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What Drives Fiscal Multipliers? The Role of Private Wealth and Debt*

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JEL ref.: C22, E62, H30.

Keywords: multiplier effects; fiscal policy; asset markets; credit markets.

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What Drives Fiscal Multipliers? The Role of Private Wealth and Debt*

Sebastian Gechert† Rafael Mentges‡

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1 Introduction

The current literature on fiscal multipliers is trying to clear the "fiscal multiplier morass" (Leeper et al. 2011) that emerged amid the renewed interest in fiscal policy after the beginning of the financial crisis and recent attempts at consolidation in the US and

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Europe. The large bandwidth of fiscal multiplier estimations ranging from negative multipliers (implying expansionary consolidations) to large positive multipliers (implying self-defeating consolidations) is partly due to different methods of identification of fiscal shocks and the inclusion or omission of important variables (Gechert 2013).

Several identification schemes have been applied to resolve the issue of endogeneity regarding the business cycle in fiscal multiplier estimations, among them the use of the cyclically adjusted primary balance (CAPB) as a measure of exogenous discretionary fiscal policy decisions in event studies (Alesina and Ardagna 2010), as well as the recursive approach (Fatás and Mihov 2001) and the one by Blanchard and Perotti (2002) in structural vector autoregressive (SVAR) models. However, the adjustment regarding business cycle movements may not be enough in the presence of pronounced asset market movements that influence the budget and GDP over and above what is generally recognized as business cycle swings (Guajardo et al. 2011; Perotti 2011; Bornhorst et al. 2011).

The mechanism can be exemplified as follows: Consider an asset price boom that leads to higher revenues through capital gains and turnover taxation, unaccounted for by the usual elasticities of revenues and thus would falsely signal an improvement in the fiscal stance as measured by business cycle adjusted budget variables. If the asset price boom is followed by an increase in output, the positive correlation of the measure of the fiscal stance with output would be falsely deemed an example for expansionary consolidations. The very same argument holds for downturns of asset price cycles where the cyclically adjusted balance and GDP are likely to exhibit a coincidental deterioration, which could be misinterpreted as a causality running from public deficits to decreasing GDP. Both situations would lead to underestimations of average fiscal multipliers.

Our main contribution to the existing literature is the allowance for an impact of asset and credit market movements on the fiscal budget and GDP which is largely overlooked in the empirical literature on fiscal multipliers. In a first step we set up a formal framework to pin down the impact of the omission of these channels on estimated multiplier values; in a second step, we quantify the possible bias on multiplier estimations by employing established identification schemes, namely the CAPB and the SVAR approach and compare their results regarding multiplier effects in the case of inclusion vs. exclusion of private wealth and debt proxies. For the CAPB identification we use a recursive VAR and compare the results of a fiscal consolidation shock. For the structural VAR, we test the recursive identification (Fatás and Mihov 2001) as well as the Blanchard and Perotti (2002) approach in a standard SVAR based on Caldara and Kamps (2008) regarding their bias in multiplier estimations from government spending impulses. Our work is based on US quarterly data ranging from 1960:1 to 2012:4.

To the best of our knowledge, this is the first paper to quantify the potential downward bias that has been claimed by Guajardo et al. (2011) and Perotti (2011) within a VAR approach. As opposed to Yang et al. (2013), who address only the usual identification bias, we are thus able to allow for an additional omitted variable bias from movements in asset and credit markets on GDP, which could amplify the possible downward-bias on multiplier estimations; second, we can test the structural VAR identifications of Blanchard and Perotti (2002) and Fatás and Mihov (2001) for similar biases; third, we do not only look at episodes of consolidations but also at fiscal expansions; fourth, in addition to asset market movements, we allow for an influence from credit markets as they may alter the net wealth position and interfere with the influence of asset swings on the budget. What is more, we extend the formal framework of Perotti (2011) to show both the identification bias and the omitted variable bias that can occur in the presence of asset and credit market movements.

Our results confirm the hypothesis of Guajardo et al. (2011), who argue that the CAPB is a biased measure of the actual fiscal stance, and extends the findings in Yang et al. (2013), who show that with their asset price-adjusted CAPB measure, consolidations are contractionary. We find downward biased multipliers from identifications based on prior information regarding business cycle endogeneity, namely the CAPB and standard structural VAR approaches, as they overlook the influence of asset and credit market movements on GDP and the fiscal budget. Multipliers are on average about 0.3 to 0.6 units higher when taking this influence into account. These findings are robust to alternative specifications. Consolidations are thus more likely to be contractionary and could be more harmful to growth than expected from the results of some of the existing literature.

The remainder of the paper is structured as follows: Section 2 reviews the literature on the effects of fiscal policy. Section 3 explains the relation between fiscal multiplier estimations and asset and credit market variables, and their working through the wrong identification bias and the omission of these variables. Section 4 shows in detail the possible estimation biases within a formal framework. Section 5 and 6 contain an outline of the empirical strategy and a description of the data used in the estimations. In Section 7 we explain the structure and the identification methods used in the baseline models, while in Section 8 there is a discussion of the properties of the fiscal shocks in the baseline models. In Section 9 the same is done for the augmented model. Section 10 compares the effects of fiscal policy in the baseline model and in the augmented model, followed by several robustness checks in Section 11. The final section concludes.

2 Literature Review

In order to identify exogenous fiscal shocks and distinguish them from endogenous reactions, one strand of the literature relies on cyclically adjusted budget variables (Alesina and Ardagna 2010). The use of this approach has been criticized by Guajardo et al. (2011) and Perotti (2011) for insufficient identification in the presence of asset price movements that trigger a potential co-movement of GDP and cyclically adjusted budget variables leading to downward-biased multipliers. So far it has been overlooked that the same critique applies to another strand of the literature that adjusts budget variables by directly imposing restrictions from prior information on budget sensitivities and recognition and implementation lags to VAR estimations (Fatás and Mihov 2001; Blanchard and Perotti 2002).

The results from the literature on fiscal multiplier evaluations are wide-ranging, for an overview see Gechert (2013) and Mineshima et al. (2013). As a general finding, multipliers are on average positive with spending multipliers close to one and tax multipliers somewhat lower. Some strand of the literature finds a regime dependence, where multipliers are usually large in recession regimes (Auerbach and Gorodnichenko 2012). Another strand, that focuses on non-Keynesian effects of fiscal policy, and predominantly works with the CAPB identification, tends to find negative multipliers, i. e. expansionary consolidations (Alesina and Ardagna 2010; **?**; Giudice et al. 2007).

The recognition of shortcomings in the cyclical adjustment of the fiscal budget stems from a more specialized literature that deals with the sensitivity of the public budget to long swings in asset markets with a focus on questions of fiscal surveillance. Eschenbach and Schuknecht (2004) find that revenues are influenced by capital gains and turnover taxation as well as the impact of wealth effects on private demand and the revenues thereof. Public spending may increase when asset price busts call for bail-outs of private sector entities. They argue that a symmetric influence of swings of asset prices will balance out over the cycle and should not pose a problem to budget surveillance in the long run, but they point to possible asymmetries and inefficiencies when planning the budget based on this distorted information. We, however, posit that even in the case of symmetry, multiplier estimations may be biased since they rely on the short-run correlation of budget variables and GDP, which is usually given a causal interpretation running from the fiscal variable to GDP as long as the fiscal variable is cyclically adjusted, and thus deemed exogenous.

Congressional Budget Office (2013) point out that asset price movements are still unaccounted for in official cyclically adjusted data of the public balance in the US, which is the same for the EU Mourre et al. (2013). Morris and Schuknecht (2007) and Price and Dang (2011) estimate budget sensitivities to asset price cycles and calculate asset-adjusted structural balances for some OECD countries. Both papers find asset price cycles to be a major factor of unexplained movements in cyclically adjusted budgets.

Yang et al. (2013) try to improve the method of cyclical adjustment of the public budget of Alesina and Ardagna (2010) for some OECD countries by additionally regressing revenues on asset price movements and comparing their outcomes to the action-based approach of Guajardo et al. (2011). They find their corrected CAPB measure to produce significantly positive multipliers, close to those of Guajardo et al. (2011).

Bénétrix and Lane (2011) investigate the impact of private credit market fluctuations on fiscal balances. They find some evidence that credit growth has a positive influence on the public budget. Besides some indirect channels, where credit growth fuels asset prices and thus feeds the channels described above, they argue that higher private debt may be an indicator for demand shifts towards non-tradeable goods and services which could increase tax revenues. Moreover, they point to credit growth fueling inflation which could foster the fiscal drag and raise tax revenues as compared to real GDP.

Regarding the influence of asset and credit variables on output itself, Case et al. (2005) and Poterba (2000) find positive wealth effects through housing and stock markets. Lindner (2013) however finds the wealth effect for the US to be positive only after the mid 1980s and negative before.

The aforementioned literature focuses very much on wrong identifications of fiscal shocks. We argue that the problem of wrong identifications may be amplified by an omitted variable bias in estimations of fiscal multipliers, when movements in asset and credit markets cause changes to both the fiscal budget and aggregate demand. The literature on fiscal effects on output has discussed several omitted variable biases, such as the influences of international spillovers (Beetsma et al. 2006), the monetary policy reaction (Woodford 2011), the exchange rate regime (Corsetti et al. 2012), public debt (Chung and Leeper 2007; Favero and Giavazzi 2007), and liquidity or credit constraints in recessions (Eggertson and Krugman 2012). The latter are analyzed in empirical studies by distinguishing upper and lower regimes of the state of the economic cycle, and they usually find higher multipliers in recessions than in expansions (Auerbach and Gorodnichenko 2012; Fazzari et al. 2012; Ferraresi et al. 2013; Batini et al. 2012). Note that our approach is different from theirs, as we focus on a general downward bias that occurs both in the upswing and in the downswing of asset and credit markets.

3 Asset and Credit Markets and Fiscal Multiplier Estimations

We set up the hypothesis that credit market and asset market movements can have considerable effects on the estimated fiscal multiplier. We distinguish two effects relevant for the estimation of fiscal multipliers, namely, (i) the wrong identification of fiscal shocks and (ii) the omitted variable bias.

3.1 Identification Problem

As argued in the previous section, asset and credit market developments can significantly influence fiscal variables. Thus, changes in these fiscal variables due to movements in asset markets can be misinterpreted as changes in the fiscal stance if the time series that should represent the fiscal stance is contaminated by endogenous changes (Guajardo et al. 2011; Perotti 2011; Bornhorst et al. 2011). Though the literature focuses on asset market variables, credit market variables could play an important interfering role. Incorrect identifications could hold for both sides of the budget, i.e. tax revenues, and, to a lesser extent, government spending.[1]

Cyclically adjusted balances are commonly used as an estimate of the structural balance in order to gain information about the fiscal stance. However, asset price swings are usually not accounted for in the cyclical adjustment. Tax revenues, being an important part of the CAPB, are influenced mainly by capital gains and turnover taxation. Eschenbach and Schuknecht (2004) list several channels where personal and corporate income taxes are affected by asset market developments. They additionally point out that governments may also draw revenue from transaction taxes. The size of these effects depends on the tax system and can differ significantly between countries (Eschenbach and Schuknecht 2002; Girouard and Price 2004).[2] The fiscal effect also depends on the intensity of asset taxation, the dispersion of the ownership of assets, and the frequency of tax base adjustments to changing market values (Eschenbach and Schuknecht 2004).

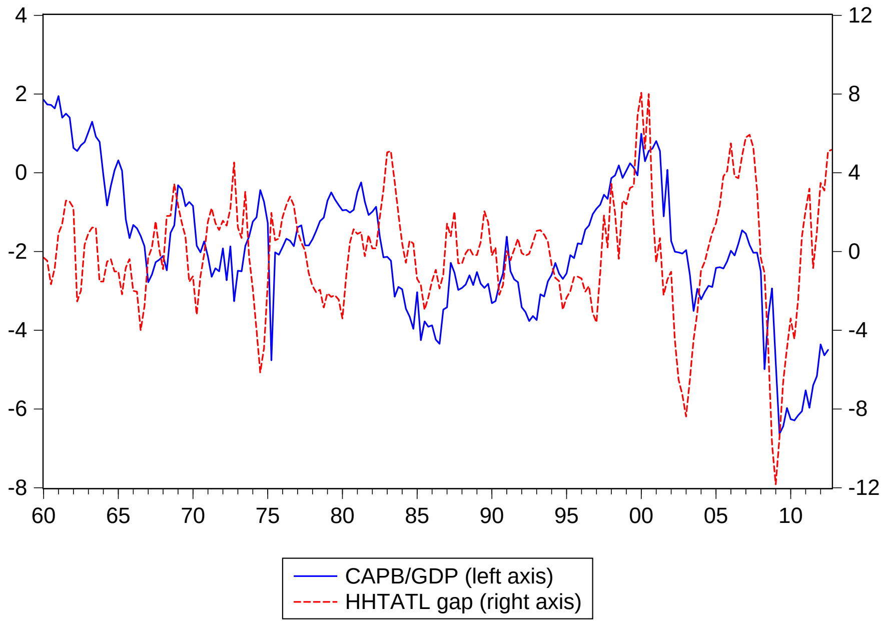
With respect to the influence of credit cycles on tax revenues, Bénétrix and Lane (2011) argue that credit growth fuels asset and property prices and increases tax revenues through this indirect channel. On the other hand, companies can save taxes by increasing their debt ratio (Miller 1977; Graham 2000). Furthermore, the United States are one of the few developed countries that allow home mortgage interest deduction.[3] Thus, an

[1]Throughout the paper tax revenues are defined as total tax revenues minus transfers (including interest payments). Government spending includes government consumption and government investment and is net of transfers.

[2]According to Girouard and Price (2004), the United States is one of the most affected countries in this regard.

[3]26 U.S.C. §163(h) of the internal revenue code.

Figure 1: CAPB-to-GDP Ratio Vs. De-trended Households' Total Assets-to-Total Liabilities Ratio



(Source: CBO, FRB Flow of Funds and authors' own calculations)

increase in overall debt can significantly alter tax revenues in one or the other direction and should be controlled for.

Government spending net of transfers seems to be less affected by asset price swings and credit market developments, but if there is a threat of large private-sector defaults due to collapsing asset prices or debt write-downs, fiscal authorities might be tempted to bail out parts of the private-sector. This is the case for 'too big to fail' banks or systemic risks to the whole financial system but also to certain industries that may be important for political reasons.

To capture the various channels described above, it seems reasonable to consider as a catch-all variable the ratio of wealth to debt, because budget variables should have a stronger correlation with the wealth-to-debt ratio than with any single asset or credit variable.

Figure 1 gives a first impression of the correlation between the de-trended wealth-to-debt ratio of households (households' total assets-to-total liabilities ratio) and the CAPB-to-GDP ratio. The most striking co-movements of these two variables are between 1997 and 2010. These two cycles represent the new economy boom and the housing bubble as well as their respective busts. Both were accompanied by a broad increase in stock prices. While the household wealth-to-debt ratio and the CAPB-to-GDP ratio before 1997 do not show such obvious co-movements as thereafter, it is nevertheless tenable to link them for a number of phases: 1985-1990, 1974-1976 and 1967-1970. Figure 1 also suggests that movements in the wealth-to-debt ratio hit the budget with some delay. This is plausible as most taxes are collected with a considerable lag.

3.2 Omitted Variable Bias

Our second argument for the inclusion of credit market and asset market measures to our estimations is the case of an omitted variable bias, which is supposed to influence estimations of government spending multipliers and tax multipliers alike. Theoretically, as shown below, the omitted variable bias could be neglected if a time series with correctly identified fiscal shocks were available and these would be uncorrelated with asset prices or credit markets. However, with imperfect identification, omission of these variables amplifies the wrong measurement of fiscal multipliers, because credit and asset market movements can have considerable effects on aggregate demand that would be captured by fiscal multiplier estimations without having anything to do with fiscal policies. Long term financial cycles influence GDP over and above what is generally recognized as business cycle swings (Borio 2012).

Asset prices affect consumer spending via wealth effects, an increase in confidence and credit-worthiness (Eschenbach and Schuknecht 2004). Poterba (2000) finds evidence that both the direct wealth effect and the indirect effect through confidence in future economic development help explain consumption during the observed period.

IMF (2012) reports evidence of recessions being more severe and more protracted if they were preceded by strong increases in household debt. According to their view, it is the particular combination of high leverage ratios and house price busts that seem to explain the severity of the economic downturn. Those findings hold for the case where changes in gross household debt imply little change in economy-wide net debt. Dynan (2012) finds evidence for a positive correlation between leverage and spending declines during the financial crisis. The difference in spending declines between highly leveraged households and their less-leveraged counterparts have even been above what would have been predicted by wealth effects alone. This means that households reduce consumption if they feel uncomfortably indebted. Thus, increasing government spending or reducing taxes could attenuate this reduction, but would show no positive effect in an empirical setting when private debt is not controlled for.

In general, economic developments that are caused or prolonged by private asset price busts and debt deleveraging may bias the measured impact of expansionary fiscal policy downwards if they are not controlled for. If households spend more when the value of their assets increases relative to their debt, and if they deleverage when their assets melt down while their debt remains high, a measure of these private sector dynamics, such as the wealth-to-debt ratio, should be taken into account in a fiscal multiplier estimation.

4 A Formal Framework

To phrase our arguments in a more formal way, we build on a simple static model by Perotti (2011), then developed to demonstrate differences between multiplier estimations using the CAPB approach and the narrative approach (Romer and Romer 2010). We extend it in order to suit our purpose. The model is based on two simplified equations, one for the budget surplus and a GDP equation:

$$\Delta s = \alpha_{sy}\Delta y + \alpha_{sf}\Delta f + \beta_{sy}\Delta y + \epsilon_s \tag{1}$$

$$\Delta y = \gamma_1\epsilon_s + \gamma_2\beta_{sy}\Delta y + \gamma_3\Delta f + \epsilon_y \tag{2}$$

The first equation shows the budget surplus as a share of GDP (s) and the variables that cause changes in the surplus, namely, the log of real GDP (y), the log of the household wealth-to-debt ratio (f), and exogenous discretionary changes to the budget by the policymaker (ϵ_s), which are independent of the cyclical position of the economy. The log of real GDP enters the equation twice due to the two different impacts it has on the GDP: α_{sy} stands for the automatic stabilizers and β_{sy} reflects endogenous discretionary (countercyclical) policy measures.

The second equation is a simplified GDP reaction function. GDP reacts to changes in the exogenous discretionary component of fiscal policy (ϵ_s), but also to its endogenous discretionary component ($\beta_{sy}\Delta y$). Furthermore, unlike Perotti, who models his financial market variable as white noise positively correlated with GDP, we give the wealth-to-debt ratio f a more pivotal role and model it as a non-stochastic variable.

We follow Blanchard and Perotti (2002) who argue that, due to recognition and implementation lags, $\beta_{sy} = 0$ when quarterly data are used. Therefore, we can drop the second term of equation 2:

$$\Delta y = \gamma_1\epsilon_s + \gamma_3\Delta f + \epsilon_y \tag{3}$$

In the following, we isolate the biases caused by the identification problem and the omission of the wealth-to-debt ratio.

4.1 Identification Bias

The measure of the cyclically adjusted surplus ($\Delta s^{capb} = \Delta s - \alpha_{sy}\Delta y$), when quarterly data are used ($\beta_{sy} = 0$), can be described as

$$\Delta s^{capb} = \alpha_{sf}\Delta f + \epsilon_s \tag{4}$$

which includes both the true exogenous shocks to the budget as well as the disturbances from f.

An OLS regression of Δy on Δs^{capb}, which should represent the impact of a fiscal contraction, gives the negative of the multiplier[4]

$$\gamma^{capb} = \frac{Cov(\Delta y, \Delta s^{capb})}{Var(\Delta s^{capb})} = \frac{\sum_i(\Delta s_i^{capb} - \overline{\Delta s^{capb}})\Delta y_i}{\sum_i(\Delta s_i^{capb} - \overline{\Delta s^{capb}})\Delta s_i^{capb}}. \tag{5}$$

Next we insert equation (4) in the numerator:

$$\gamma^{capb} = \frac{\sum_i(\alpha_f\Delta f + \epsilon_s - (\overline{\alpha_f\Delta f} + \overline{\epsilon_s}))\Delta y_i}{\sum_i(\Delta s_i^{capb} - \overline{\Delta s^{capb}})\Delta s_i^{capb}} \tag{6}$$

We rearrange this equation, isolating the true multiplier, γ_1, in order show the difference to the estimated multiplier, γ^{capb}.

$$\begin{aligned}\gamma^{capb} &= \frac{\sum_i(\alpha_f\Delta f - \overline{\alpha_f\Delta f})\Delta y_i + \sum_i(\epsilon_s - \overline{\epsilon_s})\Delta y_i}{\sum_i(\Delta s_i^{capb} - \overline{\Delta s^{capb}})\Delta s_i^{capb}} \\ &= \frac{Cov(\alpha_f\Delta f, \Delta y)}{\sum_i(\Delta s_i^{capb} - \overline{\Delta s^{capb}})\Delta s_i^{capb}} + \frac{Cov(\epsilon_s, \Delta y)}{\sum_i(\Delta s_i^{capb} - \overline{\Delta s^{capb}})\Delta s_i^{capb}}\end{aligned} \tag{7}$$

Using the information that the true negative value of the multiplier must be $\gamma_1 = Cov(\epsilon_s, \Delta y)/Var(\epsilon_s)$ we get

$$\gamma^{capb} = \gamma_1\frac{Var(\epsilon_s)}{Var(\Delta s^{capb})} + \frac{Cov(\alpha_f\Delta f, \Delta y)}{Var(\Delta s^{capb})} \tag{8}$$

[4] Remember that an increase of s^{capb} is supposed to represent a fiscal contraction. Fiscal multipliers are usually defined as the GDP reaction to a fiscal expansion. Thus, all γ presented in this section are effectively the negatives of the multiplier. Nevertheless, for convenience, we will refer to them as the multiplier.

or, after rearranging the denominator in the same way,

$$\gamma^{capb} = \gamma_1 \cdot \frac{Var(\epsilon_s)}{Var(\alpha_f \Delta f) + Var(\epsilon_s) + 2Cov(\alpha_f \Delta f, \epsilon_s)} + \frac{Cov(\alpha_f \Delta f, \Delta y)}{Var(\alpha_f \Delta f) + Var(\epsilon_s) + 2Cov(\alpha_f \Delta f, \epsilon_s)}. \quad (9)$$

Both terms show how the estimation of the multiplier with a cyclically adjusted fiscal budget is downward-biased in the presence of movements of f when it affects both GDP and the fiscal budget. The first term decreases the absolute value of the multiplier because the variance of ϵ_s is likely to be smaller than the variance of Δs^{capb}. The second term must be positive and therefore increases the estimated negative value of the multiplier γ^{capb}. Keep in mind that a positive change of ϵ_s marks an improvement in the fiscal stance (fiscal consolidation) and that adding the positive value of the second term thus downward-biases the estimated value of the multiplier, which is usually defined as the GDP reaction to a fiscal expansion.

4.2 Omitted Variable Bias

The model can also be used to isolate the omitted variable bias explained in Section 3. Ignoring, at least for a while, the identification problem, an OLS regression of GDP on the budget surplus would give the following coefficient:

$$\hat{\gamma_1} = \frac{Cov(\Delta s, \Delta y)}{Var(\Delta s)} \quad (10)$$

Now, let us assume that $\Delta y = \gamma_1 \epsilon_s + \gamma_3 \Delta f + \epsilon_y$ is the population model, but we estimate $\Delta y = \gamma_1 \Delta s_i + \epsilon_y$. Then the estimated value of γ_1 is

$$\hat{\gamma_1} = \frac{\sum_i (\Delta s_i - \overline{\Delta s}) \Delta y_i}{\sum_i (\Delta s_i - \overline{\Delta s}) \Delta s_i} = \frac{\sum_i (\Delta s_i - \overline{\Delta s})(\gamma_1 \Delta s_i + \gamma_3 \Delta f_i + \epsilon_{y,i})}{\sum_i (\Delta s_i - \overline{\Delta s}) \Delta s_i} \quad (11)$$

It can easily be shown that this term is equal to

$$\hat{\gamma_1} = \gamma_1 + \gamma_3 \frac{\sum_i (\Delta s_i - \overline{\Delta s}) \Delta f_i}{\sum_i (\Delta s_i - \overline{\Delta s}) \Delta s_i} + \frac{\sum_i (\Delta s_i - \overline{\Delta s}) \epsilon_{y,i}}{\sum_i (\Delta s_i - \overline{\Delta s}) \Delta s_i} \quad (12)$$

While the last term should be equal to zero in a well-specified structural VAR (meaning no simultaneous equation bias), the second term can only be zero in the case of a correctly identified fiscal shock which is unrelated to f. However, in the case of a correlation between the budget surplus and the wealth-to-debt ratio, there is a bias in the expected value of $\hat{\gamma}_1$; a positive correlation implies a positive difference of $\hat{\gamma}_1 - \gamma_1$, i. e. a downward biased multiplier.

5 Empirical Strategy

In order to test our hypothesis we follow a three-step approach. First, in Section 7, we set up three baseline VAR models of standard identification approaches, namely, using the CAPB as a measure of exogenous fiscal shocks and applying two variants of the SVAR methodology that impose restrictions to derive exogenous changes of budgetary decisions within the estimation.

The CAPB is tested in a four-variable VAR model, including the CAPB-to-GDP ratio, GDP, the GDP deflator and the short term real effective federal funds rate, identified by recursive ordering. Usually in the literature GDP effects of CAPB shocks are tested in an OLS framework, defining episodes of fiscal consolidations, with the CAPB seen as the fiscal stance (Alesina and Ardagna 2010). We, however, opt for a recursive VAR approach in order to provide a single coherent framework for all our tests, and to account for both an identification as well as an omitted variable bias; moreover, with the recursive VAR, we only impose contemporaneous exogeneity of the CAPB variable within the same quarter, exploiting recognition and implementation lags, while allowing for endogenous discretionary and automatic movements thereafter. With this strategy, we can disentangle the possible misidentification bias coming from endogenous discretionary reactions of policymakers to the business cycle from the one that is central to our study, namely, the endogeneity of cyclically adjusted budget variables to movements in asset and credit markets.

The baseline for the SVAR methodology is a five-variable VAR model of government spending net of transfers, GDP, the GDP deflator, tax revenues net of transfers and the short-term real effective federal funds rate, akin to the standard model in Caldara and Kamps (2008). We identify the SVAR both via a recursive approach (Fatás and Mihov 2001) and via the Blanchard-Perotti method (Blanchard and Perotti 2002).

In a second step in Section 8, the structural shocks derived from these three baseline models are tested for their orthogonality with respect to households' wealth-to-debt ratio (total assets to total liabilities). Correlation of these shocks with the wealth-to-debt ratio points to identification and omitted variable biases in our baseline models.

Thus, in a third step in Section 10, we augment our baseline VAR models with the wealth-to-debt ratio as an additional endogenous variable and compare the fiscal multipliers derived from these augmented models to their baseline counterparts. Given our hypothesis, we would expect increased multipliers from the augmented models.

Following the argumentation in Section 3 and our results obtained in Section 8 it seems straightforward that both asset and credit market movements, need to be recognized in a well-specified empirical model. In order to economize on degrees of freedom, we decided to use a single variable, the wealth-to-debt ratio, that reflects both sides of the markets. A possible downside is that we lose additional information which may be relevant for our estimation, because the choice of the ratio over including both variables acts as a restriction on the effects of both variables. This is why we run extended models, including both asset and credit variables separately, in Section 11.

6 Data

Estimations are based on US quarterly data from 1960:1 to 2012:4 and subsamples. Population, government budget series and GDP with its components stem from BEA tables. The GDP deflator, the effective federal funds rate, stock market and credit market data are taken from the FRED data base. Households' total assets and liabilities are provided by the Flow of Funds data of the FRB.

The series for CAPB-to-GDP ratio, which should represent the structural budget balance (s), stems from the Congressional Budget Office and already ends at 2012:3. Inflation (p) is the annualized growth rate of the GDP deflator; the real effective federal funds rate (r) is deflated by p.

Nominal volumes are deflated by the GDP deflator and expressed in per capita terms, transformed to logs and multiplied by 100 to scale them in line with the variables in percentages. We thus have the log of real per capita government current spending net of transfers (g), the log of real per capita revenues net of transfers (τ), the log of real GDP per capita (y), and the log of households' total-assets-to-liabilities ratio (f). For robustness tests on the financial market variable we construct the log of households' real per capita total assets and total liabilities separately, as well as the log of the deflated S&P 500 index and the log of real per capita non-financial private sector debt. Series are seasonally adjusted by the original sources or by the X12 procedure implemented in Eviews.

All variables included have been tested for a unit root by the augmented Dickey-Fuller test and have been found to be I(1) at the 5 percent critical level. Johansen tests in Table 5 in Appendix B by and large show cointegration with a rank of one for most specifications, however, test results become more valid for the augmented models including the wealth-to-debt ratio. Cointegration makes it feasible to apply a classic

VAR approach to non-stationary data as has been shown by Phillips and Durlauf (1986); West (1988); Fanchon and Wendel (1992) for example. Sims et al. (1990) argue that non-stationarity even without cointegration does not pose a problem to consistency of the estimators, notwithstanding a possible loss in efficiency for small samples.

7 Structure and Identification of the Baseline Models

We basically follow the terminology of the AB-model in Lütkepohl (2006: 364) to specify the structural shocks. The structural form of the VAR model can be expressed as

$$\boldsymbol{A}\boldsymbol{\Gamma}(L)X_t = \boldsymbol{A}v + \boldsymbol{B}\varepsilon_t \tag{13}$$

with X_t being the K-dimensional vector of endogenous variables and v representing the vector of exogenous variables, namely, a constant and a linear time trend. $\boldsymbol{\Gamma}(L)$ is a 4th-order lag polynomial of the $K \times K$ matrix $\boldsymbol{\Gamma}$, containing the coefficients of the endogenous variables and their lags.[5] ε_t is a K-dimensional vector of structural form disturbances (exogenous shocks). $\boldsymbol{A}$ and $\boldsymbol{B}$ are $K \times K$ factorization matrices and contain the contemporaneous dependencies among the endogenous variables and the structural shocks respectively. To give them an economic meaning for our application, $\boldsymbol{A}$ carries the automatic responses of the variables to shocks in the other variables, such as the sensitivity of taxes to changes in GDP, while $\boldsymbol{B}$ contains the discretionary reactions to innovations in the endogenous variables.

A formal derivation of the identification of the structural model from the reduced-form VAR and of the impulse-response functions (IRF) can be found in Appendix A.

7.1 CAPB Identification

Let us now turn to the setting of restrictions on $\boldsymbol{A}$ and $\boldsymbol{B}$ for our specific VAR models. In general, restrictions are set from prior economic information on elasticities, assumptions on institutional settings and recognition, implementation or response lags.

To measure the effects of fiscal policy changes with the CAPB in our baseline setting, we set up a four-variable VAR as in (13) with a lag order of four and the vector of endogenous variables

$$X_t = \begin{bmatrix} s_t & y_t & p_t & r_t \end{bmatrix}'. \tag{14}$$

[5] $\boldsymbol{\Gamma}(L)$ needs to be invertible for the VAR to be stable. That is, the coefficient matrices of $\boldsymbol{\Gamma}(L)$ must be absolutely summable. In other words, the coefficients of higher order of $\boldsymbol{\Gamma}(L)$ must converge to zero (Lütkepohl 2006: 27).

For identification of the CAPB-VAR we follow a simple Choleski decomposition, with the variables ordered as in (14) for the following reasons: The CAPB-to-GDP ratio is ordered first since it is taken to represent structural changes in fiscal policy stripped of automatic endogenous reactions to the other variables. Moreover, as argued in Fatás and Mihov (2001), due to recognition and implementation lags, discretionary fiscal policy should not respond to developments in other economic variables within the same quarter and should thus be contemporaneously exogenous, i. e. ordered first. Interest rate changes are ordered last since they are deemed not to provoke immediate changes in the other variables due to response lags, but could react to changes in other variables immediately. With regards to the two other variables, we follow the literature and order inflation after GDP; however, results are robust to a reversed ordering of the two variables.

With the recursive ordering, $\boldsymbol{A}$ becomes a lower triangular matrix with unit entries on the main diagonal. All entries above the diagonal are set to zero to reflect that no contemporaneous influence among the variables is assumed in this direction. Contemporaneous influences in the opposite direction are reflected by the α_{ij} items which can now be estimated. Note that no restrictions are set on the lagged interdependencies of the variables such that, e. g. inflation may influence GDP with a lag of one quarter. The $\boldsymbol{B}$ matrix reduces to a simple identity matrix for the Choleski decomposition. Thus, for our baseline specification we have

$$\boldsymbol{A} = \begin{bmatrix} 1 & 0 & 0 & 0 \\ -\alpha_{ys} & 1 & 0 & 0 \\ -\alpha_{ps} & -\alpha_{py} & 1 & 0 \\ -\alpha_{rs} & -\alpha_{ry} & -\alpha_{rp} & 1 \end{bmatrix} \qquad \boldsymbol{B} = \begin{bmatrix} 1 & 0 & 0 & 0 \\ 0 & 1 & 0 & 0 \\ 0 & 0 & 1 & 0 \\ 0 & 0 & 0 & 1 \end{bmatrix}. \tag{15}$$

After solving, the structural shocks for the time series of the four variables can be derived. The properties of the structural shocks will be analyzed in Section 8, after setting up the structure and identifying restrictions for the other two VAR models. We will also leave impulse-response analysis until Section 10, where we directly compare the baseline and augmented models and derive fiscal multipliers.

7.2 Structural VAR Identification – Recursive Approach

Instead of relying on cyclically adjusted budget variables to identify exogenous changes in the fiscal stance, the literature has developed alternative models that impose prior information on budget sensitivities directly to the estimation of the structural VAR model. With such a model, one can evaluate fiscal multipliers of spending and revenue components. The baseline specification is a five-variable fourth-order structural VAR model as in (13) with

$$X_t = \begin{bmatrix} g_t & y_t & p_t & \tau_t & r_t \end{bmatrix}'. \tag{16}$$

With respect to identification, the first approach that we will deal with is the recursive approach (RA) as applied by Fatás and Mihov (2001). It, again, uses the principle of contemporaneous one-way causality that is imposed by a Choleski ordering. We order the variables as they appear in (16).

The reasoning behind this ordering is close to that of the CAPB VAR. Due to recognition and implementation lags, the discretionary part of government spending net of transfers should not respond to developments in other economic variables within the same quarter. Moreover, government spending net of transfers is deemed insensitive to business cycle fluctuations. In line with Caldara and Kamps (2008), the tax variable, which is tax revenues net of transfers, is ordered after GDP and after inflation to capture its sensitivity to the business cycle. Note that this ordering implicitly assumes that there is no contemporaneous impact of taxes on GDP and inflation, which is questionable; however, the dilemma cannot be solved sufficiently within the recursive approach because, if taxes were ordered prior, one would implicitly assume a contemporaneous output and price elasticity of zero. The other variables are ordered as in the previous section. Under these assumptions, the factorization matrices become

$$\boldsymbol{A} = \begin{bmatrix} 1 & 0 & 0 & 0 & 0 \\ -\alpha_{yg} & 1 & 0 & 0 & 0 \\ -\alpha_{pg} & -\alpha_{py} & 1 & 0 & 0 \\ -\alpha_{\tau g} & -\alpha_{\tau y} & -\alpha_{\tau p} & 1 & 0 \\ -\alpha_{rg} & -\alpha_{ry} & -\alpha_{rp} & -\alpha_{r\tau} & 1 \end{bmatrix} \qquad \boldsymbol{B} = \begin{bmatrix} 1 & 0 & 0 & 0 & 0 \\ 0 & 1 & 0 & 0 & 0 \\ 0 & 0 & 1 & 0 & 0 \\ 0 & 0 & 0 & 1 & 0 \\ 0 & 0 & 0 & 0 & 1 \end{bmatrix}. \tag{17}$$

After solving by a Choleski decomposition, the structural shocks can be retrieved for later use in Section 8.

7.3 Structural VAR Identification – Blanchard-Perotti Approach

We now turn to the second standard identification approach in the SVAR literature, the Blanchard-Perotti (BP) approach (Blanchard and Perotti 2002). Equation (16) is used again to specify a five-variable fourth-order structural VAR, but, in line with Caldara and Kamps (2008), we restrict the factorization matrices of the baseline specification as follows.

$$\boldsymbol{A} = \begin{bmatrix} 1 & 0 & -\alpha_{gp} & 0 & 0 \\ -\alpha_{yg} & 1 & 0 & -\alpha_{y\tau} & 0 \\ -\alpha_{pg} & -\alpha_{py} & 1 & -\alpha_{p\tau} & 0 \\ 0 & -\alpha_{\tau y} & -\alpha_{\tau p} & 1 & 0 \\ -\alpha_{rg} & -\alpha_{ry} & -\alpha_{rp} & -\alpha_{r\tau} & 1 \end{bmatrix} \qquad \boldsymbol{B} = \begin{bmatrix} 1 & 0 & 0 & 0 & 0 \\ 0 & 1 & 0 & 0 & 0 \\ 0 & 0 & 1 & 0 & 0 \\ \beta_{\tau g} & 0 & 0 & 1 & 0 \\ 0 & 0 & 0 & 0 & 1 \end{bmatrix} \tag{18}$$

In contrast to the recursive approach, the BP approach uses additional prior assumptions on budget elasticities of tax revenues and institutional settings for identification. Leaving $\beta_{\tau g}$ unrestricted and setting $\beta_{g\tau} = 0$ implies that in the process of setting up the public budget, spending decisions are taken prior to revenue decisions, an assumption which has been shown to be robust by (Blanchard and Perotti 2002). For reasons of comparison, we follow Caldara and Kamps (2008), who draw on Perotti (2005) in setting the output and price elasticities of government spending and revenues for the full sample such that $\alpha_{\tau y} = 1.85, \alpha_{\tau p} = 1.25$ and $\alpha_{gp} = -.5$.[6] Imposing these restrictions, assuming that they are correct, has the advantage that we can leave the contemporaneous reaction of GDP and inflation to changes in net taxes unrestricted and have them determined by the data. After solving this third model, we can now analyse the properties of the structural shocks of all three models.

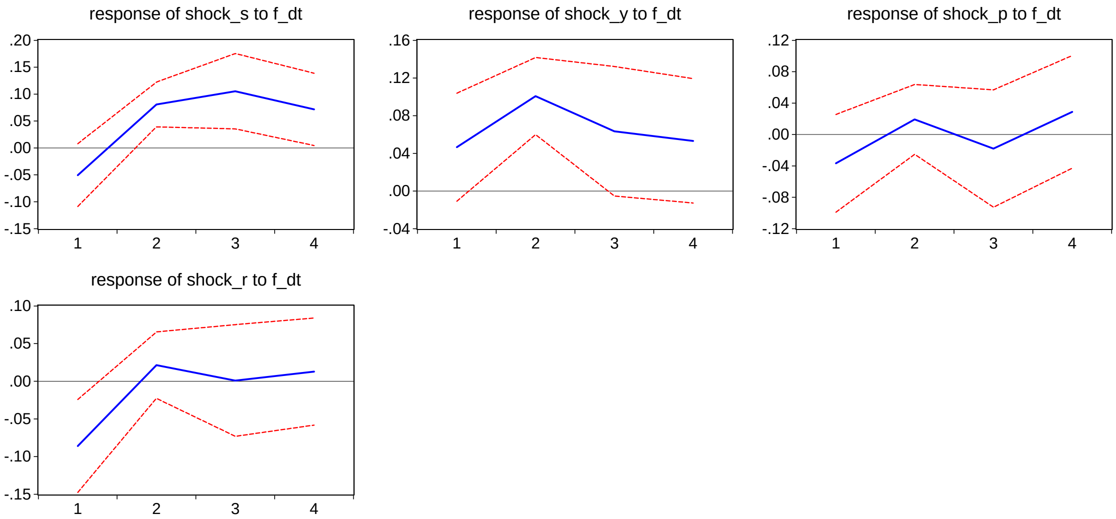
8 Properties of the Baseline Structural Shocks

If the specification of the baseline models is correct, the structural shocks derived from these models should be independent of other influences and should represent truly exogenous changes in the fiscal stance. However, our hypothesis is that private wealth and debt changes have an influence on the public budgetary position and on GDP over and above the usual business cycle fluctuations. We test this hypothesis for each of the three models against the null of no influence for the vector of shocks ε_t via the dynamic OLS model

$$\varepsilon_t = \alpha \sum_{i=0}^{4} fdetr_{t-i}\Gamma_{t-i} + e_t \tag{19}$$

[6]Perotti (2005) argues that the government's nominal wage bill does not instantaneously react to inflation, which is why he assumes that real wage payments, representing a large share of government spending, decrease with a shock to inflation.

Figure 2: Influence of wealth-to-debt ratio on Structural Residuals of CAPB VAR – Baseline



with $fdetr$ being the de-trended households' total assets over total liabilities ratio, using the same lag structure as for the VAR models. Impulse-responses are reported in Figures 2, 3 and 4.

In line with our theoretical reasoning, in Figure 2 the structural shocks derived for the CAPB-to-GDP ratio and GDP show a significantly positive correlation with changes in households' wealth-to-debt ratio. That is, an increase in the wealth-to-debt ratio comes with an increase of the budgetary position and GDP over and above the usual business cycle fluctuations.

Figures 3 and 4 show that the shocks derived from the RA and BP baseline VAR models co-move with the wealth-to-debt ratio of households, with government spending being negatively correlated and GDP and taxes being positively correlated to the wealth-to-debt ratio. Again, these results fit the arguments developed in Section 3. These results justify the augmenting of the baseline VAR models with the wealth-to-debt ratio as an additional endogenous variable, which will be done in the next section.

9 Structure and Identification of the Augmented Models

We augment our four-variable VAR model, which was set up to estimate fiscal multipliers from a change in the CAPB, and the two five-variable VAR models, which were built to estimate multipliers of government spending, by an additional endogenous variable,

Figure 3: Influence of wealth-to-debt ratio on Structural Residuals of RA VAR – Baseline

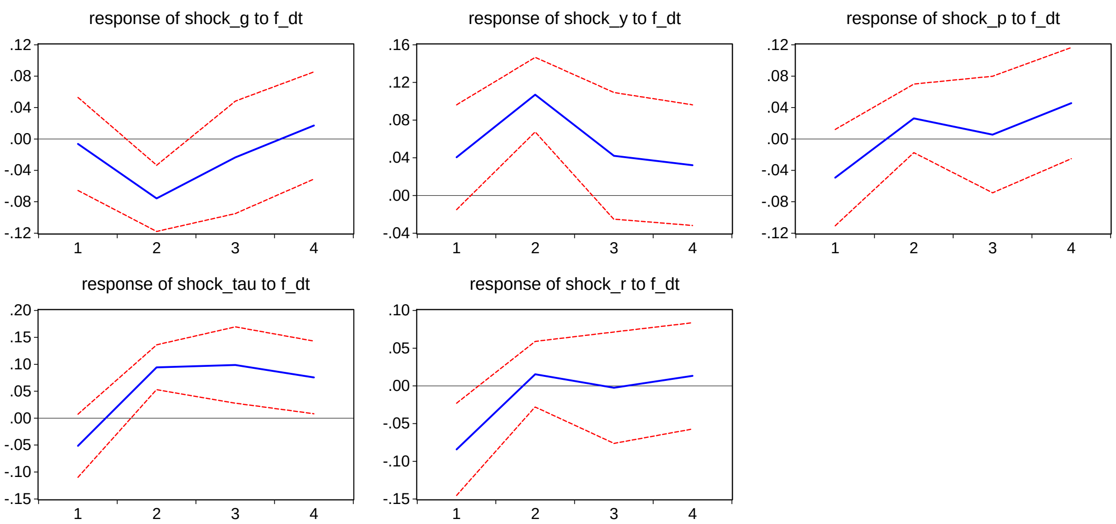
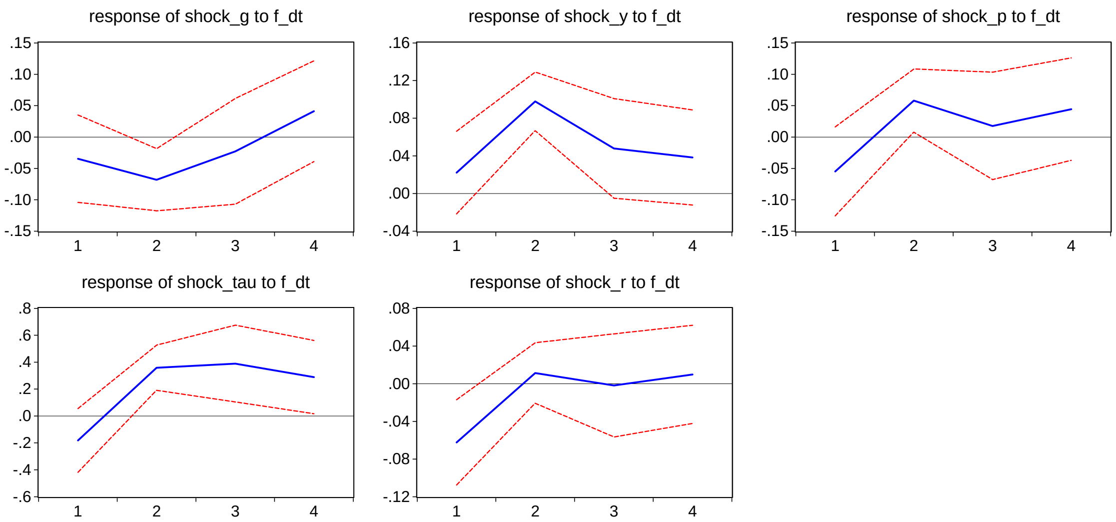


Figure 4: Influence of wealth-to-debt ratio on Structural Residuals of BP VAR – Baseline



which is the log of the wealth-to-debt ratio of households (households total assets over total liabilities, f). Since we do not want to rule out a contemporaneous dependency of households' leverage ratio on income and interest rates, and because we expect that the channels of influence from private wealth and debt on the public budget and GDP take some time to materialize, we order f last in the VAR. Results are, however, robust to ordering f first, as will be shown in Section 11. So far, for the CAPB VAR, we have

$$X_t^a = \begin{bmatrix} s_t & y_t & p_t & r_t & f_t \end{bmatrix}' \tag{20}$$

and the factorization reads

$$\boldsymbol{A} = \begin{bmatrix} 1 & 0 & 0 & 0 & 0 \\ -\alpha_{ys} & 1 & 0 & 0 & 0 \\ -\alpha_{ps} & -\alpha_{py} & 1 & 0 & 0 \\ -\alpha_{rs} & -\alpha_{ry} & -\alpha_{rp} & 1 & 0 \\ -\alpha_{fs} & -\alpha_{fy} & -\alpha_{fp} & -\alpha_{fr} & 1 \end{bmatrix} \qquad \boldsymbol{B} = \begin{bmatrix} 1 & 0 & 0 & 0 & 0 \\ 0 & 1 & 0 & 0 & 0 \\ 0 & 0 & 1 & 0 & 0 \\ 0 & 0 & 0 & 1 & 0 \\ 0 & 0 & 0 & 0 & 1 \end{bmatrix}. \tag{21}$$

While including the additional variable does not make the CAPB a better estimate of the fiscal stance per se, one has to keep in mind that we use it in a structural VAR, with the inclusion of the wealth-to-debt ratio working as an additional filter, whereby the identified fiscal shocks are more likely to be exogenous.

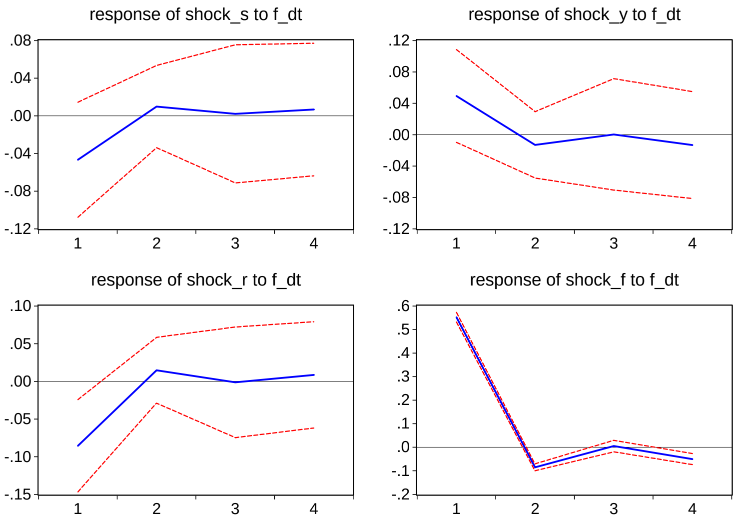
With respect to the augmented VAR models of the RA and the BP type, the vector of endogenous variables now is

$$X_t^a = \begin{bmatrix} g_t & y_t & p_t & \tau_t & r_t & f_t \end{bmatrix}' \tag{22}$$

with factorization of the RA model

$$\boldsymbol{A} = \begin{bmatrix} 1 & 0 & 0 & 0 & 0 & 0 \\ -\alpha_{yg} & 1 & 0 & 0 & 0 & 0 \\ -\alpha_{pg} & -\alpha_{py} & 1 & 0 & 0 & 0 \\ -\alpha_{\tau g} & -\alpha_{\tau y} & -\alpha_{\tau p} & 1 & 0 & 0 \\ -\alpha_{rg} & -\alpha_{ry} & -\alpha_{rp} & -\alpha_{r\tau} & 1 & 0 \\ -\alpha_{fg} & -\alpha_{fy} & -\alpha_{fp} & -\alpha_{f\tau} & -\alpha_{fr} & 1 \end{bmatrix} \qquad \boldsymbol{B} = \begin{bmatrix} 1 & 0 & 0 & 0 & 0 & 0 \\ 0 & 1 & 0 & 0 & 0 & 0 \\ 0 & 0 & 1 & 0 & 0 & 0 \\ 0 & 0 & 0 & 1 & 0 & 0 \\ 0 & 0 & 0 & 0 & 1 & 0 \\ 0 & 0 & 0 & 0 & 0 & 1 \end{bmatrix} \tag{23}$$

Figure 5: Influence of wealth-to-debt ratio on Structural Residuals of CAPB VAR – Augmented



and factorization of the BP model

$$\boldsymbol{A} = \begin{bmatrix} 1 & 0 & -\alpha_{gp} & 0 & 0 & -\alpha_{gf} \\ -\alpha_{yg} & 1 & 0 & -\alpha_{y\tau} & 0 & -\alpha_{yf} \\ -\alpha_{pg} & -\alpha_{py} & 1 & -\alpha_{p\tau} & 0 & -\alpha_{pf} \\ 0 & -\alpha_{\tau y} & -\alpha_{\tau p} & 1 & 0 & -\alpha_{\tau f} \\ -\alpha_{rg} & -\alpha_{ry} & -\alpha_{rp} & -\alpha_{r\tau} & 1 & -\alpha_{rf} \\ -\alpha_{fg} & -\alpha_{fy} & -\alpha_{fp} & -\alpha_{f\tau} & -\alpha_{fr} & 1 \end{bmatrix} \qquad \boldsymbol{B} = \begin{bmatrix} 1 & 0 & 0 & 0 & 0 & 0 \\ 0 & 1 & 0 & 0 & 0 & 0 \\ 0 & 0 & 1 & 0 & 0 & 0 \\ \beta_{\tau g} & 0 & 0 & 1 & 0 & 0 \\ 0 & 0 & 0 & 0 & 1 & 0 \\ 0 & 0 & 0 & 0 & 0 & 1 \end{bmatrix}. \tag{24}$$

Note that for the BP model we do not restrict the elasticities of the five baseline variables to the wealth-to-debt ratio with zeros, but impose the average elasticities derived from model (19), reported in Table 4.

After solving the three augmented models, we again retrieve the structural shocks and repeat the exercise of (19) to check whether the structural shocks are correlated with the wealth-to-debt ratio. Figures 5, 6 and 7 show the constructed impulse-responses. As can be seen, the structural shocks are now largely orthogonal to our additional variable, except, of course, for those of the wealth-to-debt ratio f itself.

We can now move on to a comparative impulse-response analysis of our baseline and augmented models.

Figure 6: Influence of wealth-to-debt ratio on Structural Residuals of RA VAR – Augmented

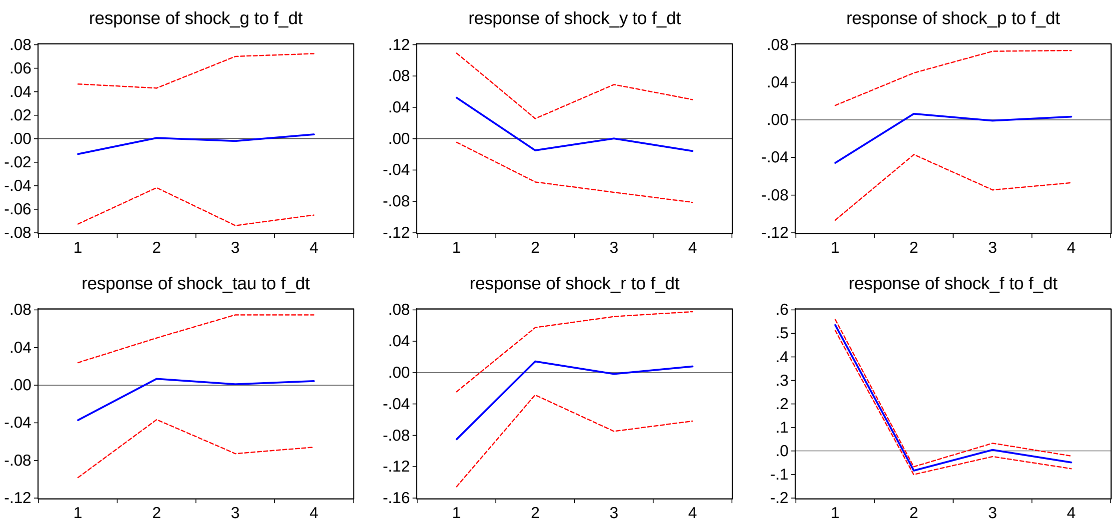


Figure 7: Influence of wealth-to-debt ratio on Structural Residuals of BP VAR – Augmented

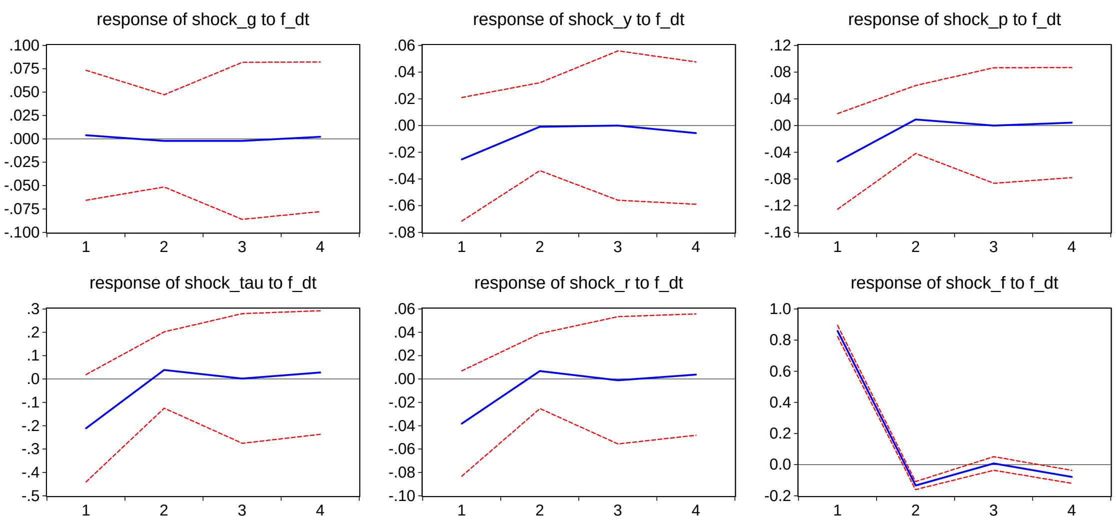


Table 1: Multipliers for Baseline and Augmented Models – Full Sample (1960-2012)

Model	f	Impact	Cumulative			Peak	
		Quarter 1	10	20	30	(Quarter)	
CAPB base.		0.14	0.40	0.45	0.36	0.31	(6)
CAPB augm.	HHTATL	0.31	1.17	1.64	1.91	0.63	(4)
RA base.		0.86	0.52	0.33	0.12	0.94	(3)
RA augm.	HHTATL	1.02	0.92	0.65	0.46	1.32	(3)
BP base.[a]		0.77	0.12	-0.38	-0.86	0.83	(3)
BP augm.[a]	HHTATL	1.00	0.65	0.12	-0.29	1.34	(3)

[a] Identifying restrictions for the BP approach can be found in Table 4 in Appendix B.

10 Effects of Fiscal Policy Changes – Baseline vs. Augmented Models

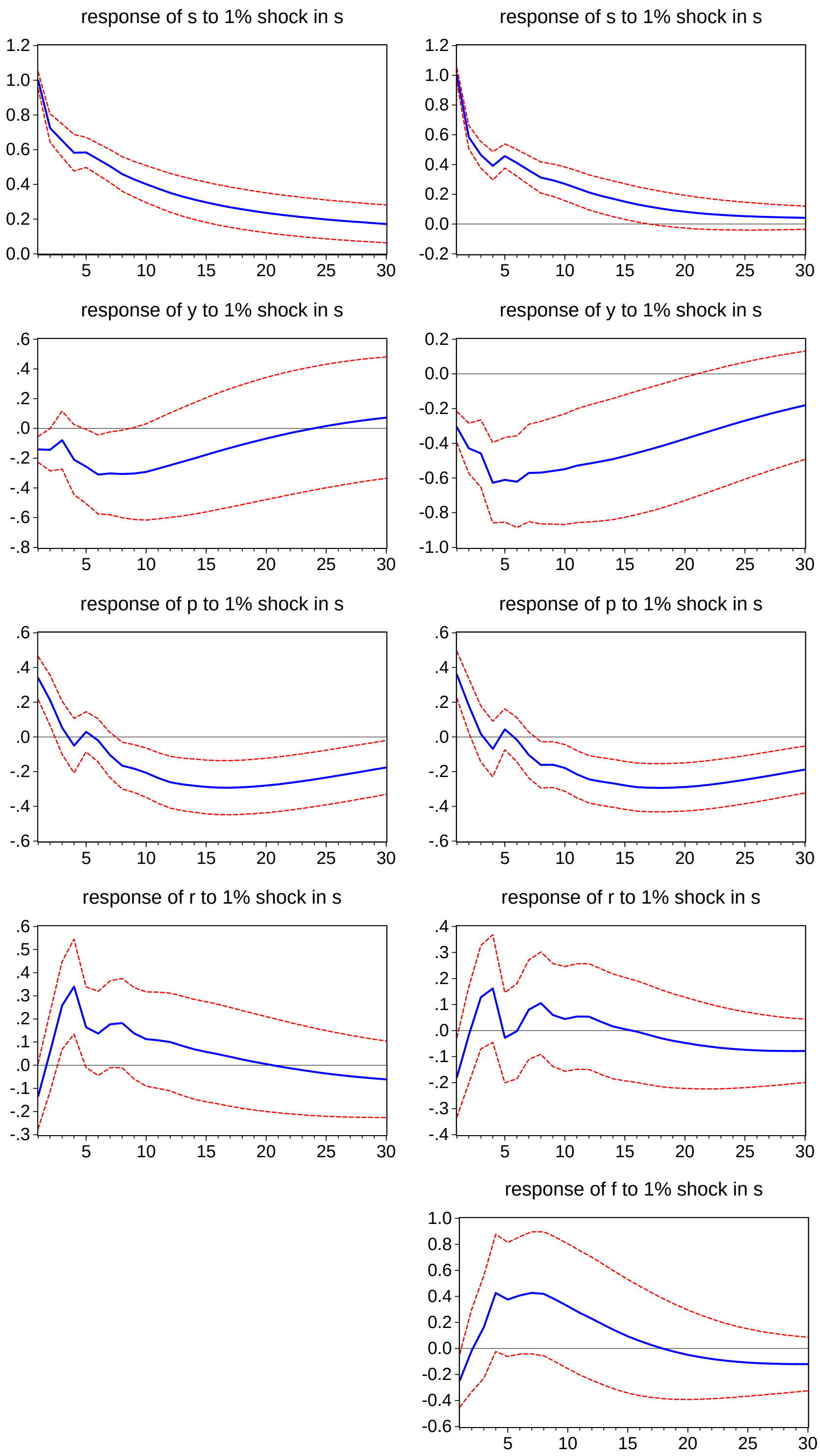
The previous sections have shown that there are potential identification and omitted variable biases in standard approaches to estimating fiscal multipliers with respect to changes in private debt and wealth. In order to quantify the impact of the bias, we now compare impulse-responses of shocks to budget variables in the baseline models to those of the augmented models.

First, we simulate a 1% of GDP improvement in the CAPB-to-GDP ratio, which is interpreted as a fiscal consolidation. Figure 8 presents the IRFs for the baseline and augmented model, respectively. Both models show a transitory, but lasting contraction in GDP after the fiscal consolidation. The reaction is more pronounced and lasting, however, for the model that controls for households' wealth-to-debt ratio; the response function of GDP remains significant for a much longer horizon. Impact and peak multipliers more than double and there is an absolute difference in the peak multiplier of about 0.3. The cumulative multipliers are much higher for the augmented model (about three times as high), though reliability of the results naturally lowers with an increasing horizon.

For digits of multipliers at selected horizons, refer to Table 1, where $HHTATL$ represents the wealth-to-debt ratio of households. It displays the multipliers derived from the IRFs for selected horizons. Multipliers are calculated either as the *impact* response of GDP divided by the initial fiscal impulse (FI)

$$k = \frac{\Delta y_t}{\Delta FI_t}, \tag{25}$$

Figure 8: Impulse Response Functions for CAPB VAR – Baseline (left), Augmented (right)



or as the *cumulative* response function of GDP divided by the cumulative fiscal impulse function

$$k = \frac{\sum_h \Delta y_{t+h}}{\sum_n \Delta FI_{t+h}}, \tag{26}$$

or as the *peak* response of GDP with respect to the initial fiscal impulse

$$k = \frac{\max_h \Delta y_{t+h}}{\Delta FI_t}, \tag{27}$$

where $\Delta(\cdot)$ marks deviation from the steady state.

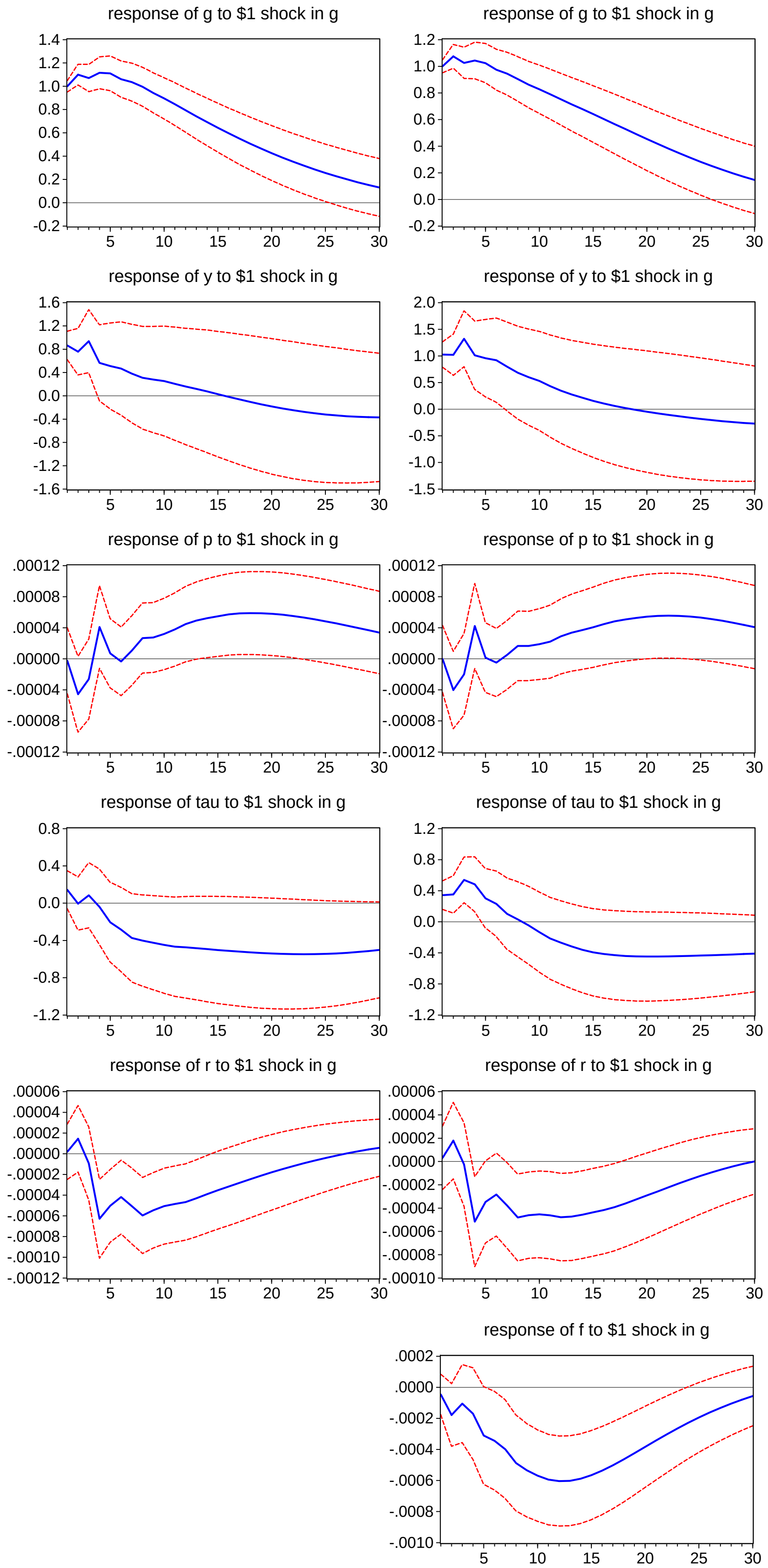
The other variables react similarly for the two models, with inflation increasing instantaneously – perhaps due to the consolidation being tax-driven to some extent –, followed by a long-lasting decline in line with the slowing down of the economy. The short-term real interest rate plausibly declines on impact after the consolidation and increases later on, indicating a very slow reaction of the nominal interest rate itself and the real rate being largely driven by inflation.[7] The households wealth-to-debt ratio itself exhibits an instantaneous fall, followed by a pronounced increase later on. Given that the consolidation to some extent consists of tax hikes, an instantaneous fall in the wealth-to-debt ratio is plausible in terms of consumption smoothing. The subsequent increase could reflect households' deleveraging as a reaction to falling GDP after the consolidation shock.

Turning to the RA model, results by and large reproduce those of the CAPB VAR. Figure 9 shows the impulse-responses to a \$ 1 per capita increase in government spending, i. e. a fiscal expansion. In both cases the change in GDP as measured in \$ per capita is positive, with a slight net crowding-out effect for the baseline case while there is net crowding-in during the first quarters for the augmented model. The GDP response remains significantly positive for twice as long in the augmented model. Even though differences in the multipliers are not that pronounced in relative terms as compared to the CAPB VAR, the absolute difference of the impact and peak multiplier is even slightly higher, ranging from 0.3 to 0.4. Over the whole set of horizons, cumulative multipliers of baseline and augmented models also differ in the range of 0.3 to 0.4.

The behavior of the other variables in the model is plausible and in line with what has been said for the CAPB VAR model. Inflation increases only slightly and only after some quarters. The real interest rate jumps for one quarter but then falls, again, reflecting

[7] The real interest rate in all the models we tested largely reflects the change in inflation, while there is no remarkable stand-alone reaction of the interest rate. Testing an alternative model with the nominal effective federal funds rate, did not alter the IRFs of the other variables.

Figure 9: Impulse Response Functions for RA VAR – Baseline (left), Augmented (right)



the dynamics of the inflation rate by which it seems to be largely driven. Behavior of the interest rate is largely in line with findings in Dungey and Fry (2009); Chung and Leeper (2007); Mountford and Uhlig (2009). Revenues rise on impact, and significantly-so for the augmented model, but turn insignificant soon after. The wealth-to-debt ratio decreases significantly after some quarters, possibly a reaction of households venturing higher indebtedness due to the rise in GDP.

For the baseline and augmented models following the BP approach, impulse-responses of a government spending shock of $ 1 per capita are presented in Figure 10, respectively. The responses are pretty much in line with those of the RA approach, resulting in positive multipliers for both versions on impact and peak, with a slight crowding-out for the baseline model, while there is a slight crowding-in for the augmented version. However, for the BP approach, the GDP response turns insignificant and negative more quickly, yielding insignificantly negative multipliers at later horizons. The difference between the baseline and the augmented model, again, remains rather stable at a level of 0.3 to 0.6 over the whole set of horizons.

The other variables react similarly for both versions. Inflation responds positively and remains significant over a long horizon while the real interest rate, again, to a large part reflects the dynamics of the inflation rate. The negative interest rate reaction is consistent with findings in the literature (Dungey and Fry 2009; Chung and Leeper 2007; Mountford and Uhlig 2009). Taxes are significantly positive on impact and turn significantly negative after some years. For the augmented model, the wealth-to-debt ratio, in line with the impulse-responses of the RA specification, turns significantly negative.

Generally, we find empirical support for our hypothesis. Estimated multipliers are considerably larger when controlling for private debt and wealth levels in otherwise standard models. This is the case for several established approaches to identification over different horizons, see Table 1 for a summary.

11 Robustness

We test the robustness of our results against the dimensions of sample size, alternative control variables, alternative aggregate demand components and alternative budget variables.

Results for multipliers with respect to alternative sample sizes can be found in Table 2. The first rows show results for a sample excluding the recent crisis years. Multipliers for the augmented models are still on a higher level than those of the baseline models

Figure 10: Impulse Response Functions for BP VAR – Baseline (left) and Augmented (right)

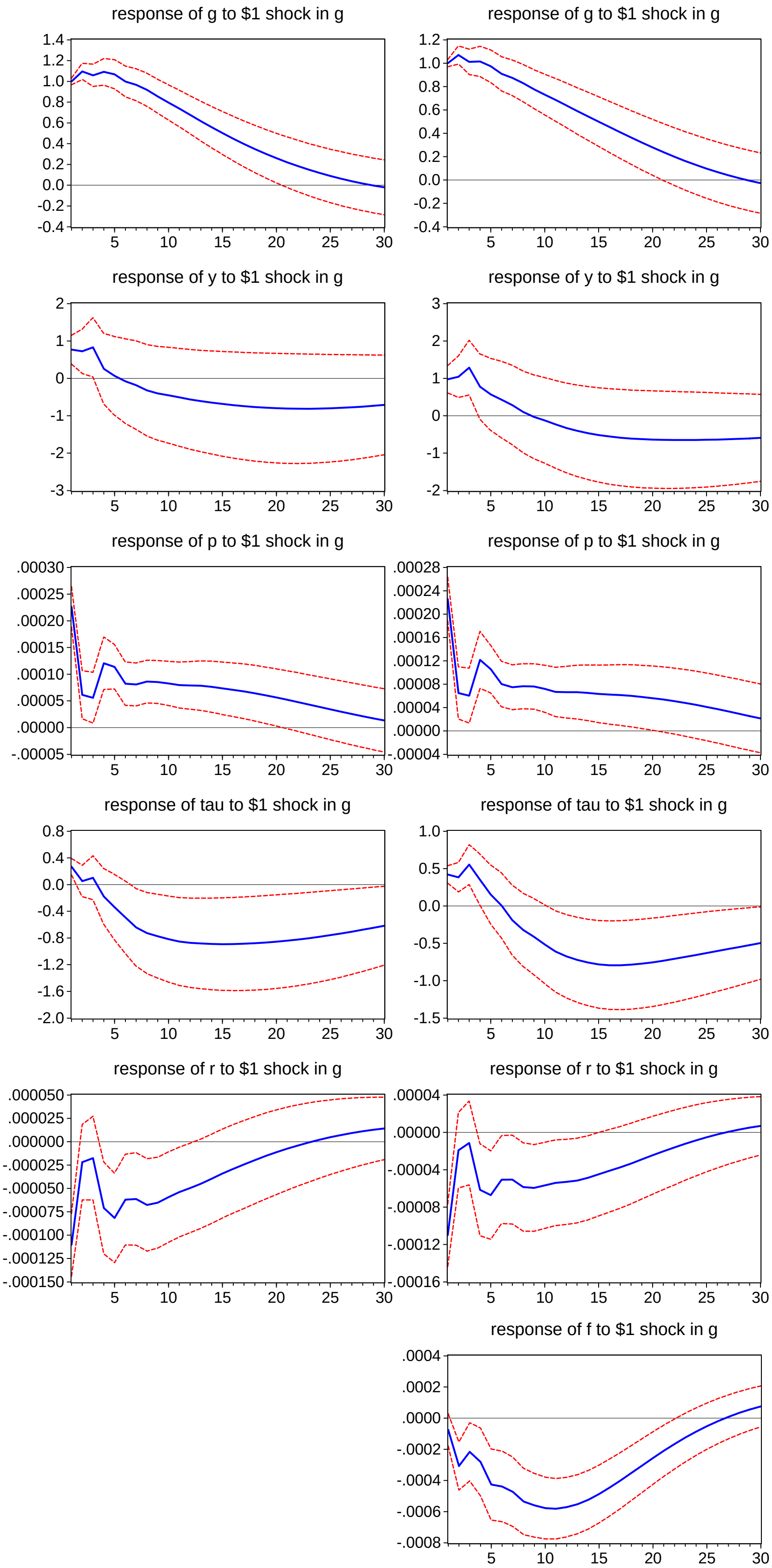


Table 2: Robustness of Multipliers for Baseline and Augmented Models – Sample

Model	f	Impact Quarter 1	Cumulative 10	Cumulative 20	Cumulative 30	Peak	(Quarter)
1960-2007							
CAPB base.		0.15	0.80	1.16	1.22	0.69	(9)
CAPB augm.	HHTATL	0.26	1.04	1.43	1.53	0.70	(6)
RA base.		0.83	1.00	1.26	1.34	1.27	(10)
RA augm.	HHTATL	0.97	1.14	1.23	1.30	1.29	(6)
BP base.[a]		0.90	1.01	1.19	1.25	1.25	(3)
BP augm.[a]	HHTATL	1.11	1.33	1.28	1.32	1.71	(3)
1960-1985							
CAPB base.		0.21	1.68	2.90	3.08	1.38	(11)
CAPB augm.	HHTATL	0.22	1.44	2.70	2.87	1.05	(10)
RA base.		0.77	1.48	2.16	2.10	2.44	(13)
RA augm.	HHTATL	0.76	1.12	1.99	1.90	1.81	(13)
BP base.[a]		0.74	1.39	2.08	2.03	1.97	(13)
BP augm.[a]	HHTATL	0.80	1.01	1.83	1.73	1.26	(13)
1985-2012							
CAPB base.		0.22	-0.47	-1.07	-1.50	0.22	(1)
CAPB augm.	HHTATL	0.38	-0.04	-0.66	-1.14	0.47	(2)
RA base.		1.52	1.32	0.08	-0.91	2.54	(3)
RA augm.	HHTATL	1.59	1.51	0.18	-0.79	2.77	(3)
BP base.[a]		2.66	3.16	0.82	-0.33	4.04	(3)
BP augm.[a]	HHTATL	3.01	3.45	1.00	-0.16	4.42	(3)

[a] Identifying restrictions for the BP approach can be found in Table 4 in Appendix B.

(with the exception of the RA model for longer horizons), but differences are smaller. Estimated multipliers are generally higher than in the full sample. However, results from the augmented model are more robust to the exclusion of the crisis years as multipliers are rather similar, while the baseline model is more sensitive to inclusion vs. exclusion of the crisis years. Put differently, the augmented model absorbs the specific effects of the crisis, while the baseline specification does not seem to handle them robustly. This is reasonable as the crisis in the US was largely driven by a private sector asset meltdown, which the augmented model can take into account.

The mid and lower rows of Table 2 show the results from a split of the full sample in half. These results may be taken with a pinch of salt since degrees of freedom shrink a lot. Multipliers for the early period are much higher for both specifications on longer horizons while they are lower on impact. Again, the results of the augmented models are more robust to the choice of sample. This time, however, multipliers from the baseline models exceed those of the augmented models in most cases on longer horizons. This specialty of the earlier subsample is in line with Lindner (2013), who finds that the wealth effect on private demand was negative before the mid-1980s and only positive afterwards, due to structural changes in financial market behavior. With a negative wealth effect, multipliers would be upward-biased in the presence of financial market movements.

For the later period, multiplier effects seem to be more pronounced for shorter horizons, but more short-lived and at a lower level on medium and longer horizons, a finding that also shows up in the US estimations of Perotti (2005), Bilbiie et al. (2008) and Caldara and Kamps (2008). However, in line with the full sample results, there is a considerable and rather stable difference in the estimated multipliers, with those of the augmented models exceeding those of the baseline by about 0.2 to 0.4 for the whole set of horizons.

In any case, results of the augmented models show more subsample stability than the baseline models, which is probably due to its capacity to capture the changing influence of financial market movements over the decades.

Table 3 summarizes robustness checks in other dimensions. In the first rows, an alternative ordering of the variables in the Choleski-decomposed models is applied, with the wealth-to-debt ratio ordered first instead of last. Results do not change much as compared to Table 1, except for the augmented CAPB model on longer horizons, whose multipliers are now somewhat lower, but still considerably above those of the baseline model.

In the following rows, we tested alternative control variables, namely, instead of using the wealth-to-debt ratio, we put both households' total assets and total liabilities, both

Table 3: Robustness of Multipliers for Baseline and Augmented Models – Specification

Model	f	Impact Quarter 1	Cumulative 10	20	30	Peak (Quarter)	
Alternative Ordering - f first							
CAPB augm.	HHTATL first	0.30	0.92	1.26	1.46	0.55	(4)
RA augm.	HHTATL first	1.03	0.98	0.72	0.53	1.36	(3)
Alternative Controls							
CAPB base.		0.14	0.40	0.45	0.36	0.31	(6)
CAPB augm.	HHTA + HHTL	0.33	0.96	1.07	0.96	0.60	(4)
CAPB augm.	S&P500 + NFPSD	0.29	1.37	2.33	2.74	0.64	(10)
RA base.		0.86	0.52	0.33	0.12	0.94	(3)
RA augm.	HHTA, HHTL	1.04	1.62	1.69	1.39	1.74	(7)
RA augm.	S&P500, NFPSD	1.03	2.04	2.24	1.80	2.13	(10)
BP base.		0.77	0.12	-0.38	-0.86	0.83	(3)
BP augm.	HHTA, HHTL	0.93	1.21	1.29	1.14	1.35	(3)
BP augm.	S&P500, NFPSD	0.95	1.77	2.00	1.54	1.50	(11)
Response of Private Consumption							
CAPB base.[a]		0.17	0.59	0.58	0.45	0.39	(6)
CAPB augm.[a]	HHTATL	0.29	1.10	1.40	1.53	0.58	(6)
RA base.		0.38	0.27	0.16	0.04	0.49	(2)
RA augm.	HHTATL	0.50	0.54	0.46	0.36	0.64	(2)
Shock to Government Consumption							
RA base.		0.52	0.57	0.66	0.52	0.83	(10)
RA augm.	HHTATL	0.77	1.18	1.20	1.09	1.34	(6)
BP base.		0.33	0.00	-0.24	-0.60	0.33	(1)
BP augm.	HHTATL	0.63	0.69	0.41	0.14	0.93	(3)

[a] These numbers should not be misinterpreted as multipliers, since they show the percentage change of private consumption per capita to an increase in the structural deficit of 1% of GDP.

in real terms per capita and in logs ($HHTA, HHTL$) into the model. Alternatively, we used the log of the deflated S&P 500 index ($S\&P500$) and the log of real non-financial private sector debt per capita ($NFPSD$) as proxies for private wealth and debt, respectively. These augmented models now include six endogenous variables in the case of the CAPB VAR and seven endogenous variables in the case of the RA and BP specification. We find very robust results for shorter horizons of the impulse-responses with considerably higher multipliers for the augmented models vs. the baseline models. In comparison to the augmented models with the wealth-to-debt ratio, differences are even more pronounced for longer horizons in Table 3, especially in the $S\&P500, NFPSD$ cases; however, confidence bands of the GDP response are then wide.

The next couple of rows of Table 3 present the results of an exercise where GDP is replaced by private consumption expenditures in the VAR models. Due to a lack of prior information on elasticities of the other variables to changes in private consumption, we did not perform this robustness test for the BP approach. For the other methods of identification, however, our earlier results are confirmed. There is crowding in of private consumption, which is much stronger for the augmented models.

Results remain robust when general government spending is replaced by government consumption in the vector of endogenous variables, as displayed in the lower rows of Table 3. Multipliers are lower on average, but the difference between the baseline and augmented models persists.

12 Conclusions

We have investigated whether movements in credit and asset markets imply both a biased identification and an omitted variable bias in standard multiplier estimation techniques that rely on prior information regarding endogeneity of the fiscal budget with respect to the normal business cycle. In line with a growing literature (Guajardo et al. 2011; Perotti 2011; Yang et al. 2013), we have argued that in the presence of movements in asset and credit markets standard approaches can lead to wrong identifications that downward-bias the estimated multiplier both in a market upswing and downswing.

To test this hypothesis, we set up a formal framework to pin down the impact of the omission of these channels on estimated multiplier values; the derivation shows that there should be a downward-bias of estimated multipliers in the presence of movements in the wealth-to-debt ratio in both directions. We then quantify the possible bias on multiplier estimations by employing empirical models of established identification schemes, namely the CAPB and two versions of the structural VAR approach, and compare their results vis-à-vis multiplier effects in the case of inclusion vs. exclusion of private debt and wealth proxies. For the CAPB identification we use a recursive VAR and compare the results of a fiscal consolidation shock. For the structural VAR, we test the recursive identification (Fatás and Mihov 2001) as well as the Blanchard and Perotti (2002) approach in a standard VAR based on Caldara and Kamps (2008) to estimate the multipliers from government spending impulses.

Our results confirm the hypothesis of Guajardo et al. (2011). We find downward-biased multipliers from identifications based on prior information regarding business cycle endogeneity, such as using the CAPB and standard structural VAR approaches, as they overlook the influence of asset and credit market movements on GDP and the fiscal budget. Multipliers are on average about 0.3 to 0.6 units higher when taking these influences into account. These findings are robust to numerous alternative specifications. Fiscal consolidations thus are more likely to be contractionary and could be more harmful to growth than expected from the results of some of the previous literature.

This line of research could be extended to other country samples, especially to members of the Euro Area. Moreover, future research could take into account whether there is an asymmetry of the bias in the upswing and downswing of the financial cycle, which would connect our findings with those of a state dependence of the fiscal multiplier (Auerbach and Gorodnichenko 2012; Batini et al. 2012; Fazzari et al. 2012; Ferraresi et al. 2013).

References

Alesina, A. and S. Ardagna (2010). Large changes in fiscal policy: taxes versus spending. *NBER/Tax Policy & the Economy 24*(1), 35–68.

Auerbach, A. J. and Y. Gorodnichenko (2012). Measuring the Output Responses to Fiscal Policy. *American Economic Journal: Economic Policy 4*(2), 1–27.

Batini, N., G. Callegari, and G. Melina (2012). Successful Austerity in the United States, Europe and Japan. IMF Working Paper WP/12/190.

Beetsma, R., M. Giuliodori, and F. Klaassen (2006). Trade spill-overs of fiscal policy in the European Union: a panel analysis. *Economic Policy 41*(48), 639–687.

Bénétrix, A. S. and P. Lane (2011). Financial Cycles and Fiscal Cycles. Mimeo.

Bilbiie, F. O., A. Meier, and G. J. Müller (2008). What Accounts for the Changes in U.S. Fiscal Policy Transmission? *Journal of Money, Credit, and Banking 40*(7), 1439–1469.

Blanchard, O. and R. Perotti (2002). An Empirical Characterization of the Dynamic Effects of Changes in Government Spending and Taxes on Output. *Quarterly Journal of Economics 117*(4), 1329–1368.

Borio, C. (2012). The financial cycle and macroeconomics: What have we learnt? BIS Working Papers 395.

Bornhorst, F., G. Dobrescu, A. Fedelino, J. Gottschalk, and T. Nakata (2011). When and How to Adjust Beyond the Business Cycle? A Guide to Structural Fiscal Balances. IMF Technical Notes and Manuals 11/02.

Caldara, D. and C. Kamps (2008). What are the effects of fiscal policy shocks? A VAR-based comparative analysis. European Central Bank working paper series 877.

Case, K. E., J. M. Quigley, and R. J. Shiller (2005). Comparing Wealth Effects: The Stock Market versus the Housing Market. *Advances in Macroeconomics 5*(1), 1–36.

Chung, H. and E. M. Leeper (2007). What Has Financed Government Debt? NBER working paper 13425.

Congressional Budget Office (2013). The Effects of Automatic Stabilizers on the Federal Budget as of 2013. Technical report.

Corsetti, G., A. Meier, and G. J. Müller (2012). What Determines Government Spending Multipliers? IMF Working Paper WP/12/150.

Dungey, M. and R. Fry (2009). The Identification of Fiscal and Monetary Policy in a Structural VAR. *Economic Modelling 26*(6), 1147–1160.

Dynan, K. (2012). Is A Household Debt Overhang Holding Back Consumption? Mimeo.

Eggertson, G. B. and P. Krugman (2012). Debt, Deleveraging, and the Liquidity Trap: A Fisher-Minsky-Koo Approach. *Quarterly Journal of Economics 127*(3), 1469–1513.

Eschenbach, F. and L. Schuknecht (2002). Asset prices and fiscal balances. European Central Bank working paper series 141.

Eschenbach, F. and L. Schuknecht (2004). Budgetary risks from real estate and stock markets. *Economic Policy 19*(39), 313–346.

Fanchon, P. and J. Wendel (1992). Estimating VAR models under non-stationarity and cointegration: alternative approaches for forecasting cattle prices. *Applied Economics 24*(2), 207–217.

Fatás, A. and I. Mihov (2001). The Effects of Fiscal Policy on Consumption and Employment: Theory and Evidence. CEPR Discussion Papers 2760.

Favero, C. and F. Giavazzi (2007). Debt and the Effects of Fiscal Policy. NBER working paper 12822.

Fazzari, S. M., J. Morley, and I. Panovska (2012). State-Dependent Effects of Fiscal Policy. Mimeo.

Ferraresi, T., A. Roventini, and G. Fagiolo (2013). Fiscal Policies and Credit Regimes: A TVAR Approach. University of Verona Department of Economics Working Paper Series 3.

Gechert, S. (2013). What fiscal policy is most effective? A Meta Regression Analysis. IMK working paper 117.

Girouard, N. and R. Price (2004). Asset Price Cycles, "One-Off" Factors and Structural Budget Balances. OECD Economics Department Working Papers 391.

Giudice, G., A. Turrini, and J. i. t. Veld (2007). Non-Keynesian Fiscal Adjustments? A Close Look at Expansionary Fiscal Consolidations in the EU. *Open Economies Review 18*(5), 613–630.

Graham, J. R. (2000). How Big Are the Tax Benefits of Debt? *Journal of Finance 55*(5), 1901–1941.

Guajardo, J., D. Leigh, and A. Pescatori (2011). Expansionary Austerity: New International Evidence. IMF Working Paper WP/11/158.

IMF (2012). World Economic Outlook April 2012: Growth Resuming, Dangers Remain. Technical Report World Economic and Financial Surveys, Washington DC.

Leeper, E. M., N. Traum, and T. B. Walker (2011). Clearing up the fiscal multiplier morass. NBER working paper 17444.

Lindner, F. (2013). The housing wealth effect on consumption reconsidered. IMK working paper 115.

Lütkepohl, H. (2006). *New Introduction to Multiple Time Series Analysis.* Berlin u. a.: Springer.

Miller, M. H. (1977). Debt and Taxes. *Journal of Finance 32*(2), 261–275.

Mineshima, A., M. Poplawski-Ribeiro, and A. Weber (2013). Size of Fiscal Multipliers. Mimeo.

Morris, R. and L. Schuknecht (2007). Structural balances and revenue windfalls: the role of asset prices revisited. European Central Bank working paper series 737.

Mountford, A. and H. Uhlig (2009). What are the Effects of Fiscal Policy Shocks? *Journal of Applied Econometrics 24*(6), 960–992.

Mourre, G., G.-M. Isbasoiu, D. Paternoster, and M. Salto (2013). The cyclically-adjusted budget balance used in the EU fiscal framework: an update. European Commission Economic Papers 478.

Perotti, R. (2005). Estimating the effects of fiscal policy in OECD countries. CEPR Discussion Papers 4842.

Perotti, R. (2011). The "Austerity Myth": Gain Without Pain? NBER working paper 17571.

Phillips, P. C. and S. N. Durlauf (1986). Multiple Time Series Regression with Integrated Processes. *Review of Economic Studies 53*(4), 473–495.

Poterba, J. M. (2000). Stock Market Wealth and Consumption. *Journal of Economic Perspectives 14*(2), 99–118.

Price, R. and T.-T. Dang (2011). Adjusting Fiscal Balances for Asset Price Cycles. OECD Economics Department Working Papers 868.

Romer, C. D. and D. H. Romer (2010). The macroeconomic effects of tax changes: estimates based on a new measure of fiscal shocks. *American Economic Review 100*(3), 763–801.

Sims, C. A., J. H. Stock, and M. W. Watson (1990). Inference in Linear Time Series Models with Some Unit Roots. *Econometrica 58*(1), 113–144.

West, K. D. (1988). Asymptotic Normality, When Regressors Have a Unit Root. *Econometrica 56*(6), 1397–1417.

Woodford, M. (2011). Simple Analytics of the Government Expenditure Multiplier. *American Economic Journal: Macroeconomics 3*(1), 1–35.

Yang, W., J. Fidrmuc, and S. Ghosh (2013). Macroeconomic Effects of Fiscal Adjustment: A Tale of two Approaches. CESifo working paper 4401.

Appendix A

In order to solve the structural model and identify the structural shocks ε_t that are central for quantitative policy simulations, we first need to estimate the VAR in reduced form

$$\mathbf{\Gamma}(L)X_t = v + \boldsymbol{A}^{-1}\boldsymbol{B}\varepsilon_t \tag{28}$$

$$\mathbf{\Gamma}(L)X_t = v + u_t \tag{29}$$

and pick the K-dimensional vector of reduced form residuals u_t.

$$u_t = \boldsymbol{A}^{-1}\boldsymbol{B}\varepsilon_t \tag{30}$$

Equation (30) represents the relation between reduced form shocks u_t and structural form shocks ε_t. Due to the autoregressive structure of the model the reduced form residuals u_t are almost certainly correlated with each other and therefore inappropriate to simulate exogenous policy changes. Thus, in a second step we solve for the structural shocks via

$$\varepsilon_t = \boldsymbol{B}^{-1}\boldsymbol{A}u_t. \tag{31}$$

This is done by taking the $K \times K$ variance-covariance matrix Σ_u of the reduced form residuals and by assuming ortho-normality of the structural shocks ($\varepsilon_t \sim (0, \Sigma_\varepsilon = I_K)$).[8] From (30) follows that

$$\Sigma_u = \boldsymbol{A}^{-1}\boldsymbol{B}\Sigma_\varepsilon\boldsymbol{B}'(\boldsymbol{A}^{-1})' = \boldsymbol{A}^{-1}\boldsymbol{B}\boldsymbol{B}'(\boldsymbol{A}^{-1})'. \tag{32}$$

Since (32) is over-parameterized, as it contains $2K^2$ unknowns and only $K(K+1)/2$ equations, we need to impose at least $2K^2 - K(K+1)/2$ restrictions from prior economic information on some parameters of $\boldsymbol{A}$ and $\boldsymbol{B}$ in order to calculate their remaining items. Before we describe setting of restrictions in more detail for our specific VAR models, we briefly set out the remainder of the procedure to derive *impulse-response functions* (IRFs): With just identified matrices $\boldsymbol{A}$ and $\boldsymbol{B}$, we are able to derive the structural shocks from (31). Afterwards, the *structural vector moving average representation* (SVMA) of the VAR can be determined:

$$X_t = \mu + \boldsymbol{\Theta}(L)\varepsilon_t = \mu + \sum_{h=0}^{p} \boldsymbol{\Theta}_i\varepsilon_{t-i} \tag{33}$$

with $\boldsymbol{\Theta}(L) = \boldsymbol{\Gamma}(L)^{-1}\boldsymbol{A}^{-1}\boldsymbol{B}$, $\mu = \boldsymbol{\Gamma}(L)^{-1}v$ and h being the respective horizon of interest. Note that $\boldsymbol{\Gamma}(L)$ must be invertible to allow for a MA representation.

Finally, the IRFs of the endogenous variables i to unit structural shocks to variable j at horizon h can be computed from the SVMA via

$$\boldsymbol{\Upsilon}_{i,j,h} = \frac{\partial x_{i,t+h}}{\partial \varepsilon_{j,t}}. \tag{34}$$

They show the deviation of variable i at horizon h from a steady state path of the model when the system is hit by an exogenous shock to variable j and can be interpreted as multipliers if they are scaled correctly.

Appendix B

[8] The assumption of ortho-normality is not restrictive. It ensures that the structural shocks are random and independent of one another and it pre-sizes their variance to easily interpret impulse-responses later on. No information is lost, since the settings made here will be reflected in the coefficients of the $\boldsymbol{A}$ and $\boldsymbol{B}$ matrices.

Table 4: Identifying restrictions set for the BP models

	$\alpha_{\tau y}$ [a]	α_{gp} [a]	$\alpha_{\tau p}$ [a]	α_{gf}	α_{yf}	α_{pf}	$\alpha_{\tau f}$	α_{rf}
Full Sample	1.85	-0.5	1.25	-0.05	0.06	0	0.1	0.03
1960-2007	1.85	-0.5	1.25	-0.07	0.04	0	0.1	0.03
1960-1985	1.75	-0.5	1.09	-0.07	-0.02	0	-0.2	0
1985-2012	1.97	-0.5	1.40	-0.04	0.04	0	0.1	0

[a] Source: Perotti (2005)

Table 5: Johansen tests for cointegration

Sample 1960:1 2012:4
Lags interval for differencec endog: 1 to 2
Selected (0.05 level*) Number of Cointegrating Relations by Model

Data Trend	None	None	Linear	Linear	Quadratic
Test Type	No Intercept	Intercept	Intercept	Intercept	Intercept
No Trend	No Trend	No Trend	Trend	Trend	
CAPB baseline 4 variables					
Trace	1	1	0	0	0
Max-Eig	1	1	0	0	0
CAPB augmented 5 variables					
Trace	1	1	1	1	1
Max-Eig	1	1	1	1	1
SVAR baseline 5 variables					
Trace	1	1	0	0	0
Max-Eig	1	1	0	0	0
SVAR augmented 6 variables					
Trace	1	2	0	0	1
Max-Eig	1	2	1	1	1

*Critical values based on MacKinnon-Haug-Michelis (1999)

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