj} parameters for $j = p, v, g$ will be left unrestricted. In the first case, a rise in PIT rate, reduces disposable income, consumption and firm's profits, leading to a reduction in CIT revenue. In the second case, under presence of price rigidities, a rise in VAT rate likely has an effect on profits and hence β_{vp} is left unrestricted. Finally, random government expenditure shocks can have an effect on random tax policy decisions introduced afterwards to e.g. finance increased expenditure, and hence β_{pg}, β_{cg} and β_{vg} will be unrestricted.

According to Blanchard and Perotti (2002), automatic changes in tax revenues due to contemporaneous output shocks can be approximated by setting α_{iy} parameters for $i = p, v, c$ equal to tax revenue to output elasticities for each instrument, the estimation of which is described in detail in Section 5. Together with α_{gy}, β_{gp}, β_{gc}, β_{gv}, β_{pc}, β_{pv} and β_{vc} set equal to zero, the three elasticity restrictions on α_{cy}, α_{vy} and α_{cy} complete the requirement for 10 restrictions on matrices A and B

together. Thus, equation (3.2.6) becomes:

$$\begin{pmatrix} 1 & -\alpha_{yg} & -\alpha_{yp} & -\alpha_{yv} & -\alpha_{yc} \\ 0 & 1 & 0 & 0 & 0 \\ -\alpha^*_{py} & 0 & 1 & 0 & 0 \\ -\alpha^*_{vy} & 0 & 0 & 1 & 0 \\ -\alpha^*_{cy} & 0 & 0 & 0 & 1 \end{pmatrix} \begin{pmatrix} u^y_t \\ u^g_t \\ u^p_t \\ u^v_t \\ u^c_t \end{pmatrix} = \begin{pmatrix} 1 & 0 & 0 & 0 & 0 \\ 0 & 1 & 0 & 0 & 0 \\ 0 & \beta_{pg} & 1 & 0 & 0 \\ 0 & \beta_{vg} & \beta_{vp} & 1 & 0 \\ 0 & \beta_{cg} & \beta_{cp} & \beta_{cv} & 1 \end{pmatrix} \begin{pmatrix} \varepsilon^y_t \\ \varepsilon^g_t \\ \varepsilon^p_t \\ \varepsilon^v_t \\ \varepsilon^c_t \end{pmatrix} \tag{3.2.7}$$

where * denotes coefficients set equal to exogenous elasticities. Equation (3.2.7) can be rewritten as the following restricted system:

$$u^g_t = \varepsilon^g_t \tag{3.2.8}$$

$$u^p_t = \alpha^*_{py} u^y_t + \beta_{pg}\varepsilon^g_t + \varepsilon^p_t \tag{3.2.9}$$

$$u^v_t = \alpha^*_{vy} u^y_t + \beta_{vg}\varepsilon^g_t + \beta_{vp}\varepsilon^p_t + \varepsilon^v_t \tag{3.2.10}$$

$$u^c_t = \alpha^*_{cy} u^y_t + \beta_{cg}\varepsilon^g_t + \beta_{cp}\varepsilon^p_t + \beta_{cv}\varepsilon^v_t + \varepsilon^c_t \tag{3.2.11}$$

$$u^y_t = \alpha_{yg} u^g_t + \alpha_{yp} u^p_t + \alpha_{yv} u^v_t + \alpha_{yc} u^c_t + \varepsilon^y_t \tag{3.2.12}$$

Given exogenous α^*_{iy} for $i = p, v, c$, it is possible to construct *cyclically adjusted* fiscal policy shocks (Blanchard and Perotti, 2002). Cyclically adjusted shocks are defined as $u^{i,CA}_t \equiv u^i_t - \alpha^*_{iy} u^y_t$ for $i = g, p, v, c$, and applying this for (3.2.8)-(3.2.11) gives:

$$u^{g,CA}_t = \varepsilon^g_t \tag{3.2.13}$$

$$u^{p,CA}_t = \beta_{pg}\varepsilon^g_t + \varepsilon^p_t \tag{3.2.14}$$

$$u^{v,CA}_t = \beta_{vg}\varepsilon^g_t + \beta_{vp}\varepsilon^p_t + \varepsilon^v_t \tag{3.2.15}$$

$$u^{c,CA}_t = \beta_{cg}\varepsilon^g_t + \beta_{cp}\varepsilon^p_t + \beta_{cv}\varepsilon^v_t + \varepsilon^c_t \tag{3.2.16}$$

$$u^y_t = \alpha_{yg} u^g_t + \alpha_{yp} u^p_t + \alpha_{yv} u^v_t + \alpha_{yc} u^c_t + \varepsilon^y_t \tag{3.2.17}$$

Equations (3.2.14)-(3.2.16) can now be estimated recursively by OLS, since $cov(\varepsilon^i_t, \varepsilon^j_t) = 0$ for any $i \neq j$ in accordance with $\sum_\varepsilon$. Equation (3.2.17) can be estimated by IV, using ε^g_t, ε^p_t, ε^v_t and ε^c_t as instruments for u^g_t, u^p_t, u^v_t and u^c_t respectively, since $cov(\varepsilon^i_t, \varepsilon^y_t) = 0$ for any $i \neq y$. In this manner, all parameters from A and B matrices in equations (3.2.6) and (3.2.7) are identified, providing estimates of C_0^{-1} and $\sum_\varepsilon$, as required for estimation of structural impulse response functions.

4 Data

4.1 Structural VAR variables

As discussed in Section 3.1, five variables form the structural VAR: output y_t, government expenditure g_t, personal income tax revenue τ_p, value added tax revenue τ_v and corporate tax revenue τ_c, all divided by GDP deflator normalised to 2002Q1 and expressed using natural logarithms. Logarithmic transformation allows for estimation of fiscal policy effects on real output to be expressed in terms of percentage changes. Quarterly data on nominal GDP, GDP deflator and government expenditure is obtained from Eurostat, ranging from 2002Q1 to 2017Q2, while data on tax revenues is obtained from Central Statistical Office of Poland, with the same time range. The number of observations (62) is small and thus can question the consistency of OLS estimators used to estimate the reduced form VAR. In this case, however, the small number of observations has to be accepted, as quarterly GDP has been recorded in Poland since 2002Q1, while meaningful annual data is limited to observations since Poland's conversion to capitalism in 1989. Upon inspection of time series plots in Figures 1 and 2, all series present positive linear trends, with tax variables presenting a notable pattern of seasonality, even after seasonal adjustment using dummy variables.[1] This case was noted in the original study of Blanchard and Perotti (2002) and it involves the quarterly dependence of tax collections on GDP. When most taxes are collected in quarter 2, the collected revenue depends on GDP throughout the past year, while the remaining taxes collected in other quarters do not depend on GDP. As a solution, Blanchard and Perotti (2002) propose a VAR with quarter dependent coefficients, which will be discussed further in Section 5.2.

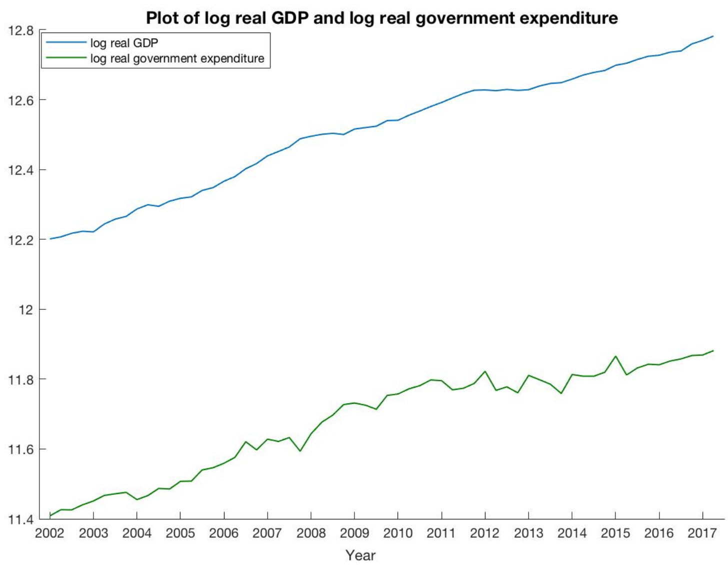


Figure 1: Plot of real GDP vs. real government expenditure (logs)

[1]Original tax series obtained from Central Statistical Office were not seasonally adjusted.

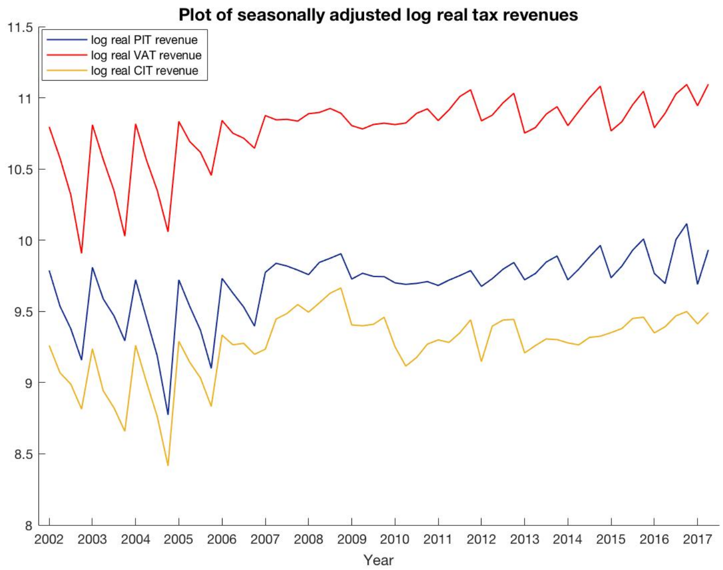


Figure 2: Plot of real tax revenues (logs)

Unit root testing was carried out using ADF tests, with lags selected according to PACFs and lag significance[2]. Deterministic trends were included in all tests, as positive trends are visible upon inspection. As shown in Appendices 2 and 3, all series except for output were declared I(1) at 10% significance level. Tax series were problematic, as tests with 4 lags declared them as I(1), but showed insignificant lags 1-3, with the only fully significant regression having zero lags, which instead declared tax series as I(0). Nevertheless, I treat these series as I(1) to avoid spurious regressions. Output series was initially found to be I(2), but upon observing a structural break during 2008 financial crisis, Perron's variant of ADF test found output to be I(1) at 5% significance level.

4.2 Variables used for exogenous elasticity estimation

According to Perotti (2005), exogenous tax to output elasticities, denoted by α_{iy} for $i = p, v, c$ are defined as the product of tax to tax base elasticity $\epsilon_{i,ib}$ and tax base to output elasticity $\epsilon_{ib,y}$, such that for $i = p, v, c$:

$$\alpha_{iy} = \epsilon_{i,ib} \times \epsilon_{ib,y} \tag{4.2.1}$$

In (4.2.1), tax to tax base elasticities $\epsilon_{i,ib}$ are obtained from Price, Dang and Boetv (2015), while tax base to output elasticities $\epsilon_{ib,y}$ are obtained from regressions of tax base on output (estimation discussed in Section 5.1). An exception to this is VAT tax to output elasticity α_{vy}, which was set to 1.00. This is because VAT is a tax on total expenditure E with tax rate t_v, such that VAT revenue is defined as:

$$\tau_v = t_v E \tag{4.2.2}$$

[2]PACF's in Appendix 1

Since output y is equal to expenditure E, taking logs of both sides of equation (4.2.2) and differentiating τ_v with respect to y, yields:

$$ln(\tau_v) = ln(t_v) + ln(E) \tag{4.2.3}$$

$$\frac{\partial ln(\tau_v)}{\partial ln(y)} = \alpha_{vy} = 1 \tag{4.2.4}$$

Personal income tax base is the aggregate income subject to taxation, defined as the average wage $\overline{w}$ multiplied by the number of employed persons N:

$$PITbase = \overline{w}N \tag{4.2.5}$$

In order to construct (4.2.5), the data on $\overline{w}$ and N was obtained from Central Statistical Office of Poland and seasonally adjusted using dummy variables. Upon inspection from Figure 3, the log series exhibits a positive trend, with ACF and PACF suggesting a unit root. Based on the PACF, the ADF test included three lags and deterministic trend, identifying the series as I(1). The corporate tax base is equal to aggregate corporate profits, as these are subject to corporate taxation in Poland. Since data on exact corporate profits is unavailable, gross operating surplus is used as a proxy. The gross operating surplus series, denoted by $CITbase$ is obtained from Eurostat in seasonally adjusted quantities. Based on time series plot in Figure 3, the log series exhibits a positive trend, with ACF and PACF suggesting a unit root. Consequently, ADF test with one lag and deterministic trend was carried out, identifying the series as I(1).

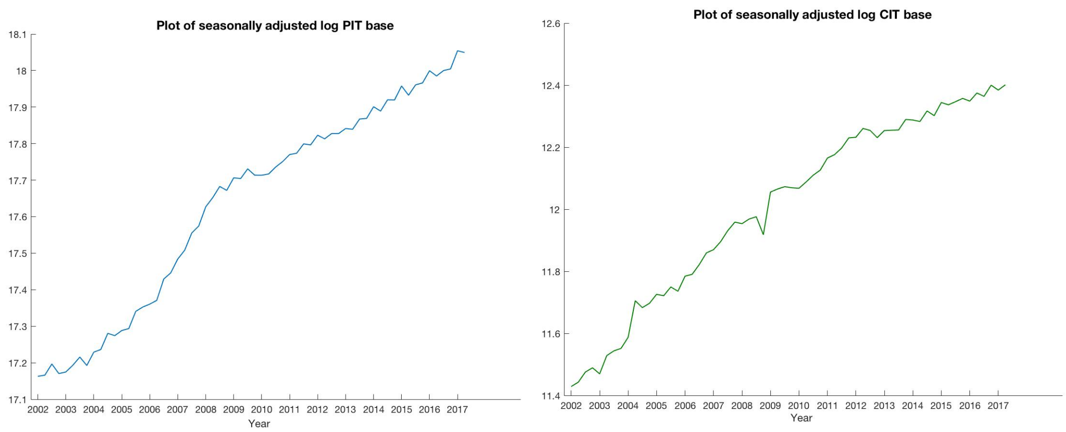


Figure 3: Plots of log PIT and CIT bases

5 Estimation and Results

5.1 Estimation of exogenous tax base to output elasticities

The tax base to output elasticities $\epsilon_{ib,y}$ were estimated through log tax base on log output regressions. Since $lnPITbase$ series was found to be I(1), and there was no cointegration between

$lnPITbase$ and output y, I estimate a first difference model[3]:

$$\Delta ln(PITbase)_t = \delta_0 + \sum_{k=0}^{K} \delta_{k+1} \Delta ln(y)_{t-k} + \sum_{i=1}^{p} \delta_{K+i} \Delta ln(PITbase)_{t-i} + \varepsilon_t \qquad (5.1.1)$$

where p and K are chosen according to PACF in Appendix 5 and lag significance. Results for this regression are presented in Appendix 8. While 2 lags were significant, serial correlation was present for lags 1-2, yielding biased estimates. In this case, to ensure consistency I choose the model with zero lags and use Newey-West standard errors, with estimated coefficient of 0.73.

In the CIT regression, cointegration was found between $lnCITbase$ and output y, and therefore I estimate an ECM:

$$\Delta ln(CITbase)_t = \delta_0 + \sum_{k=0}^{K} \delta_{k+1} \Delta ln(y)_{t-k} + \sum_{i=1}^{p} \delta_{K+i} \Delta ln(CITbase)_{t-i} + \gamma \hat{\varepsilon}_{t-1} + \varepsilon_t \qquad (5.1.2)$$

where $\hat{\varepsilon}_{t-1}$ is an estimated residual from static regression:

$$ln(CITbase)_t = \phi_0 + \phi_1 ln(y)_t + \varepsilon_t \qquad (5.1.3)$$

ECM results are displayed in Appendix 9, where the preferred specification yields a coefficient of 1.57. The elasticities used in SVAR identification (denoted by *) are summarised in Table 1:

Table 1: Tax to output elasticities for SVAR identification (*)

Elasticity	Tax/Base	Base/Output	Tax/Output
	$\epsilon_{i,ib}$	$\epsilon_{ib,y}$	$\alpha_{iy} = \epsilon_{i,ib} \times \epsilon_{ib,y}$
PIT	1.93	0.73	1.41*
CIT	2.30	1.57	3.61*
VAT			1.00*

5.2 Estimation of reduced form VAR

As discussed in Section 4.1, all series are I(1), raising the possibility of spurious regressions, and to resolve this, I look for cointegrating vectors using Johansen's test for multivariate cointegration with unrestricted constant specification. Schwartz Bayesian Information Criterion suggested that reduced form VAR(1) is the preferred model, and thus, one lag was used in Johansen's test with unrestricted constant. The test found a single linear combination which is I(0) (cointegrating rank of 1), implying that VAR(1) yields consistent estimates[4]. In this case, a common procedure involves the estimation of a VECM, which takes the cointegrating vector into account. In this case, however, to maintain comparability with existing literature, I pursue a structural VAR approach. Additionally, considering that each series was identified to have a positive trend, a 5×1 vector of trends was

[3]see Appendix 7 for cointegration tests

[4]See Appendices 10 and 11 for lag selection and Johansen's test

included to avoid spurious relationships due to deterministic trends.

The seasonality of tax series is a serious issue in estimation of C_0^{-1}, where regressions (3.2.14)-(3.2.17) that form coefficients in C_0^{-1} require stationary and serially uncorrelated residuals. As a solution, Blanchard and Perotti (2002) propose the inclusion of both additive and multiplicative dummy variables in each regression, such that all coefficients in the VAR are dependent on each quarter. As shown in time series plots, sample ACF, PACF and stationarity tests in Appendices 12-14, use of multiplicative dummies (quarter dependence) largely removes serial correlation and nonstationarity, which are both present in a model without quarter dependence. In particular, the ACF and PACF for residuals from model with quarter dependence present barely significant lags and ADF tests suggest that all residuals are I(0), thus satisfying necessary criteria for the residuals to be used for regressions that form C_0^{-1} and showing that seasonality of tax series is the likely reason for serial correlation and nonstationarity in tax residuals instead of lag order misspecification. With these modifications, the final estimated VAR(1) regression took the following form:

$$Y_t = d + \beta q + D_{11}Y_{t-1} + D_{12}(Y_{t-1} \times q_2) + D_{13}(Y_{t-1} \times q_3) + D_{14}(Y_{t-1} \times q_4) + \gamma t + u_t \qquad (5.2.1)$$

where $q \equiv [q_2, q_3, q_4]'$ is a 3×1 vector of quarterly dummy variables with associated 5×3 coefficient matrix β, where quarter 1 is the reference quarter, t is a 5×1 vector of linear trends for each variable with associated 5×5 matrix γ, d is a 5×1 vector of constants and u_t is a 5×1 vector of reduced form residuals, discussed in Section 3.2. Estimation of model (5.2.1) implies that all results, involving structural impulse response functions, will be quarter dependent. Since this study, along with much of the literature is concerned with average impulse response functions to tax policy shocks, I proceed with the approach of Blanchard and Perotti (2002) and use (5.2.1) to estimate the structural matrix C_0^{-1} and the structural covariance matrix $\sum_\varepsilon$, and then use the model without quarter dependent coefficients to estimate the average, quarter-independent structural impulse responses. This model is given by:

$$Y_t = d + \beta q + D_1 Y_{t-1} + \gamma t + u_t \qquad (5.2.2)$$

Estimates of the matrix C_0^{-1} from (5.2.1) and structural covariance matrix $\sum_\varepsilon$ are given in Appendix 15, where notably, the structural covariance matrix presents variances of structural shocks on the diagonal and zero covariances between structural shocks, as required for an appropriate identification procedure. These estimates, along with D_1 coefficient matrix from (5.2.2), will be used for identification as outlined in Section 3.2 and computation of structural impulse response functions in Section 5.3.

5.3 Results: Structural Impulse Response Functions and Evaluation

The structural impulse response functions were calculated according to the method outlined in Kilian and Lutkepohl (2017). Given the identified matrix C_0^{-1} and H coefficients from a moving average representation of reduced form VAR(1), Ψ_i, the structural impulse response matrices Θ_i are calculated by post-multiplying each Ψ_i by C_0^{-1}:

$$\Theta_1 = \Psi_1 C_0^{-1} = I_m C_0^{-1} = C_0^{-1}$$
$$\Theta_2 = \Psi_2 C_0^{-1}$$
$$\vdots$$
$$\Theta_H = \Psi_H C_0^{-1}$$

Since there are 5 variables, Ψ_i and Θ_i are 5×5 matrices, implying 25 structural impulse response functions, where a single structural IRF coefficients $\theta_{jk,i}$ are jk'th elements of each consecutive matrix Θ_i for $i = 1, ..., H$. The structural IRFs for the main shocks of interest are plotted in Figure 4 below. Since all variables are in logs of real quantities, the coefficients of structural IRF's, $\theta_{jk,i}$, can be interpreted as elasticities, that is, percentage changes of output due to 1% shock in tax revenues or government expenditure.

As depicted in Figure 4, a structural shock in PIT revenue causes a sharp decline in output, which reaching a minimum of -0.0523% in Q2. An explanation for this is a fall in consumption that results from a fall in disposable income, which is further supported in Figure 5, where a structural PIT shock causes a fall in VAT revenue, reflecting a fall in consumption as VAT is a tax on expenditure. In addition, PIT shock has an initially positive impact on profits and hence, the CIT revenue, as firms reduce their labour demand and employment. In the long run, however, corporate profits and CIT revenue decline due to falling consumption, as shown in response of CIT revenue to a PIT shock in Figure 5. Similarly, a structural CIT shock adversely affects firms' profits and investment, causing a fast decline in output and reaching a minimum of -0.0278% in Q1, which is smaller compared to a PIT shock. Furthermore, a structural CIT shock causes a sharp decline in PIT revenue, through falling wages and employment, which starts after three months, coinciding with most contract termination periods. The PIT revenue, following a CIT shock quickly returns to around zero, thus giving a possible explanation of firms restructuring their production process following a profit tax shock, which can have a short, but significant impact on wages and employment in the process.

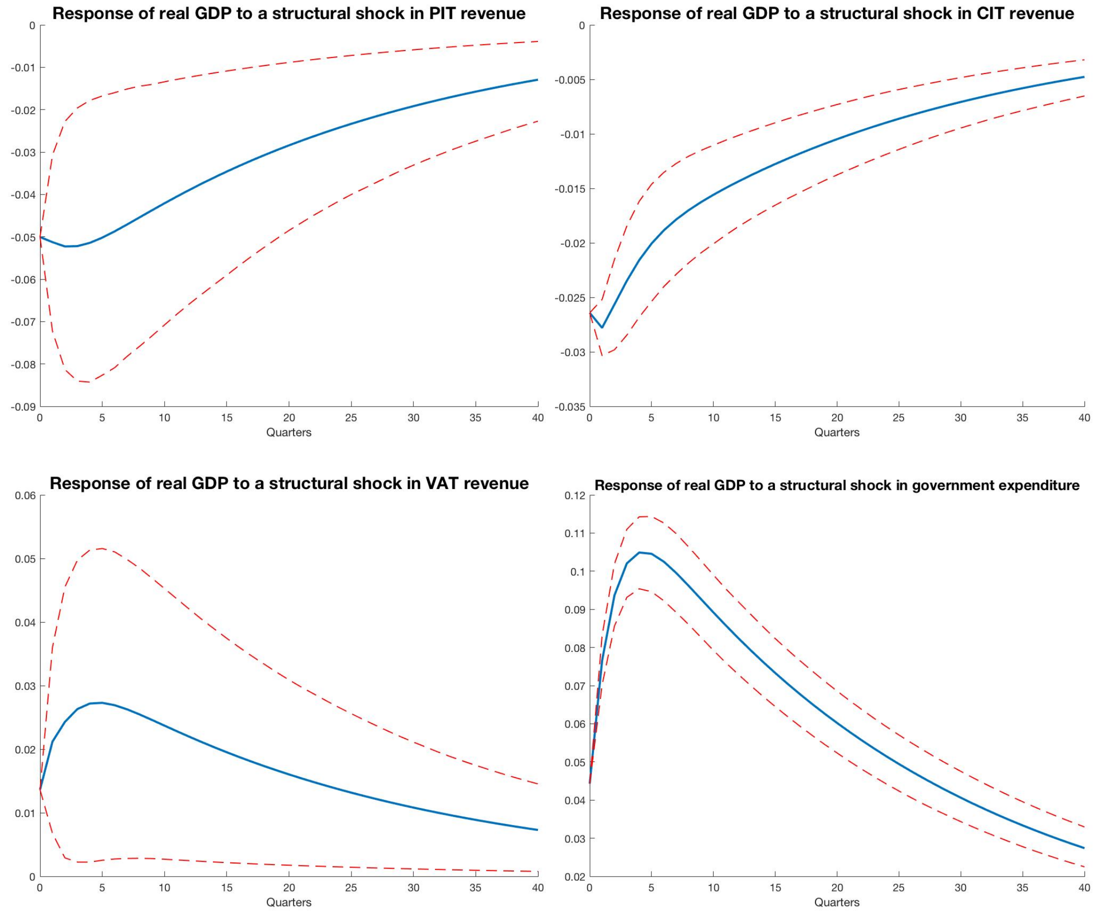


Figure 4: Structural impulse response functions for tax revenues and government expenditure shocks

A particularly interesting result is the positive response of output to a structural VAT shock, which peaks at 0.0273% in Q5. An explanation of this effect lies within firm behaviour, as suggested by response of CIT revenue to a VAT shock in Figure 5, where the shape of the IRF reflects that VAT increase causes initially falling (until Q4) and then rising profits. This timing coincides with evidence from Jankiewicz and Kolodziejczyk (2008), who identify that most firms in Poland change their prices once a year, with little firms doing so more often than quarterly. Their study also identifies that reasons for lack of immediate price responses involves menu costs and implicit contracts of firms with their consumers. This leads to final good sellers being reluctant to change prices immediately following a VAT increase, and thus, initially accepting the VAT burden. In contrast, upstream producers who possess more market power are able to increase their prices more quickly (e.g. in Q1), leading to falling profits of final good sellers, meaning that final good sellers decide to initially bear the tax burden. At the same time, as consumers expect prices to rise in future as a result of VAT increase, they increase their present consumption and investment, which results in a positive response of real GDP.

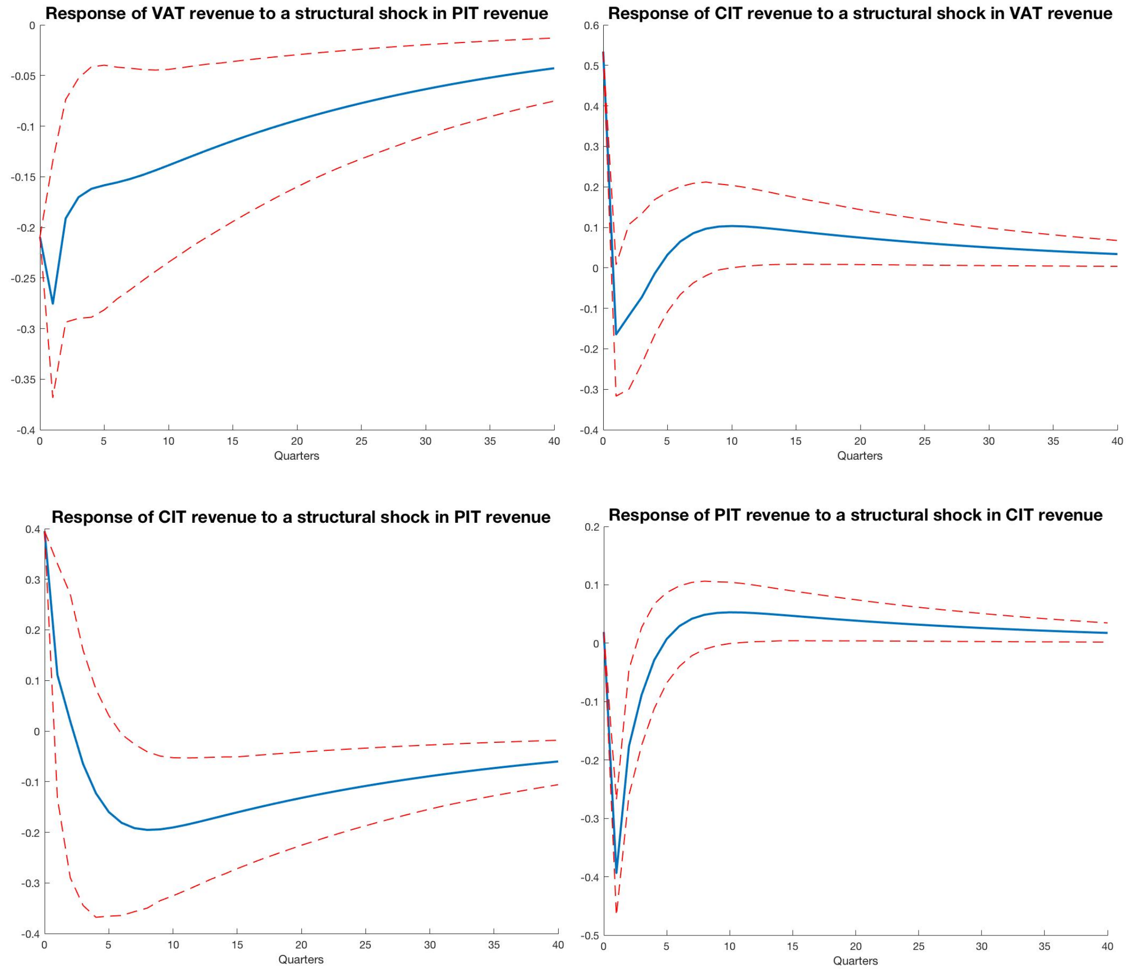


Figure 5: Additional relevant structural impulse responses

Therefore, this result presents that under such characteristics of the economy, VAT is the best way to finance the public debt without causing a contractionary effect on output. It should also be noted that most signs and magnitudes of tax and government expenditure (around 0.1%) shocks on output are comparable to existing literature, particularly the evidence from Germany by Tenhofen et al (2006). In order to assess the suitability of structural impulse response functions for statistical inference, 95% confidence intervals were computed using bootstrapping procedure with 2000 replications, as outlined in Kilian and Lutkepohl (2017)[5]. As shown in Figure 4, the bootstrap confidence intervals, plotted as red dotted lines, suggest that positive response of output to VAT shocks and negative response to PIT and CIT shocks are both significant, which contrasts with existing literature that mostly finds insignificant effects of indirect taxes (e.g. VAT) on output.

While the results depicted in Figure 4 rely on restrictions justified by empirical information on the structure of tax transition mechanisms, timing of policy changes and characteristics of price

[5]The bootstrapping procedure used in this study, along with my own Matlab code used for this is outlined in Appendix 16

and wage rigidity, there exist alternative ordering assumptions. One includes abandoning the fixed quarterly price assumption and instead, ordering $Y_t \equiv [y_t, g_t, \tau_p, \tau_c, \tau_v]'$ by assuming that $\beta_{cv} = 0$, that is, structural shocks in VAT revenue have no contemporaneous effect on CIT revenue, meaning that firms quickly adjust their prices, such that profits are unaffected. Although in my view, short-run corporate profits are affected by VAT innovations, this alternative assumption of flexible price adjustments has no impact on the sign of output responses and only affects the magnitudes of output changes, as demonstrated in Figure 5, which presents structural IRFs from different possible orderings on the same graph. Similarly, another possibility involves arguing that government expenditure decisions respond to prior changes in tax decisions, that is, ordering $Y_t \equiv [y_t, \tau_p, \tau_c, \tau_v, g_t]'$ and imposing $\beta_{pg}, \beta_{vg}, \beta_{cg} = 0$. While the presence of budget deficit throughout the sample period and timing of particular policy changes suggest this is not the case, even if it would be, the results remain unchanged apart from the magnitudes. Thus, even considering such alternative possibilities, which are possible, but not necessarily empirically supported, the important result that VAT results in a small positive shock is robust to such orderings.

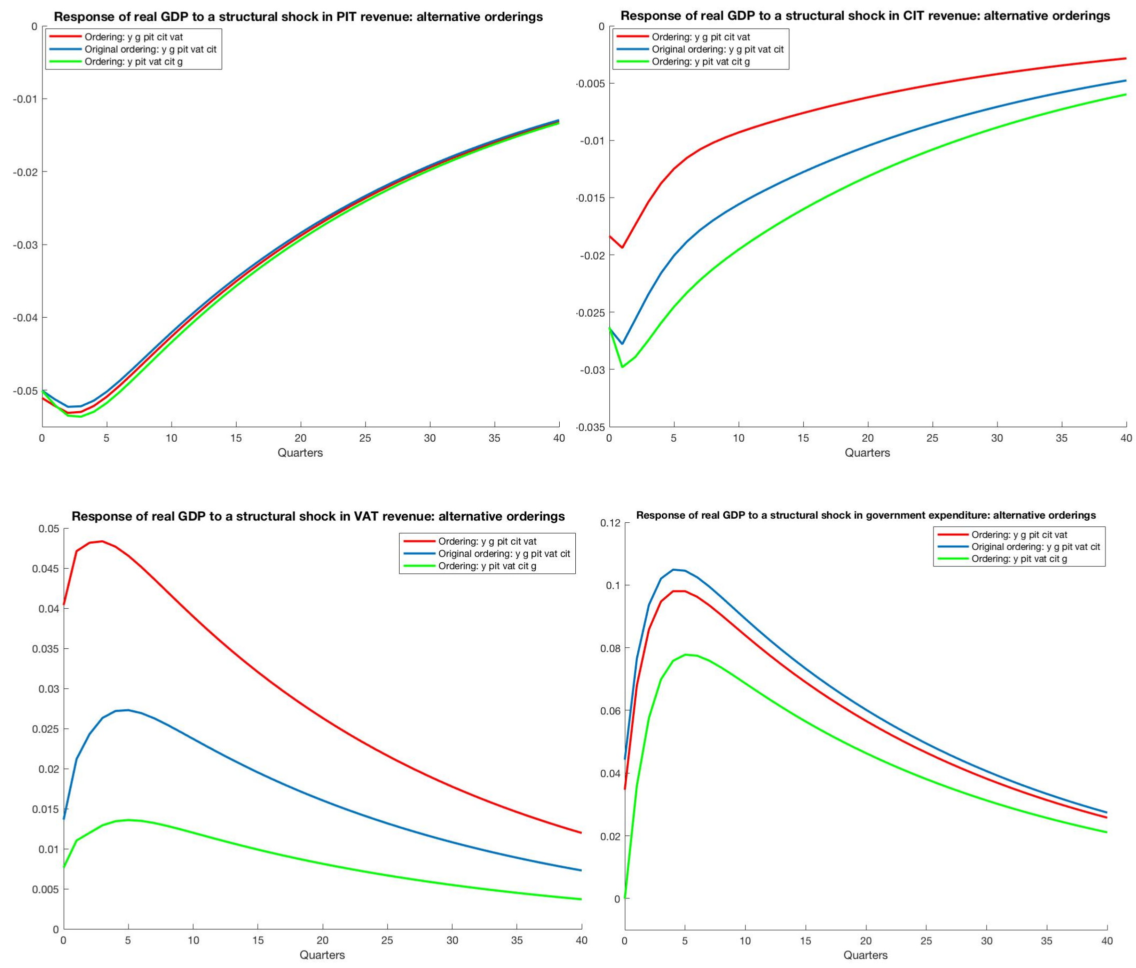


Figure 6: Key structural impulse response functions under alternative proposed orderings

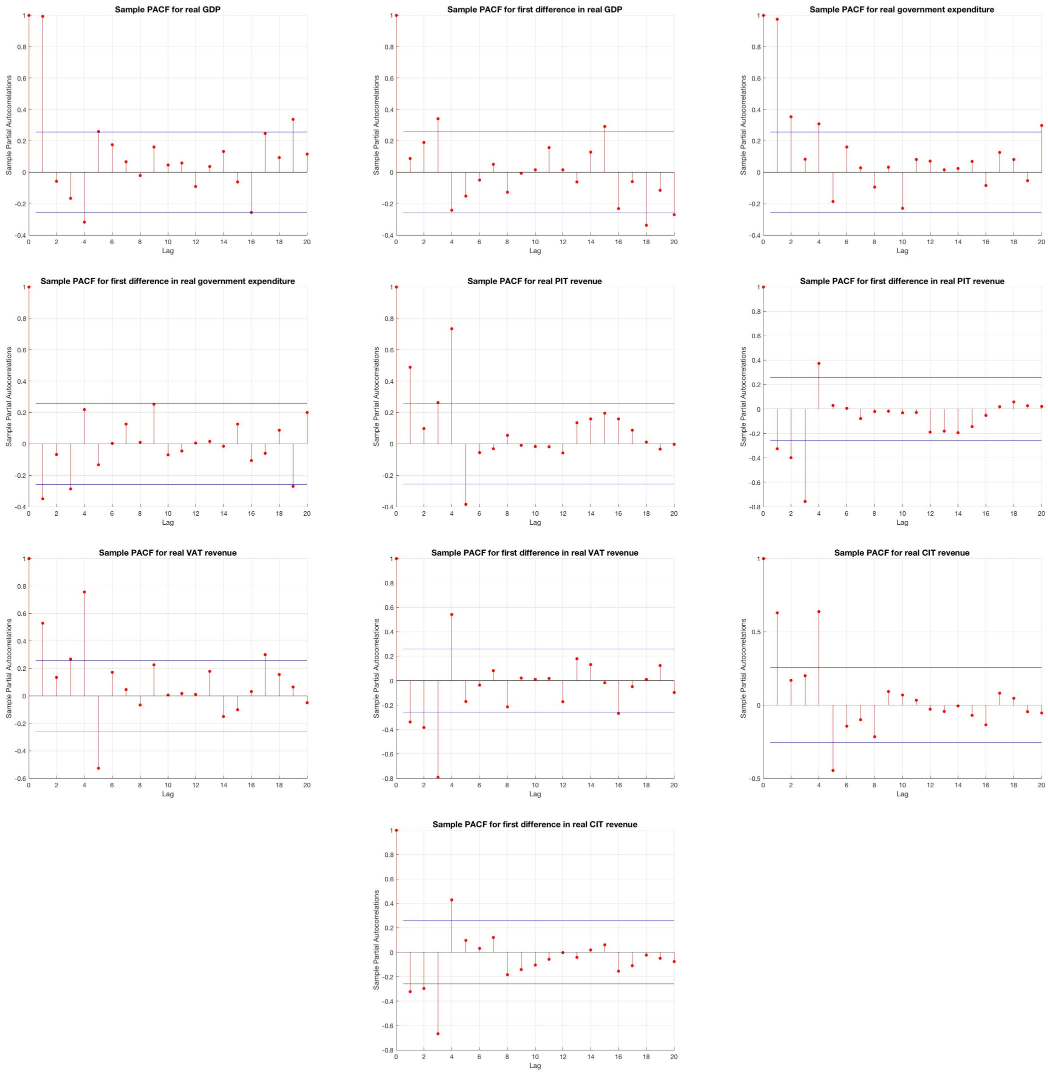
6 Conclusion

The evidence from structural impulse response functions, obtained from an estimated SVAR and an identification scheme based on theoretical and institutional characteristics of the Polish economy, suggests that VAT is the best tax instrument to be used to finance the public debt, without a significant adverse effect on output, and instead with a small, significant positive effect. In addition, the results show that PIT and CIT rate increases produce significant adverse effects on output, with PIT being the worst alternative due to presence of adverse consumption effects. The policy interpretation of these results should be cautious, as the positive effects of VAT increase on output only hold for small changes in the VAT rate, and may not be applicable for large changes, suggesting the possibility of nonlinear dynamics, which are not present in this model.

There are several possible extensions that could improve upon this sudy. The first one is the inclusion of a greater number of observations, providing more consistent estimates of model parameters and hence, the structural IRFs, which may undermine the positive effect of VAT shocks on output. This is particularly important, as the estimation of reduced form VAR with multiplicative dummies used up a lot of degrees of freedom, which as mentioned before questions the consistency of OLS estimators. Since transmission mechanisms of tax changes are a crucial aspect of disaggregated fiscal policy analysis, a microfounded model, such as DSGE, could be a useful extension through providing a mathematical description of underlying microeconomic mechanisms, which could more riguoruosly explain why VAT shocks result in a positive responses of output, instead of using outside research to explain the estimated results. Finally, the limitation of much SVAR fiscal policy research, and especially disaggregated tax SVAR approaches is the lack of monetary policy variables in the model, especially interest rate and inflation. As mentioned in the discussion of transmission mechanisms, increases in VAT vs. PIT or CIT have different effects on prices, where a rise in VAT over time has an upward pressure on prices, while PIT increase, for instance, raises a large possibility of downward pressure on prices resulting from adverse consumption effects. Through an inclusion of monetary variables in the model, it could be possible to understand monetary and fiscal policy interactions for each particular tax, where introducing monetary policy into the model could suggest that a different tax instrument is preferred, based on its impact on output.

7 Appendices

Appendix 1: Sample partial autocorrelation functions for real GDP, government expenditure and tax revenue series, including first differences (all in natural logs)



Appendix 2: ADF unit root test for the first difference in log real GDP (1) and Perron's version of the test for log real GDP growth with structural break in 2008 (2)

	(1) First difference in real GDP	(2) First difference in real GDP
1st lag coefficient	-0.483997	-0.6745458
	(-2.47)	(-3.03)
1st lag of difference	-0.5037049	-0.3727153
	(-2.88)	(-2.04)
2nd lag of difference	-0.3413119	-0.2789889
	(-2.62)	(-2.13)
Constant	0.0048032	0.0093621
	(2.35)	(3.05)
Level dummy		-0.0041707
		(-2.05)
Pulse dummy		-0.0016006
		(-0.23)
Test statistic	-2.470	-3.03
10% c.v.	-2.597	-2.597
Reject H_0 ?	No	Yes
Observations	58	58

t statistics in parentheses

H_0: unit root present in the series

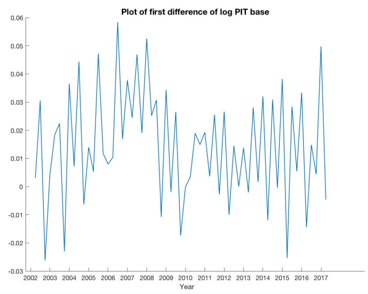
Appendix 3: ADF unit root tests for log real tax revenues and log real government expenditure

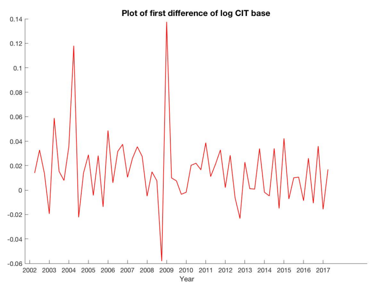
	(1) PIT revenue	(2) VAT revenue	(3) CIT revenue	(4) Government expenditure
1st lag coefficient	-0.3796328	-0.2135725	-0.2131756	-0.0698021
	(-2.30)	(-2.16)	(-1.89)	(-0.91)
1st lag of difference	-0.1417825	-0.1363233	-0.1548821	-0.3764104
	(-0.75)	(-1.00)	(-1.01)	(-2.67)
2nd lag of difference	-0.2318691	-0.1617526	-0.2010446	-0.1786006
	(-1.45)	(-1.34)	(-1.48)	(-1.22)
3rd lag of difference	-0.2916802	-0.1865711	-0.2848471	-0.2930685
	(-2.08)	(-1.73)	(-2.28)	(-2.19)
4th lag of difference	0.4307466	0.5596578	0.4645694	
	(3.42)	(5.67)	(3.88)	
Constant	3.574315	2.25336	1.933545	0.8208861
	(2.30)	(2.18)	(1.90)	(0.93)
Trend	0.0035931	0.0019459	0.0016099	0.0002905
	(2.10)	(1.78)	(1.22)	(0.46)
Test statistic	-2.301	-2.161	-1.889	-0.906
10% c.v.	-3.176	-3.176	-3.176	-3.176
Reject H_0 ?	No	No	No	No
Observations	57	57	57	58

t statistics in parentheses

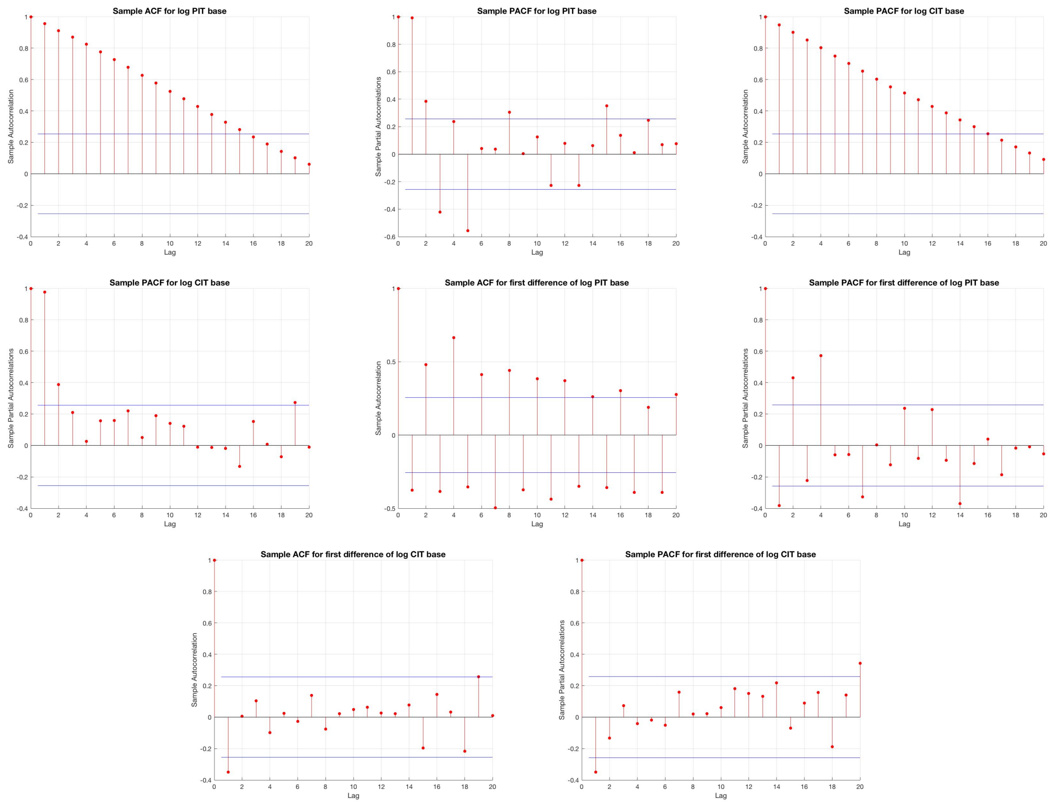
H_0: unit root present in the series

Appendix 4: Plots of first differences in log PIT and CIT bases





Appendix 5: ACF and PACF for log PIT and CIT bases and their first differences



Appendix 6: ADF unit root tests for log PIT and CIT bases

	(1) $ln(PITbase)$	(2) $ln(CITbase)$
1st lag coefficient	-0.0599875	-0.0801149
	(-1.39)	(-1.20)
1st lag of difference	-0.1861652	-0.3569397
	(-1.48)	(-2.82)
2nd lag of difference	0.4550898	
	(3.58)	
Constant	1.040769	0.9569794
	(1.41)	(1.24)
Trend	0.0008853	0.008583
	(1.28)	(0.77)
Test statistic	-1.387	-1.197
10% c.v.	-3.176	-3.176
Reject H_0 ?	No	No
Observations	59	60

t statistics in parentheses

H_0: unit root present in the series

Appendix 7: ADF tests for cointegration between PIT/CIT bases and output

Static regressions		
	(1)	(2)
Regression	PIT base on output (logs)	CIT base on output (logs)
Coefficient	1.010827	1.058809
	(81.67)	(91.65)
ADF tests of estimated residuals		
1st lag coefficient	-0.1937698	-0.5368908*
	(-2.51)	(-4.83)
1st lag of difference	-0.2356053	
	(-1.82)	
2nd lag of difference	0.396734	
	(3.21)	
Constant	-0.0004973	0.0006702
	(-0.26)	(0.24)
Test statistic	-2.512	-4.829
10% c.v.	-3.0462	-3.0462
Reject H_0 ?	No	Yes
Observations	59	61

t statistics in parentheses

H_0: unit root present in the series, implies no cointegration

*All lags of differences were insignificant, but H_0 was rejected in all cases

Appendix 8: Regressions of PIT base on output for elasticity calculations

	(1)	(2)	(3)
	$\Delta ln(PITbase)_t$	$\Delta ln(PITbase)_t$	$\Delta ln(PITbase)_t$
$\Delta ln(PITbase)_{t-1}$	-0.268	-0.422	
	(-2.88)	(-4.42)	
$\Delta ln(PITbase)_{t-2}$	0.413		
	(4.62)		
$\Delta ln(y)_t$	0.772	0.805	0.730*
	(3.71)	(3.57)	(3.20)
Constant	0.00119	0.00909	0.00380
	(0.30)	(2.28)	(0.88)
Serial correlation?	Yes	Yes	Yes
Observations	59	60	61

t statistics in parentheses

Breusch-Godfrey tests with 4 lags used to test for serial correlation

* No lagged dependent variables present, so the coefficient remains unbiased

Newey-West standard errors used to account for serial correlation

Appendix 9: Error correction model for CIT base on output regression

	(1)
	$\Delta ln(CITbase)_t$
$\Delta ln(y)_t$	1.568
	(5.76)
$\hat{\varepsilon}_{t-1}$	-0.531
	(-4.88)
Constant	-0.00682
	(-1.41)
Serial correlation?	No
Observations	61

t statistics in parentheses

Breusch-Godfrey tests with 4 lags used to test for serial correlation

Insignificant lags of $\Delta ln(CITbase)$ dropped from the ECM

Appendix 10: VAR model lag selection criteria

Lag	LR	AIC	HQIC	SBIC
0		-9.08536	-8.80492	-8.36202
1	462.06	-16.4436	-15.8126	-14.816*
2	31.466	-16.1126	-15.1311	-135809
3	52.042	-16.1491	-14.817	-12.7132
4	130.27	-17.5824	-15.8998	-13.2424
5	84.639	-18.201*	-16.1678*	-12.9568
6	40.571*	-18.0326	-15.6489	-11.8842

* denotes optimal lag

quarterly dummies included as exogenous

Appendix 11: Johansen's cointegration test with unrestricted constant

Rank	Parameters	LL	Eigenvalue	Trace statistic	5% c.v.
0	5	476.51147		100.8561	68.52
1	14	503.6219	0.58888	46.6352*	47.21
2	21	513.80649	0.28389	26.2660	29.68
3	26	520.99207	0.20990	11.8949	15.41
4	29	525.9472	0.14995	1.9846	3.76
5	30	526.93951	0.03201		

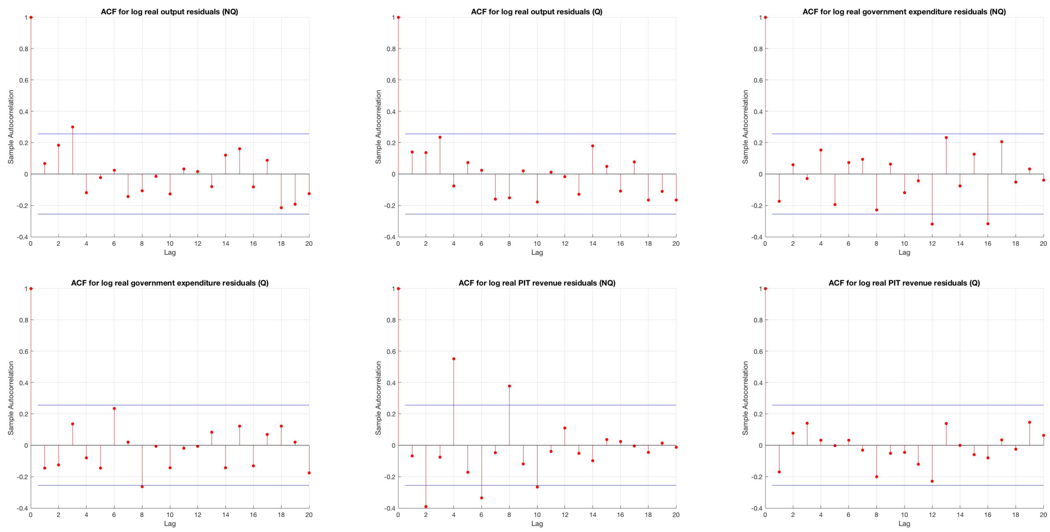
* denotes 1 cointegrating vector

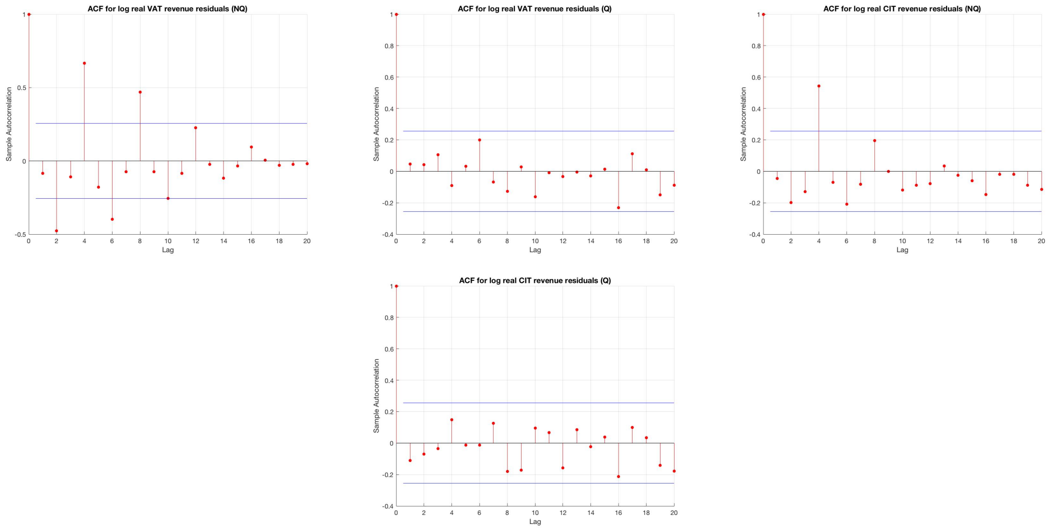
61 observations and one lag used in the specification

Restricted trend specification also gives 1 cointegrating vector

Appendix 12: Plots of reduced form residuals with (blue) vs. without multiplicative dummies (red)

Appendix 13: ACF for non quarter-dependent (NQ - additive dummies only) vs. quarter-dependent (Q - additive and multiplicative dummies) model residuals





Appendix 14: Brief summary of ADF unit root tests on non quarter-dependent vs. quarter-dependent model residuals

Residual	Non quarter-dependent (additive dummies only)		
	Lags	Test statistic	Reject H_0?
u_t^y	0	-7.165	Yes
u_t^g	0	-9.073	Yes
u_t^p	3	-2.591	No
u_t^v	3	-2.144	No
u_t^c	3	-2.351	No
Residual	**Quarter-dependent (additive and multiplicative dummies)**		
	Lags	Test statistic	Reject H_0?
u_t^y	0	-6.639	Yes
u_t^g	0	-8.804	Yes
u_t^p	0	-9.061	Yes
u_t^v	0	-7.270	Yes
u_t^c	0	-8.338	Yes

Optimal lags selected based upon PACF and lag significance in ADF regressions
H_0: unit root present
10% McKinnon critical value used: -2.596

Appendix 15: Estimated structural matrix of contemporaneous coefficients C_0 and estimated structural covariance matrix $\sum_\varepsilon$[6]

$$\mathbb{E}(\varepsilon_t\varepsilon_t') = \textstyle\sum_\varepsilon = \begin{pmatrix} 0.00003 & -0.00000 & 0.00000 & 0.00000 & -0.00000 \\ -0.00000 & 0.00024 & 0.00000 & 0.00000 & -0.00000 \\ 0.00000 & 0.00000 & 0.00173 & 0.00000 & -0.00000 \\ 0.00000 & 0.00000 & 0.00000 & 0.00081 & -0.00000 \\ -0.00000 & -0.00000 & -0.00000 & -0.00000 & 0.00342 \end{pmatrix}$$

$$C_0 = \begin{pmatrix} 1.0000 & -0.0201 & 0.0345 & -0.0298 & 0.0297 \\ 0.0000 & 1.0000 & 0.0000 & 0.0000 & 0.0000 \\ -1.4100 & 0.1535 & 1.0000 & -0.0000 & -0.0000 \\ -1.2249 & 0.4353 & 0.1595 & 1.0000 & -0.0000 \\ -2.1948 & 0.9349 & -0.6527 & -0.4850 & 1.0000 \end{pmatrix}$$

Appendix 16: Bootstrap confidence interval estimation procedure and Matlab code

1. Draw 61 samples with replacement from estimated reduced form residuals $u_t \equiv [u_t^y, u_t^g, u_t^p, u_t^v, u_t^c]'$:

```
uys=datasample(uy,61)
ugs=datasample(ug,61)
upits=datasample(upit,61)
uvats=datasample(uvat,61)
ucits=datasample(ucit,61)
E=[uys ugs upits uvats ucits]
```

2. Use each of the 61 samples of u_t, denoted by u_t^*, and the estimated reduced form coefficient D_1 to create a new data generating process for each variable in $Y_t^* \equiv [y_t^*, g_t^*, \tau_p^*, \tau_v^*, \tau_c^*]'$. The vector of first observations (from period 0), Y_0^* is equal to the first observations from original series of Y_t, while observations $Y_1^*, Y_2^*, ..., Y_T^*$ are generated as follows:

$$Y_1^* = D_1 Y_0^* + u_1^*$$
$$Y_2^* = D_1 Y_1^* + u_2^*$$
$$\vdots$$
$$Y_T^* = D_1 Y_{T-1}^* + u_T^*$$

```
% Constructing the NaN matrix to be filled
Ys=NaN(61,5)
% The rows can be accessed as:
Ys(1,:)
% Constructing the first two observations:
Ys(1,:)=Y(1,:)'
Ys(2,:)=D1*(M(1,:)')+E(2,:)'
% Bootstrap loop in general:
for i=3:61
    Ys(i,:)=D1*(Ys(i-1,:)')+E(i,:)'
end
```

[6]Covariances between structural shocks, which all show values of 0.00000 or -0.00000 were extremely small, usually to the powers ranging from -20 to -22. Note that in the C_0 matrix, a negative coefficient means that contemporaneous effect is positive - all signs are reversed.

3. Estimate the reduced form VAR based on boostrapped data generating process in step 2 and compute the structural impulse responses using matrix C_0^{-1} as outlined in Section 5.3. In this study, I compute 41 impulse response coefficients.

```
% Accessing each data generating process for VAR estimation
y2=Ys(:,1)
g2=Ys(:,2)
pit2=Ys(:,3)
vat2=Ys(:,4)
cit2=Ys(:,5)
% Estimating VAR based on resampled data generating processes
[EstMdl2,EstSE,logL,E]=estimate(Mdl,[y2 g2 pit2 vat2 cit2],'X',[q2 q3 q4])
a=cell(41,1)
D1s=EstMdl2.AR{1}
for p=1:41
    a{p}=(D1s^(p-1))*Ci
end
```

4. Repeat steps 1-3 2000 times, where each of the 2000 computed structural impulse response matrix series is saved as a column in a (1,2000) array in Matlab

5. Access the jk'th element of each matrix to obtain particular structural impulse response functions and for each IRF, create a matrix with 41 columns (one column for each structural IRF coefficient) and 2000 rows (for each coefficient, 2000 bootstrap values):

```
% Example: structural PIT shock on output
% 2000 bootstrap values for each coefficient
Byp1=NaN(2000,1)
for k=1:2000
    Byp1(k,1)=BR{1,k}{1,1}(1,3)
end
% BR is the array of all structural impulse response matrices produced in 2000
bootstrap replications
% This is the first column, for the first IRF coefficient
Byp(:,1)=Byp1
% Repeat this for Byp(:,2), Byp(:,3) and so on
```

6. Each structural IRF coefficient now has a distribution composed of 2000 observations, which appears bell-shaped. Take 2.5th and 97.5th percentile to calculate the 95% confidence interval, that is, upper and lower bounds for each structural IRF coefficient and then plot the upper and lower bounds together with the original structural IRF:

```
% Calculating 97.5 and 2.5 percentiles:
% Lower bound
Bypl1=NaN(41,1)
for l=1:41
    Bypl1(l,1)=prctile(Byp(:,l),2.5)
end
% Upper bound
Byph1=NaN(41,1)
for l=1:41
    Byph1(l,1)=prctile(Byp(:,l),97.5)
end
% Plotting the graph
hold on
plot(IRFpit)
plot(Bypl1)
plot(Byph1)
hold off
```

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