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Essais sur la politique fiscale

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Sommaire

Cette thèse est composée de trois essais qui traitent de la politique fiscale. L'évaluation empirique des effets macroéconomiques de cette politique est à la fois difficile et controversée, puisque certaines hypothèses doivent être imposées afin d'isoler les changements exogènes et imprévus des taxes et des dépenses publiques. Les études basées sur les Vecteurs Autorégressifs Structurels (SVAR) identifient les chocs fiscaux en imposant des restrictions arbitraires sur les interactions entre les variables du système.

Dans le premier essai de la thèse, nous identifions les chocs fiscaux en exploitant l'hétéroscédasticité conditionnelle des innovations structurelles, ce qui nous permet de relâcher les hypothèses d'identification couramment utilisées dans la littérature empirique. Nous utilisons cette méthodologie afin d'évaluer les effets macroéconomiques des chocs fiscaux aux États-Unis au cours des périodes pré- et post-1979. D'abord, nos résultats montrent des différences considérables entre les deux périodes en ce qui concerne les effets de ces chocs sur l'activité économique. De plus, nous constatons que l'augmentation des dépenses publiques est, en général, plus efficace que les réductions d'impôts pour stimuler l'activité économique. Enfin, notre étude contribue significativement à la littérature existante en testant formellement les restrictions d'identification communément utilisées.

Très peu d'études empiriques ont évalué les effets des chocs fiscaux, en particulier les chocs de taxes, sur l'ajustement extérieur des économies. Le deuxième essai de la thèse mesure les effets des chocs de taxes et dépenses publiques sur le compte courant et le taux de change réel pour un échantillon de quatre pays industrialisés. Notre analyse est basée sur un SVAR non contraint où les chocs sont identifiés en exploitant la variance conditionnelle des perturbations structurelles. Les résultats qui découlent de ce travail sont les suivants : (i) il y a très peu d'évidence en faveur de l'hypothèse des déficits jumeaux (i.e. relation positive entre les déficits extérieur et budgétaire), (ii) les effets d'une baisse non anticipée des taxes ne sont pas compatibles avec les prédictions des modèles économiques standards, excepté pour le cas américain, et (iii) les approches d'identification traditionnelles sous évaluent la dépréciation réelle conditionnellement à un choc positif de dépenses publiques.

Le troisième essai de la thèse documente et explique le comouvement entre les déficits extérieur et budgétaire pour un échantillon de pays en développement couvrant la période post-1960. D'abord, les estimations indiquent que la covariance empirique entre ces déficits est toujours positive et est statistiquement significative pour la plupart des pays considérés. Ceci est cohérent avec les résultats antérieurs obtenus à partir de régressions effectuées sur des données longitudinales. De plus, et pour la plupart des pays de l'échantillon, les données permettent de répliquer la covariance empirique à partir d'un modèle de petites économies ouvertes à générations imbriquées avec des biens de consommations hétérogènes. En outre, la covariance prédite est induite par des chocs qui sont étroitement liés aux conditions internes telles que les ressources domestiques et les mesures fiscales et, dans une moindre mesure, aux conditions extérieures telles que le taux d'intérêt mondial, le taux de change réel, et les termes de l'échange. Cette conclusion contraste avec celle documentée à partir des études à formes réduites, caractérisant soit le comportement individuel du déficit extérieur, ou encore, celui du déficit budgétaire.

Mots clés: Politique fiscale, Dépenses gouvernementales, Taxes, Vecteur autorégressif structurel, Identification, Compte courant, Taux de change, Décomposition de covariance, Conditions internes et externes, Modèle de petite économie ouverte à générations imbriquées avec des biens hétérogènes.

Summary

This thesis is composed of three essays related to fiscal policy. Measuring the effects of discretionary fiscal policy is both difficult and controversial, as some identifying assumptions need to be made to isolate exogenous and unanticipated changes in taxes and government spending. Studies based on structural vector autoregressions (SVAR) typically achieve identification by restricting the contemporaneous interaction of fiscal and non-fiscal variables in a rather arbitrary way.

In the first essay of the thesis, we relax those restrictions and identify fiscal policy shocks by exploiting the conditional heteroscedasticity of the structural disturbances. We use this methodology to evaluate the macroeconomic effects of fiscal policy shocks in the U.S. before and after 1979. Our results show substantive differences in the economy's response to government spending and tax shocks across the two periods. Importantly, we find that increases in public spending are, in general, more effective than tax cuts in stimulating economic activity. A key contribution of this study is to provide a formal test of the identifying restrictions commonly used in the literature.

Relatively little empirical evidence exists about countries' external adjustment to changes in fiscal policy and, in particular, to changes in taxes. The second essay of the thesis addresses this question by measuring the effects of tax and government spending shocks on the current account and the real exchange rate in a sample of four industrialized countries. Our analysis is based on a SVAR in which the interaction of fiscal variables and macroeconomic aggregates is left unrestricted. Identification is instead achieved by exploiting the heteroscedasticity of the structural disturbances. Three main findings emerge: (i) the data provide little support for the twin-deficit hypothesis (i.e. a positive relation between the external and budget deficits) , (ii) the estimated effects of unexpected tax cuts are generally inconsistent with the predictions of standard economic models, except for the U.S., and (iii) the puzzling real depreciation triggered by an expansionary public spending shock is substantially larger in magnitude than predicted by traditional identification approaches.

The third essay documents and explains the positive comovement between the external and budget deficits of developing countries covering the post-1960 period. First, the estimates indicate that the empirical covariance between these deficits is always positive and is statistically significant for many cases. This is consistent with previous findings obtained from panel regressions. Second, the empirical covariance is close to that predicted from a tractable small open-economy, overlapping-generation model with heterogeneous goods. Also, the predicted covariance is induced by shocks which are closely related to internal conditions such as domestic resources and fiscal policies, and to a much lesser extent to external conditions such as the world interest rate, real exchange rate, and terms of trade. This analysis explaining the joint behavior of the external and budget deficits contrasts with earlier single-equation studies characterizing the individual behavior of either the external deficit or budget deficit.

Keywords: Fiscal policy, Government spending, Taxes, Structural vector auto-regression, Identification, Current account, Exchange rate, Covariance decomposition, Internal and external conditions, Small-open economy overlapping-generation model with heterogeneous goods.

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Contribution des auteurs

Le premier essai intitulé «Measuring the Effects of Fiscal Policy» est co-écrit avec les professeurs Hafedh Bouakez et Michel Normandin. Le deuxième essai intitulé «Fiscal Policy and External Adjustment: New Evidence» est également corédigé avec les professeurs Hafedh Bouakez et Michel Normandin. Enfin, le troisième essai intitulé «External and Budget Deficits in Some Developing Countries» est rédigé en collaboration avec le professeur Michel Normandin. Les contributions des auteurs sont égales.

Introduction générale

La récente crise économique mondiale est très différente des précédentes, ce qui remet à l'avant-scène l'efficacité des politiques économiques conventionnelles. Dans un environnement où les taux d'intérêt sont presque nuls, la mise en œuvre de la politique monétaire par le biais des instruments traditionnels devient problématique. Pour cette raison, certaines banques centrales ont eu recours à des instruments monétaires non traditionnels, tels que l'assouplissement du crédit et l'assouplissement quantitatif. Dans ce contexte, la politique fiscale se présente également comme un instrument alternatif. Malheureusement, les théories économiques procurent des prédictions différentes en ce qui a trait aux effets de la politique fiscale sur l'allocation des ressources domestiques. À titre d'exemple, l'approche keynésienne prédit une réponse positive de la consommation privée suite à un accroissement des dépenses publiques. Cette réponse est engendrée par un effet de revenu positif, appelé également l'effet multiplicateur. À l'opposé, la théorie classique met l'accent sur un effet de richesse négatif afin de prédire une corrélation négative entre les consommations privée et publique, conditionnellement à un choc de dépenses gouvernementales.

Sur le plan empirique, les études se classent selon deux méthodologies: l'approche narrative et les Vecteurs Autorégressifs Structurels (SVAR). Afin de documenter les chocs de politique fiscale, l'approche narrative se base sur des informations institutionnelles. Par exemple, Ramey (2011) considère les périodes de guerre comme un indicateur pertinent des chocs exogènes et non anticipés de dépenses gouvernementales. Aussi, Romer et Romer (2010) considèrent les changements de législation fiscale, et isolent exclusivement ceux visant à stimuler la croissance à long terme afin d'identifier les chocs de taxes. Par ailleurs, les SVAR consistent à imposer des restrictions sur les relations contemporaines entre les variables du système afin d'isoler les chocs fiscaux. À titre d'exemple, il est couramment supposé dans la littérature que, au cours d'un trimestre, les dépenses gouvernementales ne réagissent pas aux changements de la production et de taxes (Fatás et Mihov, 2001a, Blanchard et Perotti, 2002, et Galí, López-Salido et Vallés, 2007). Dans ce contexte, les dépenses gouvernementales représentent l'indicateur de la politique fiscale, et les variations qui y sont associées correspondent à des chocs de dépenses gouvernementales. Afin de compléter l'identification, des restrictions additionnelles reliées au mécanisme de transmission des chocs fiscaux doivent être imposées. Par exemple, dans le cas d'un choc de dépenses gouvernementales, Fatás et Mihov (2001a) et Galí, López-Salido et Vallés (2007) considèrent que la production ne réagit pas de façon contemporaine aux changements des taxes.

Toutefois, les études empiriques n'apportent pas une réponse claire et unique quant aux effets des chocs de politique fiscale sur l'allocation des ressources domestiques. En effet, les études basées sur l'approche narrative confirment généralement les prédictions de la théorie classique. En revanche, les résultats obtenus à partir des SVAR sont fréquemment en concordance avec les prédictions de la théorie keynésienne. Cette ambiguïté, aussi bien sur le plan théorique qu'empirique, justifie la nécessité d'apporter plus de lumière sur ce sujet.

Le premier essai de cette thèse évalue empiriquement les effets des chocs de politique fiscale sur la production, la consommation et l'investissement pour le cas américain. Pour ce faire, nous estimons un SVAR qui relâche les hypothèses d'identification imposées dans les études empiriques existantes. Nous exploitons l'hétéroscédasticité conditionnelle des innovations associées aux variables du système afin d'identifier les chocs de taxes et de dépenses gouvernementales. De plus, la spécification que nous proposons emboîte les études basées sur l'approche des SVAR. Ceci nous permet de tester les restrictions couramment utilisées dans la littérature empirique qui sont associées aux différents indicateurs de la politique fiscale et aux mécanismes de transmission des chocs fiscaux.

Afin d'estimer les effets des chocs fiscaux sur l'activité économique américaine, nous considérons des données trimestrielles couvrant la période s'étalant de 1960 à 2007. En raison d'un bris structurel dans les données, l'échantillon est divisé en deux sous périodes. Le test proposé par Andrews et Ploberger (1994) suggère que la date de ce changement structurel correspond au deuxième trimestre de l'année 1979. En ce qui concerne la période pré-1979, nos résultats suggèrent que les indicateurs de la politique fiscale sont les dépenses gouvernementales ainsi que les taxes ajustées pour les changements cycliques. Par ailleurs, les données rejettent les restrictions impliquant que le déficit primaire est l'indicateur de la politique fiscale. De plus, les tests statistiques confirment la validité des restrictions reliées au mécanisme de transmission des chocs de dépenses imposées par le schéma d'identification récursif (e.g, Fatás et Mihov, 2001a et Galí, López-Salido et Vallés, 2007) et celui proposé par Blanchard et Perotti (2002). Enfin, les résultats soulignent que les restrictions imposées par Blanchard et Perotti (2002) pour documenter la transmission des chocs de taxes ne sont pas validées par les données.

Pour ce qui a trait à la période post-1979, les dépenses gouvernementales représentent le seul indicateur de la politique fiscale. Toutefois, les restrictions additionnelles imposées dans les études existantes afin de caractériser la transmission des chocs de dépenses sont statistiquement rejetées. Malgré le rejet des taxes ajustées pour les changements cycliques en tant qu'indicateur de la politique fiscale, le mécanisme de transmission des chocs de taxes proposé par Blanchard et Perotti (2002) est statistiquement valide.

Le rejet des restrictions liées à l'indicateur de la politique fiscale et/ou au mécanisme de transmission des chocs se traduit par des différences importantes entre les réponses dynamiques et les multiplicateurs documentés par notre spécification et ceux obtenus à partir des spécifications contraintes. Les réponses dynamiques non contraintes associées à un choc positif de dépenses gouvernementales indiquent une augmentation du PIB réel et de la consommation privée. Cette augmentation est plus persistante au cours de la période post-1979. Ce résultat est conforme avec les prédictions de la théorie keynésienne. Par ailleurs, un choc positif de taxes entraîne une baisse plus marquée du PIB réel et de la consommation privée durant la période pré-1979. Enfin, nos résultats montrent qu'en général la production et la consommation privée sont plus sensibles aux chocs de dépenses qu'aux chocs de taxes, alors que l'inverse est vrai pour l'investissement privé.

Ramey (2011) montre que les chocs de dépenses gouvernementales identifiés à partir des études basées sur des SVAR sont causés au sens de Granger par les périodes de guerre. Dans ce cas, les chocs sont anticipés et les réponses dynamiques qui en découlent sont biaisées. En utilisant le même test que Ramey (2011), nous montrons que les chocs de dépenses gouvernementales identifiés à partir de notre spécification ne sont pas sujets à cette critique. De plus, nous vérifions également que les chocs de taxes obtenus à partir de notre spécification ne sont pas causés au sens de Granger par les dates correspondant aux changements de législation fiscale identifiées par Romer et Romer (2010). Par conséquent, les réponses et les multiplicateurs dynamiques, associés aux chocs de taxes que nous identifions, ne sont pas biaisés.

Très peu d'études empiriques explorent les implications des chocs de politique fiscale sur l'ajustement extérieur des économies. À notre connaissance, Kim et Roubini (2008) sont les seuls à estimer l'effet des chocs de taxes sur le compte courant et le taux de change réel. Aussi, la réaction de ces variables aux chocs de dépenses gouvernementales a fait l'objet de seulement quelques travaux (Corsetti et Müller 2006; Kim et Roubini 2008; Müller 2008; Monacelli et Perotti 2010; Enders, Müller et Scholl 2011).

Basées sur des SVAR, ces études s'intéressent généralement au cas américain et indiquent une dépréciation réelle suite à une politique fiscale expansionniste. En ce qui a trait à la réponse du compte courant, Kim et Roubini (2008) réfutent l'hypothèse des déficits jumeaux, soit une relation positive entre les déficits extérieur et budgétaire. Spécifiquement, un choc négatif de taxes entraîne une amélioration du compte courant accompagnée d'une détérioration du solde budgétaire. Pour le cas d'une augmentation non anticipée des dépenses gouvernementales, la réponse du compte courant diffère d'un pays à l'autre. Somme toute, ces résultats sont en contradiction avec les enseignements théoriques. En effet, une large classe de modèles macroéconomiques de petites économies ouvertes prédisent qu'une politique fiscale expansionniste mène à une appréciation réelle de la devise et à une détérioration du compte courant.

Le deuxième essai de la thèse estime les effets des chocs fiscaux sur le taux de change réel et le compte courant pour le cas de quatre pays industrialisés, à savoir l'Australie, le Canada, le Royaume Uni et les États-Unis. Par rapport aux études existantes, notre travail se distingue dans trois dimensions. D'abord, relativement à Kim et Roubini (2008), nous étudions les effets des chocs de taxes sur le compte courant et le taux de change réel pour un plus grand nombre de pays. De plus, nous proposons une méthodologie qui relâche les restrictions d'identification imposées dans les études antérieures. Notre stratégie d'identification nous permet de vérifier les implications de ces restrictions sur les réponses dynamiques des variables aux chocs fiscaux. Dans la même veine que le premier essai, nous identifions les chocs fiscaux en exploitant les fluctuations temporelles des variances conditionnelles des innovations reliées aux variables du système.

Pour le cas américain, nos résultats indiquent qu'une baisse non anticipée des taxes entraîne une détérioration du compte courant américain et une augmentation du déficit budgétaire. Ceci suggère la validité de l'hypothèse des déficits jumeaux. De plus, les réponses dynamiques impliquent une appréciation réelle suite à un choc négatif de taxes. Il est important de signaler que ces résultats représentent des nouvelles évidences empiriques qui confirment les prédictions des modèles macroéconomiques standards. De plus, nous vérifions que les restrictions d'identification imposées par des études antérieures peuvent mener à des réponses dynamiques erronnées. Par exemple, les schémas d'identification proposés par Kim et Roubini (2008) ou Monacelli et Perotti (2010) impliquent qu'une baisse non anticipée des taxes détériore le solde budgétaire, améliore le compte courant et déprécie la valeur réelle de la devise. Enfin, et en ce qui concerne les effets d'une augmentation non anticipée des dépenses gouvernementales, les résultats indiquent une réponse positive (avec un certain retard) du compte courant et une détérioration du solde budgétaire, ce qui rejette la validité de l'hypothèse des déficits jumeaux. Pour ce qui a trait au taux de change réel, les résultats montrent une dépréciation plus prononcée que celle documentée dans les études antérieures.

Pour les autres pays de l'échantillon, nos résultats réfutent l'hypothèse des déficits jumeaux. En effet, le compte courant et le solde budgétaire réagissent différemment à une politique fiscale expansionniste. En ce qui concerne la réponse du taux de change réel, les résultats montrent une appréciation significative et très persistante pour le cas canadien suite à une baisse non anticipée des taxes. Une telle baisse n'a aucun effet sur le taux de change réel pour le cas de l'Australie et du Royaume-Uni. Finalement, une augmentation non anticipée des dépenses gouvernementales entraîne une dépréciation réelle significative pour le cas de l'Australie et du Royaume-Uni, alors que la réponse du taux de change canadien est nulle.

Le troisième essai de la thèse s'intéresse à la dynamique conjointe des déficits budgétaire et extérieur pour un échantillon de pays en développement. Ce travail met d'abord en évidence l'existence d'une covariance positive et significative entre les deux déficits. Ce constat est robuste aux différentes mesures du déficit extérieur, ainsi qu'à l'exclusion d'observations extrêmes. Le deuxième objectif du travail consiste à expliquer ce comouvement positif entre les déficits. Pour ce faire, nous développons un modèle de petite économie ouverte à générations imbriquées avec des biens de consommation hétérogènes. Ce modèle offre l'avantage de faire intervenir les conditions externes et internes qui sont des facteurs explicatifs fréquemment considérés pour les pays en développement. Les conditions externes sont le taux d'intérêt réel mondial, le taux de change réel et les termes de l'échange. Les conditions internes sont reliées à la production et aux variables fiscales. Aussi, le modèle capte les imperfections associées aux liens intergénérationnels et aux marchés financiers, telles que les contraintes de liquidité auxquelles les pays en développement font face.

Les paramètres du modèle sont estimés tels que les moments du second ordre prédits, et en particulier la covariance entre les déficits extérieur et budgétaire, s'approchent de leurs contreparties empiriques. La covariance prédite est ensuite décomposée afin d'identifier la contribution des chocs associés aux conditions internes et externes. Les résultats montrent que les chocs liés aux conditions internes représentent les facteurs les plus déterminants pour l'explication du comouvement positif entre les déficits extérieur et budgétaire. La prédominance des chocs associés aux conditions internes dépend positivement du degré d'imperfection des liens intergénérationnels ou des marchés financiers. Enfin, il est important de signaler que nos résultats contrastent avec ceux obtenus à partir des modèles à formes réduites analysant soit le comportement individuel du déficit extérieur ou du déficit budgétaire (Berg et Sachs 1988, Calderon, Chong et Zanforlin 2007, Calderon, Chong et Loayza 2002, Chinn et Prasad 2003, Combes et Saadi-Sedik 2006, Roubini 1991). En effet, cette littérature ne permet pas d'identifier les facteurs induisant un comouvement positif entre les déficits extérieur et budgétaire.

Essay 1. Measuring the Effects of Fiscal Policy

Abstract

Measuring the effects of discretionary fiscal policy is both difficult and controversial, as some explicit or implicit identifying assumptions need to be made to isolate exogenous and unanticipated changes in taxes and government spending. Studies based on structural vector autoregressions typically achieve identification by restricting the contemporaneous interaction of fiscal and non-fiscal variables in a rather arbitrary way. In this paper, we relax those restrictions and identify fiscal policy shocks by exploiting the conditional heteroscedasticity of the structural disturbances. We use this methodology to evaluate the macroeconomic effects of fiscal policy shocks in the U.S. before and after 1979. Our results show substantive differences in the economy's response to government spending and tax shocks across the two periods. Importantly, we find that increases in public spending are, in general, more effective than tax cuts in stimulating economic activity. A key contribution of this study is to provide a formal test of the identifying restrictions commonly used in the literature.

JEL classification: C32, E62, H20, H50, H60.

Keywords: Fiscal policy, Government spending, Taxes, Primary deficit, Structural vector autoregression, Identification.

1 Introduction

A classic question in macroeconomics is: how does fiscal policy affect economic activity and welfare? This question has received renewed interest in light of the recent financial crisis and the debate about the relevance and the nature of government intervention to stimulate the economy. To the extent that different theories provide different answers regarding the macroeconomic effects of fiscal policy, it is important to have an accurate empirical assessment of these effects. The purpose of this paper is to provide new evidence on this subject using an alternative empirical methodology that avoids potential shortcomings of existing approaches. The main challenge facing the empirical literature in this area is the difficulty to isolate exogenous and unanticipated changes in fiscal policy. One reason is that a large fraction of government revenue varies automatically with income and is, therefore, predictable. A second reason is that changes in public spending or taxes may reflect countercyclical policy actions to stabilize the economy or the government's desire to maintain the budget deficit or public debt at a given level.

The complexity of the process by which fiscal policy is conducted is not fully captured, however, in existing empirical studies that use structural vector auto-regressions (SVAR) to assess the effects of unanticipated shocks to government spending and taxes.[1] The assumptions commonly employed to identify these shocks are to a large extent arbitrary and sometimes overly restrictive, thus calling into question the validity of the ensuing results. For example, most existing studies identify government spending shocks by assuming that public spending is predetermined with respect to any other economic variable, including taxes (e.g., Fatás and Mihov, 2001a, Blanchard and Perotti, 2002, and Galí, López-Salido and Vallés, 2007). Also, following the seminal work of Blanchard and Perotti (2002), tax shocks are typically identified by purging the fraction of government revenue that changes automatically with output and by assuming that the resulting cyclically adjusted taxes do not respond to contemporaneous changes in government spending. In both cases, these exclusion restrictions–which define the policy indicator–are insufficient to achieve identification, and so additional restrictions must be imposed on

[1] A parallel empirical literature uses the narrative approach to identify exogenous and unanticipated changes in U.S. fiscal policy. Ramey and Shapiro (1998) isolate three events that led to large military buildups in the U.S. (the Korean War 1950:3, the Vietnam War 1965:1, and the Carter-Reagan defense build-up 1980:1). They identify exogenous changes in government spending with a dummy variable that traces these episodes. Ramey (2011) isolates more events that led the press to forecast increases in defense spending and provides estimates of the present value of the forecasted changes. Romer and Romer (2010) use a variety of government documents to identify, quantify and classify significant changes in federal tax legislation from 1947 to 2007.

the contemporaneous interaction of the variables included in the SVAR. These additional restrictions affect the transmission of fiscal policy shocks.

In this paper, we estimate the effects of fiscal policy shocks on GDP and domestic absorption in the U.S. using a flexible SVAR that relaxes the identifying assumptions used in previous studies. We instead achieve identification by exploiting the conditional heteroscedasticity of the innovations to the variables included in the SVAR, a methodology initially proposed by King, Sentana and Wadhwani (1994) and Sentana and Fiorentini (2001). The presence of conditional heteroscedasticity in the macroeconomic time series typically used in empirical work on fiscal policy has been documented by several existing studies.[2] Our empirical approach avoids imposing a priori assumptions about the implicit indicator of fiscal policy or its transmission mechanism, as it leaves unrestricted the contemporaneous interaction among fiscal instruments and between those instruments and the remaining variables of interest. Importantly, it also allows us to test various identifying restrictions commonly imposed in the literature, which are otherwise untestable under the usual assumption of conditional homoscedasticity of the shocks.[3] To the best of our knowledge, this is the first attempt to identify fiscal policy shocks and their effects through time-varying conditional variances.[4]

Underlying our empirical framework is a simple theoretical model that imposes a minimal structure on the system to be estimated. The model casts fiscal policy in the context of a market for newly issued government bonds. The supply of bonds may or may not shift as a result of changes in taxes or public expenditures, depending on the government's implicit target or, alternatively, fiscal-policy indicator. In turn, variations in taxes and public expenditures reflect both the automatic/systematic response of these variables to changes in economic conditions, and exogenous and unpredicted shifts in policy, i.e., fiscal-policy shocks. The market-clearing condition for bonds and the government budget

[2] See, for example, Garcia and Perron (1996), Den Haan and Spear (1998), Fountas and Karanasos (2007), Fernandez-Villaverde, Guerrón-Quintana, Kuester, and Rubio-Ramírez (2010), and Fernandez-Villaverde, Guerrón-Quintana, Rubio-Ramírez, and Uribe (2011).

[3] Mountford and Uhlig (2009) propose an alternative agnostic procedure whereby fiscal-policy shocks are identified by imposing sign restrictions on the impulse responses of fiscal variables and by assuming that these shocks are orthogonal to business-cycle and monetary-policy shocks. While Mountford and Uhlig's approach leaves unrestricted many of the contemporaneous relations between the variables of interest, it still restricts the response of fiscal variables to fiscal shocks and requires the prior identification of business-cycle and monetary-policy shocks. Moreover, the sign-restriction approach does not allow formal testing of the commonly used identifying restrictions.

[4] Identification through heteroscedasticity has been recently applied to study the effects of monetary policy shocks. Rigobon and Sachs (2004) assume that there is a shift in the *unconditional* variance of the monetary policy shock on days of FOMC meetings, while Normandin and Phaneuf (2004) and Bouakez and Normandin (2010) allow the *conditional* variances of policy and non-policy shocks to follow a parametric process.

constraint then impose a cross-equation restriction on the SVAR parameters, thus ensuring that the dynamics of fiscal variables are mutually consistent. An additional advantage of our theoretical model is that it allows us to give a structural interpretation to the parametric restrictions associated with the different indicators of fiscal policy.

In order to account for a structural break in the data, we estimate our SVAR over the pre- and post-1979 periods. While the model specification is not rejected in any of the two periods, estimates of the structural parameters differ substantially from one period to the other. These differences have important implications for the dynamic effects of fiscal policy shocks and their relative contribution to the variability of output. In particular, we find that an unexpected increase in government spending leads to a larger and more persistent rise in output in the post-1979 period than in the pre-1979 period. The implied impact multiplier (defined as the dollar change in output that results from a dollar increase in the exogenous component of public spending) increases from 0.93 in the former period to 1.34 in the latter. Our results also indicate that output has become less responsive to tax shocks after 1979 and that tax cuts are, in general, less effective in stimulating economic activity than increases in government spending.

The extent to which these results differ from those obtained by imposing the commonly used identifying restrictions depends on the nature of the fiscal shock and on the sample period. In the case of government spending shocks, the discrepancies between the unrestricted and restricted results are much more evident in the post-1979 period. For example, the spending multiplier implied by the unrestricted system for this period is roughly 50 percent larger than that implied by recursive identification schemes. Test results reveal that these discrepancies are not due to the restrictions associated with the policy indicator but rather to those associated with its propagation mechanism. More specifically, the hypothesis that government spending is predetermined with respect to taxes and output cannot be statistically rejected, whereas the additional restrictions typically imposed in the literature to complete identification are strongly rejected by the data. In the case of tax shocks, important differences between the restricted and unrestricted results exist in both periods, but they originate from different sources. Before 1979, these differences reflect the lack of empirical support for the restrictions associated with the transmission of tax policy. After 1979, they reflect the fact that cyclically adjusted taxes do not seem to be the appropriate indicator of tax policy.

A fundamental question that has received considerable attention in recent years concerns the response of private consumption to a government spending shock. Standard neoclassical theory predicts that public spending crowds out private consumption due to a negative wealth effect, but the empirical literature provides mixed evidence. Generally speaking, SVAR-based studies find that consumption rises in response to an increase in government spending (e.g., Blanchard and Perotti, 2002 and Galí, López-Salido and Vallés, 2007), while those based on the narrative approach find the opposite result (e.g., Edelberg, Eichenbaum, and Fisher,1999, Burnside, Eichenbaum, and Fisher, 2004 and Ramey, 2008). To shed further light on this issue, we estimate an extended version of our SVAR that includes consumption. We find a significant crowding-in effect of public spending, which has become substantially more persistent after 1979. We also find that the effects of government spending shocks on consumption are larger than those of tax shocks, whereas the opposite is true for private investment.

It is often argued that, due to the legislative and implementation lags inherent in fiscal policy, changes in government spending and taxes are likely to be anticipated by economic agents several months before they actually take place, a phenomenon commonly referred to as fiscal foresight (see, for example, Leeper, Walker and Yang, 2008). To the extent that agents behave in a forward-looking manner, reacting to news about future fiscal policy, the SVAR approach may fail to correctly identify fiscal policy shocks and may therefore lead to biased estimates of their effects. Ramey (2011) provides suggestive evidence that the SVAR-based innovations are in fact anticipated. More specifically, she finds that the government spending shocks extracted from a standard SVAR estimated using U.S. data and identified as in Blanchard and Perotti (2002) are Granger-caused by the war dates isolated by Ramey and Shapiro (1998).

To verify whether this criticism applies to the government spending shocks implied by our SVAR, we subject them to the same test carried out by Ramey. The test provides no evidence that these shocks are Granger-caused by the war dates. In fact, we find that Ramey's results are driven by the Korean-War episode, which is not covered by our sample period. We also conduct an analogous check for our tax shocks by testing whether they are Granger-caused by the dates identified by Romer and

Romer (2010) as marking the announcements of exogenous changes in U.S. tax policy. We again find no evidence that these dates predict the SVAR tax shocks. These results suggest that the fiscal-foresight problem is not sufficiently severe to undermine the ability of the SVAR approach to identify truly unanticipated shocks to fiscal policy, at least in the sample period considered here. This is likely due to the fact that an important fraction of fiscal policy shocks are in fact unanticipated. Simulation results by Mertens and Ravn (2009a) indeed show that if the data are generated both by anticipated and unanticipated fiscal shocks and that the former explain a relatively small share of the variance of fiscal variables, the SVAR approach can be successful in uncovering the true impulse responses to an unanticipated fiscal shock. These authors also estimate the effects of unanticipated government spending shocks in the U.S. using an augmented SVAR procedure that is robust to the presence of anticipated effects and find very similar results to those obtained from a standard SVAR.

The rest of the paper is organized as follows. Section 2 presents the SVAR specification and describes the identification strategy, the estimation method and the data. Section 3 reports the estimation results and discusses the properties of the implied fiscal policy indicators, the dynamic effects of fiscal policy shocks, and the implications of imposing the commonly used identifying restrictions. Section 4 extends the baseline SVAR to study the effects of fiscal policy shocks on consumption and investment. Section 5 concludes.

2 Empirical Methodology

2.1 Specification

We start with the following SVAR:

$$Az_t = \sum_{i=1}^{m} A_i z_{t-i} + \epsilon_t, \tag{1}$$

where z_t is a vector of macroeconomic variables and ϵ_t is a vector of mutually uncorrelated structural innovations, which include fiscal shocks. Blanchard and Perotti (2002) assume that the vector z_t consists of output, government spending and taxes. In our specification, we add to this list the price of government bonds for reasons that will become apparent below. Denote by ν_t the vector of residuals (or statistical innovations) obtained by projecting z_t on its own lags. These residuals are linked to the structural innovations through

$$A\nu_t = \epsilon_t, \tag{2}$$

where $A \equiv [a_{i,j}]_{i,j=1,\dots,4}$ is the matrix that captures the contemporaneous interaction among the variables included in z_t. Extracting the structural shocks from the residuals requires knowledge of the matrix A. As is well known, however, under conditional homoscedasticity of the structural shocks, projecting z_t on its own lags does not provide sufficient information to identify all the elements of A. As discussed below, our empirical methodology relaxes the assumption that the shocks are conditionally homoscedastic, so that no arbitrary restrictions need to be imposed to identify fiscal shocks and their effects. Nonetheless, we require that the elements of A satisfy a minimal set of cross-equation restrictions that ensure that system (1) is a coherent framework for the analysis of fiscal policy.

More specifically, we consider the following model:

$$\nu_{b,t}^{d} = -\alpha\nu_{q,t} + \beta(\nu_{y,t} - \nu_{\tau,t}) + \sigma_d\epsilon_{d,t}, \tag{3}$$

$$\nu_{p,t} \equiv \nu_{g,t} - \nu_{\tau,t} = \nu_{q,t} + \nu_{b,t}^{s}, \tag{4}$$

$$\nu_{g,t} = \eta_g\nu_{y,t} + \theta_g\sigma_d\epsilon_{d,t} + \psi_g\sigma_\tau\epsilon_{\tau,t} + \sigma_g\epsilon_{g,t}, \tag{5}$$

$$\nu_{\tau,t} = \eta_\tau\nu_{y,t} + \theta_\tau\sigma_d\epsilon_{d,t} + \psi_\tau\sigma_g\epsilon_{g,t} + \sigma_\tau\epsilon_{\tau,t}. \tag{6}$$

Equation (3) is the private sector's demand for newly issued government bonds (Treasury bills), expressed in innovation form. It states that the demand for bonds, $\nu_{b,t}^{d}$, depends on the price of bonds, $\nu_{q,t}$, on disposable income, $\nu_{y,t} - \nu_{\tau,t}$, and on a demand shock, $\epsilon_{d,t}$, scaled by the parameter σ_d. The parameter α, which measures (the absolute value of) the slope of the demand curve, is assumed to be positive and different from 1, and β is a positive parameter. Rather than taking a stand on the process by which the government determines the quantity of newly issued bonds, we simply require that this quantity satisfies the (linearly approximated) government's budget constraint. The latter is given by equation (4), which states that the innovation in the primary deficit, $\nu_{p,t}$, (i.e., the difference between government spending and taxes) must be equal to the innovation in the value of debt. Note that because this constraint is expressed in innovation form, it does not include the payment for bonds that mature in period t (since those bonds were issued in period $t-1$).[5] Equations (5) and (6) describe the procedures followed by the government to determine fiscal spending and taxes. The disturbances $\epsilon_{g,t}$ and $\epsilon_{\tau,t}$ are the fiscal shocks that we aim to identify. The former is a shock to government spending and the latter is a tax shock. The terms σ_g and σ_τ are scaling parameters. Equation (5) states that government spending may change in response to changes in output or to demand and tax shocks. Equation(6) has an analogous interpretation for taxes. In these equations, the parameters η_g and η_τ measure the automatic and systematic responses of, respectively, government spending and taxes to changes in output. In this respect, η_g and η_τ do not necessarily coincide with the elasticities of fiscal variables with respect to output estimated by Blanchard and Perotti (2002), which capture only the automatic adjustment of government spending and taxes. As we explain below, different procedures to set fiscal policy will be characterized by different values of the parameters $\beta, \eta_g, \eta_\tau, \theta_g, \theta_\tau, \psi_g$ and ψ_τ.

Imposing equilibrium in the bonds market and solving for the structural innovations, ϵ_t, in terms of the residuals, ν_t, yield

$$\begin{pmatrix} a_{11} & a_{12} & a_{13} & a_{14} \\ -\frac{\beta}{\sigma_d} & \frac{\alpha-1}{\sigma_d} & \frac{1}{\sigma_d} & \frac{\beta-1}{\sigma_d} \\ \frac{\psi_g(\eta_\tau-\beta\theta_\tau)-\left(\eta_g-\beta\theta_g\right)}{\sigma_g\left(1-\psi_g\psi_\tau\right)} & \frac{(1-\alpha)\left(\theta_g-\theta_\tau\psi_g\right)}{\sigma_g\left(1-\psi_g\psi_\tau\right)} & \frac{1-\theta_g+\theta_\tau\psi_g}{\sigma_g\left(1-\psi_g\psi_\tau\right)} & \frac{(1-\beta)\left(\theta_g-\theta_\tau\psi_g\right)-\psi_g}{\sigma_g\left(1-\psi_g\psi_\tau\right)} \\ \frac{\psi_\tau\left(\eta_g-\beta\theta_g\right)-(\eta_\tau-\beta\theta_\tau)}{\sigma_\tau\left(1-\psi_g\psi_\tau\right)} & \frac{(1-\alpha)(\theta_\tau-\theta_g\psi_\tau)}{\sigma_\tau\left(1-\psi_g\psi_\tau\right)} & \frac{\psi_\tau(\theta_g-1)-\theta_\tau}{\sigma_\tau\left(1-\psi_g\psi_\tau\right)} & \frac{1+(1-\beta)(\theta_\tau-\theta_g\psi_\tau)}{\sigma_\tau\left(1-\psi_g\psi_\tau\right)} \end{pmatrix}$$
$$\times \begin{pmatrix} v_{y,t} \\ v_{q,t} \\ v_{g,t} \\ v_{\tau,t} \end{pmatrix} = \begin{pmatrix} \epsilon_{1,t} \\ \epsilon_{d,t} \\ \epsilon_{g,t} \\ \epsilon_{\tau,t} \end{pmatrix}, \tag{7}$$

where a_{1j} $(j = 1, ..., 4)$ are unconstrained parameters.

The conditional scedastic structure of system (7) is:

$$\Sigma_t = A^{-1}\Gamma_t A^{-1'}, \tag{8}$$

where $\Sigma_t = E_{t-1}(\nu_t\nu_t')$ is the (non-diagonal) conditional covariance matrix of the statistical innovations and $\Gamma_t = E_{t-1}(\epsilon_t\epsilon_t')$ is the (diagonal) conditional covariance matrix of the structural innovations. The

[5] The government budget constraint (4) omits seignorage, given that this source of revenue has historically been negligible in the U.S. during the period considered (less than 0.4 percent of GDP on average, according to our calculations).

unconditional variances of the structural innovations are normalized to unity ($I = E(\epsilon_t \epsilon_t')$). The dynamics of the conditional variances of the structural innovations are determined by

$$\Gamma_t = (I - \Delta_1 - \Delta_2) + \Delta_1 \bullet (\epsilon_{t-1}\epsilon_{t-1}') + \Delta_2 \bullet \Gamma_{t-1}. \tag{9}$$

The operator $\bullet$ denotes the element-by-element matrix multiplication, while Δ_1 and Δ_2 are diagonal matrices of parameters. Equation (9) involves intercepts that are consistent with the normalization $I = E(\epsilon_t \epsilon_t')$. Also, (9) implies that all the structural innovations are conditionally homoscedastic if Δ_1 and Δ_2 are null. On the other hand, some structural innovations display time-varying conditional variances characterized by univariate generalized autoregressive conditional heteroscedastic [GARCH(1,1)] processes if Δ_1 and Δ_2 — which contain the ARCH and GARCH coefficients, respectively — are positive semi-definite and $(I - \Delta_1 - \Delta_2)$ is positive definite. Finally, all the conditional variances follow GARCH(1,1) processes if Δ_1, Δ_2, and $(I - \Delta_1 - \Delta_2)$ are positive definite.

2.2 Identification

Under conditional heteroscedasticity, system (7) can be identified, allowing us to study the effects of fiscal policy shocks. The sufficient (rank) condition for identification states that the conditional variances of the structural innovations are linearly independent. That is, $\lambda = 0$ is the only solution to $\Gamma\lambda = 0$, such that $(\Gamma'\Gamma)$ is invertible — where Γ stacks by column the conditional volatilities associated with each structural innovation. The necessary (order) condition requires that the conditional variances of (at least) all but one structural innovations are time-varying. In practice, the rank and order conditions lead to similar conclusions, given that the conditional variances are parameterized by GARCH(1,1) processes (see Sentana and Fiorentini, 2001).

To understand how time-varying conditional volatility helps with identification, first note that the unconditional variances of the statistical and structural shocks are related through

$$\Sigma = A^{-1} A^{-1'}. \tag{10}$$

Assuming the SVAR includes n variables, the estimate of Σ allows to identify $\frac{n(n+1)}{2}$ of the n^2 elements of A, leaving $\frac{n(n-1)}{2}$ elements to be identified. Note also that (8) implies

$$\Delta\Sigma_t = \Sigma_t - \Sigma_{t-1} = A^{-1}\left[\Gamma_t - \Gamma_{t-1}\right]A^{-1'} = A^{-1}\left[\Delta\Gamma_t\right]A^{-1'}. \tag{11}$$

This set of equations allows to identify $\frac{k(k+1)}{2}$ additional parameters of A, where k is the rank of $\Delta\Gamma_t$. Hence, if $\Delta\Gamma_t$ has a rank of at least $n-1$, identification can be achieved. In our context, a necessary condition for this is that at least $n-1$ structural innovation are time-varying.

Under conditional homoscedasticity of the structural disturbances (i.e., when Δ_1 and Δ_2 are null), (8) and (10) coincide, so that (11) becomes non-informative. In this case, $\frac{n(n-1)}{2}$ arbitrary restrictions need to be imposed on the elements of A in order to achieve identification. These restrictions reflect the econometrician's belief about the relevant policy indicator and/or transmission mechanism of fiscal shocks. We now show how our empirical model nests various commonly used restrictions of both types.[6]

2.2.1 Restrictions Associated with the Policy Indicator

The third equation of system (7) shows how the government spending shock is related to the VAR residuals:

$$\epsilon_{g,t} = a_{31}\nu_{y,t} + a_{32}\nu_{q,t} + a_{33}\nu_{g,t} + a_{34}\nu_{\tau,t}, \tag{12}$$

where

$$\begin{aligned}
a_{31} &= \frac{\psi_g\left(\eta_\tau - \beta\theta_\tau\right) - \left(\eta_g - \beta\theta_g\right)}{\sigma_g\left(1-\psi_g\psi_\tau\right)},\\
a_{32} &= \frac{(1-\alpha)\left(\theta_g - \theta_\tau\psi_g\right)}{\sigma_g\left(1-\psi_g\psi_\tau\right)},\\
a_{33} &= \frac{1-\theta_g+\theta_\tau\psi_g}{\sigma_g\left(1-\psi_g\psi_\tau\right)},\\
a_{34} &= \frac{(1-\beta)\left(\theta_g - \theta_\tau\psi_g\right) - \psi_g}{\sigma_g\left(1-\psi_g\psi_\tau\right)}.
\end{aligned}$$

[6] Mountford and Uhlig (2009) propose an alternative identification strategy by imposing sign restrictions on the variables' responses. Their system, however, includes more variables than ours so that their identifying restrictions cannot be nested in our framework. We are therefore unable to test those restrictions or to reproduce their results using our set of variables.

The term on the right-hand side of equation (12) defines the fiscal-spending indicator (in innovation form). Since the coefficients a_{3j} $(j = 1, ..., 4)$ are functions of freely estimated parameters, this policy indicator is not constrained to be summarized by a single variable (or a particular subset of variables). This contrasts with existing empirical studies, which make a priori assumptions about the relevant policy indicator in order to achieve identification. Most of these studies assume that the fiscal-spending indicator is government spending (Blanchard and Perotti, 2002, Galì, López-Salido and Vallès, 2007). Fatàs and Mihov (2001b), on the other hand, use the primary deficit as a broad indicator of fiscal policy (i.e., without distinction between government spending and tax policies). The parametric restrictions under which government spending and the primary deficit measure the stance of fiscal spending are the following:

- *G indicator* (*government spending*): $\eta_g = \theta_g = \psi_g = 0$. In this case, changes in government spending are completely predetermined with respect to the current state of the economy and do not reflect any systematic/automatic response of the government. It is easy to show that under these restrictions the policy shock is proportional to the innovation to government spending ($\epsilon_{g,t} = \frac{1}{\sigma_g}\nu_{g,t}$).

- *PD indicator* (*primary deficit*): $\eta_g = \eta_\tau$, $\theta_g = \theta_\tau$ and $\psi_g = 1$. Under this scenario, the government targets the primary deficit when setting fiscal spending. Unexpected changes in the primary deficit therefore reflect purely government spending shocks ($\epsilon_{g,t} = \frac{1}{(1-\psi_\tau)\sigma_g}\nu_{p,t}$).

Analogously, the fourth equation of system (7) is

$$\epsilon_{\tau,t} = a_{41}\nu_{y,t} + a_{42}\nu_{q,t} + a_{43}\nu_{g,t} + a_{44}\nu_{\tau,t}, \tag{13}$$

where

$$\begin{aligned}
a_{41} &= \frac{\psi_\tau\left(\eta_g - \beta\theta_g\right) - \left(\eta_\tau - \beta\theta_\tau\right)}{\sigma_\tau\left(1 - \psi_g\psi_\tau\right)}, \\
a_{42} &= \frac{(1-\alpha)\left(\theta_\tau - \theta_g\psi_\tau\right)}{\sigma_\tau\left(1 - \psi_g\psi_\tau\right)}, \\
a_{43} &= \frac{\psi_\tau\left(\theta_g - 1\right) - \theta_\tau}{\sigma_\tau\left(1 - \psi_g\psi_\tau\right)}, \\
a_{44} &= \frac{1 + (1-\beta)\left(\theta_\tau - \theta_g\psi_\tau\right)}{\sigma_\tau\left(1 - \psi_g\psi_\tau\right)}.
\end{aligned}$$

Two cases of interest are nested in the rule above. The first defines the relevant indicator of tax policy as cyclically adjusted government revenue, as in Blanchard and Perotti (2002). In the second, the tax-policy indicator is the primary deficit. The corresponding restrictions are:

- *CAT indicator* (*cyclically adjusted taxes*): $\theta_\tau = \psi_\tau = 0$. In this case, tax shocks are measured with unexpected changes in the fraction of government revenue that does not vary automatically or systematically with output ($\epsilon_{\tau,t} = \frac{\nu_{\tau,t} - \eta_\tau \nu_{y,t}}{\sigma_\tau}$).

- *PD indicator* (*primary deficit*): $\eta_g = \eta_\tau$, $\theta_g = \theta_\tau$ and $\psi_\tau = 1$. In this case, tax shocks correspond to unexpected changes in the primary deficit ($\epsilon_{\tau,t} = \frac{1}{(\psi_g - 1)\sigma_\tau} \nu_{p,t}$).

2.2.2 Restrictions Associated with the Transmission Mechanism

Each of the policy indicators discussed in the previous section implies 3 different restrictions on the elements of A (2 in the case of the CAT indicator). Therefore, 3 additional restrictions (4 in the case of the CAT indicator) have to be imposed in order to achieve identification. These restrictions in turn determine the way in which fiscal shocks affect the endogenous variables over time. In the case of a government spending shock, the literature typically completes identification via a Cholesky decomposition of the covariance matrix of the VAR residuals, which yields three additional zero restrictions (see, for example, Fatás and Mihov, 2001a and Galí, López-Salido and Vallés, 2007). By ordering government spending first among the variables included in the VAR, this scheme implies that the matrix A in (1) is lower triangular, so that system (7) becomes

$$\begin{pmatrix} \tilde{a}_{11} & 0 & 0 & 0 \\ \tilde{a}_{21} & \tilde{a}_{22} & 0 & 0 \\ \tilde{a}_{31} & \tilde{a}_{32} & \tilde{a}_{33} & 0 \\ \tilde{a}_{41} & \tilde{a}_{42} & \tilde{a}_{43} & \tilde{a}_{44} \end{pmatrix} \begin{pmatrix} v_{g,t} \\ v_{y,t} \\ v_{\tau,t} \\ v_{q,t} \end{pmatrix} = \begin{pmatrix} \epsilon_{g,t} \\ \epsilon_{1,t} \\ \epsilon_{\tau,t} \\ \epsilon_{d,t} \end{pmatrix}. \tag{14}$$

This identification scheme can be obtained as a special case of system (7) by imposing the following restrictions: $a_{12} = a_{14} = \theta_\tau = 0$, in addition to three restrictions associated with the G indicator. Note that the ordering of the remaining variables is irrelevant when computing the effects of a government spending shock.

Blanchard and Perotti (2002) propose an alternative, non-recursive, scheme to identify the effects of a government spending shock. In the context of our four-variable SVAR, their identification scheme implies

$$\begin{pmatrix} \tilde{a}_{11} & 0 & 0 & 0 \\ \tilde{a}_{21} & \tilde{a}_{22} & \tilde{a}_{23} & 0 \\ \tilde{a}_{31} & -x\tilde{a}_{33} & \tilde{a}_{33} & 0 \\ \tilde{a}_{41} & \tilde{a}_{42} & \tilde{a}_{43} & \tilde{a}_{44} \end{pmatrix} \begin{pmatrix} v_{g,t} \\ v_{y,t} \\ v_{\tau,t} \\ v_{q,t} \end{pmatrix} = \begin{pmatrix} \epsilon_{g,t} \\ \epsilon_{1,t} \\ \epsilon_{\tau,t} \\ \epsilon_{d,t} \end{pmatrix}, \tag{15}$$

where x is the elasticity of taxes with respect to output, which is estimated outside the SVAR. The system above can be obtained by setting $a_{12} = \theta_\tau = 0$ and $\eta_\tau = x$ in (7), in addition to the restrictions associated with the G indicator. It is worth emphasizing that the recursive and non-recursive schemes given by (14) and (15) yield identical responses to a government spending shock since they imply identical estimates of the parameters $\tilde{a}_{i1}$ $(i = 1, ..., 4)$.

To identify the effects of a tax shock, Blanchard and Perotti relax the assumption that $\tilde{a}_{12} = 0$ and assume instead that taxes are predetermined with respect to government spending. This yields

$$\begin{pmatrix} \tilde{a}_{11} & \tilde{a}_{12} & -\tilde{a}_{12}/x & 0 \\ \tilde{a}_{21} & \tilde{a}_{22} & \tilde{a}_{23} & 0 \\ 0 & -x\tilde{a}_{33} & \tilde{a}_{33} & 0 \\ \tilde{a}_{41} & \tilde{a}_{42} & \tilde{a}_{43} & \tilde{a}_{44} \end{pmatrix} \begin{pmatrix} v_{g,t} \\ v_{y,t} \\ v_{\tau,t} \\ v_{q,t} \end{pmatrix} = \begin{pmatrix} \epsilon_{g,t} \\ \epsilon_{1,t} \\ \epsilon_{\tau,t} \\ \epsilon_{d,t} \end{pmatrix}. \tag{16}$$

In this specification, the precise value of x imposed by Blanchard and Perotti captures exclusively the automatic adjustment of taxes to output.[7] This system can be obtained from (7) by imposing $a_{12} = \theta_g = \eta_g = 0$ and $\eta_\tau = x$, in addition to the two restrictions associated with the CAT indicator.

Under conditional homoscedasticity, none of these identifying restrictions can be tested; thus, no formal criterion can be used to choose among competing identification schemes. This is possible,

[7] Note that the first and third equations of (16) can be rewritten as

$$\begin{aligned} v_{g,t} &= \frac{\tilde{a}_{12}}{x\tilde{a}_{11}\tilde{a}_{33}}\epsilon_{\tau,t} + \frac{1}{\tilde{a}_{11}}\epsilon_{g,t}, \\ v_{\tau,t} &= xv_{y,t} + \frac{1}{\tilde{a}_{33}}\epsilon_{\tau,t}. \end{aligned}$$

This representation is similar to that found in Blanchard and Perotti (2002, p. 1333).

however, under our identification method (which exploits the conditional heteroscedasticity of the shocks) since it leaves unrestricted the elements of A. We perform this exercise in Sections 3.2.1 and 3.3.3.

2.3 Estimation Method and Data

The elements of A, Δ_1, and Δ_2 are estimated using the following two-step procedure. We first estimate by ordinary least squares a $4-$order VAR ($m = 4$) that includes output, the price of bonds, government spending and taxes,[8] and extract the implied residuals, ν_t, for $t = m+1, ..., T$. For given values of the elements of the matrices A, Δ_1, and Δ_2, it is then possible to construct an estimate of the conditional covariance matrix Σ_t recursively, using equations (8) and (9) and the initialization $\Gamma_m = \epsilon_m \epsilon_m' = I$. Assuming that the residuals are conditionally normally distributed, the second step consists in selecting the elements of the matrices A, Δ_1, and Δ_2 that maximize the likelihood of the sample.

We use quarterly U.S. data from 1960:1 to 2007:4. In their main analysis, Blanchard and Perotti (2002) excluded the 1950s on the ground that this period was characterized by exceptionally large spending and tax shocks. Since one of our objectives is to compare our results to theirs, we restrict our sample to the post-1960 period, and we closely follow their approach in constructing the series used in estimation. Output is measured by real GDP. The price of bonds is measured by the inverse of the gross real return on 3-month treasury bills,[9] where the CPI is used to deflate the gross nominal return. Government spending is defined as the sum of federal (defense and non-defense), state and local consumption and gross investment expenditures. Taxes are defined as total government receipts less net transfer payments. The spending and tax series are expressed in real terms using the GDP deflator. The data are taken from the National Income and Products Accounts (NIPA), except for the 3-month treasury bill rate, which is obtained from the Federal Reserve Bank of Saint-Louis' Fred database.[10] Output, government spending and taxes are divided by total population (taken from Fred) and all the series are expressed in logarithm.

The transformed series are depicted in Figure 1.1. The series of output, government spending and taxes exhibit a clear upward trend, but that of the price of bonds appears to have two distinct

[8] A constant and a trend are also included among the regressors.
[9] We found the results to be robust when we measure the price of bonds using the return on 10-year treasury bonds.
[10] All the series, except the interest rate, are seasonnally adjusted at the source.

regimes separated by a break around the end of the 1970s. This observation suggests that it may not be appropriate to estimate (7) over the entire sample period. To determine the cutoff date in a more formal and precise way, we applied Andrews and Ploberger's (1994) structural break test to detect changes in the trend of the price of bonds. The test suggests that there is a break at 1979:2. We therefore consider the two sub-periods: 1960:1–1979:2 and 1979:3–2007:4.[11]

3 Results

3.1 Estimation and Test Results

Before discussing the estimates of the structural parameters and their implications, we perform a preliminary analysis to document the presence of conditional heteroscedasticity in the series used in estimation. We start by applying the multivariate ARCH test proposed by Fiorentini and Sentana (2009) to the statistical residuals obtained from the VAR. The test is applied both to the diagonal and off-diagonal elements of the matrices of the ARCH coefficients at a given lag. The test results presented in Table 1.I confirm the presence of cross-correlation in the squared statistical residuals at different lags. While the exact form of the process that governs the conditional covariances of the residuals is not exactly an ARCH(q) process but instead a non-linear combination of GARCH(1,1) processes, the rejection of the null that the ARCH coefficients are jointly insignificant nonetheless suggests the presence of heteroscedasticity in the statistical innovations, which is likely to reflect time varying conditional variances of the structural shocks.

A visual inspection of the conditional variances of the structural shocks extracted from system (7), depicted in Figure 1.1, reveals that both fiscal and non-fiscal shocks exhibit significant heteroscedasticity in both sub-samples, displaying alternating episodes of high and low volatility. This suggests that the order condition for identification (that at least three shocks have time varying conditional variances) is satisfied. This observation corroborates the findings of earlier studies that document the presence of conditional volatility in the time series of output (Fountas and Karanasos 2007), the interest rate (Garcia and Perron 1996; Den Haan and Spear 1998; Fernandèz-Villaverde et al. 2010), and fiscal variables (Fernandèz-Villaverde et al. 2011).

[11] Perotti (2004) and Favero and Giavazzi (2009) also distinguish between the pre- and post-1980 periods when measuring the effects of U.S. fiscal policy.

Unfortunately, a test of the significance of the ARCH and GARCH coefficients in (9) cannot be performed because conventional critical values are invalid under the null hypothesis of conditional homoscedasticity, given that system (7) becomes under-identified. However, it is possible to apply the multivariate ARCH test proposed by Fiorentini and Sentana (2009) using an identified version of system (7), i.e., a version in which enough restrictions are imposed on the matrix A to ensure identification.[12] Since the structural shocks are orthogonal, the test is applied only to the diagonal elements of the matrices of the ARCH coefficients at a given lag. Again, strictly speaking, this is not a test of the joint significance of the parameters in (9); however, as is well known, the latter process can be approximated by an ARCH process of a sufficiently high order. The test results, presented in Table 1.II, indicate that the null hypothesis that the ARCH coefficients are jointly equal to 0 at lags 1, 2, and 4 is generally rejected by the data, implying that the conditional variances of the structural innovations are time-varying.

To determine whether the GARCH(1,1) specification provides an adequate description of the process that governs the conditional variances of the structural innovations, we test whether there is any autocorrelation in the ratio of the squared structural innovations relative to their conditional variances. The McLeod-Li test results, reported in Table 1.III, indicate that the null hypothesis of no autocorrelation cannot be rejected at the 5 percent (and, with one exception, at the 10 percent) significance level for 1, 2 and 4 lags. This suggests that the GARCH(1,1) process is well specified.

Table 1.IV reports estimates of the structural parameters. With a few exceptions, the estimates differ substantially across the two periods, thus confirming the presence of instability and justifying the need to focus on sub-periods rather than the entire sample period. Recall that the model presented in Section 2 imposes restrictions only on the parameters α and β. The former has to be positive and different from 1, while the latter must take a positive value. The requirements for α are satisfied in both periods, but we obtain a positive estimate of β only for the post-1979 period. None of the point estimates is precise, however. Our model also implies that the linear restriction $a_{21} + a_{23} = -a_{24}$ must hold. A likelihood-ratio test of this restriction indicates that it cannot be rejected at standard significance levels in any of the two sub-periods (see Table 1.V). This suggests that system (7) represents

[12] An advantage of Fiorentini and Sentana's test is that the numerical value of the test statistic is invariant to the specific identifying restrictions imposed on A.

an adequate specification of the data, so we henceforth refer to it as the unrestricted system and to its implications as the unrestricted ones. As stated above, this system can be used to test the identifying restrictions associated with the various indicators of fiscal policy and with the different transmission mechanisms usually assumed in the literature. In this regard, it is important to emphasize that despite the uncertainty surrounding the individual estimates of structural parameters, we show below that we obtain sharp implications regarding the effects of fiscal policy shocks and that we are often able to reject the joint restrictions associated with the commonly used identifying assumptions.

3.2 Fiscal Policy Indicators

Our empirical approach implies that the appropriate indicator of fiscal policy is an unrestricted linear combination of output, the price of bonds, government spending and taxes. On the other hand, each of the commonly used indicators discussed above implies a set of parametric restrictions that we formally test below. We also graphically compare the time series of fiscal policy shocks implied by the unrestricted system and those associated with the restricted indicators. Finally, we discuss the predictability of our measured fiscal policy shocks.

3.2.1 Tests of the Restricted Indicators

The restrictions associated with each of the fiscal policy indicators are tested using a likelihood-ratio test. Panel A of Table 1.VI reports the results. Starting with spending policy, the results indicate that the restrictions associated with the G indicator cannot be rejected, whereas those associated with the PD indicator are clearly not supported by the data, especially in the 1979:3–2007:4 period. The former result corroborates the conclusion reached by Blanchard and Perotti (2002) based on institutional information that there is little evidence of a contemporaneous response of government spending to economic activity. In turn, this also suggests that the commonly used identifying assumption that innovations to government spending are exogenous is a plausible one.

As for tax policy, we find that the restrictions associated with the CAT indicator are consistent with the data only in the 1960:1–1979:2 period, but that they are soundly rejected in the post-1979 period. Thus, purging the automatic/systematic response of taxes to economic activity is not sufficient to isolate the purely exogenous component of tax changes, at least when focusing on more recent data. The message that the unrestricted model conveys is that one also needs to purge the systematic response of taxes to government spending and demand shocks.

3.2.2 Unrestricted Versus Restricted Measures of Fiscal Policy Shocks

Using the estimates of the elements of A and the statistical innovations extracted in the first step of our estimation procedure, it is straightforward to recover (via equation 2) the time series of structural shocks and, in particular, fiscal policy shocks, implied by the unrestricted system and each of the restricted policy indicators discussed above. Figure 1.3 depicts the unrestricted and restricted series of government spending shocks. Figure 1.4 shows the series of tax shocks. Table 1.VII reports the correlation coefficients between the unrestricted and restricted measures of fiscal policy shocks.

Figure 1.3 shows that the time series of government spending shocks obtained under the restrictions associated with the G indicator tracks very closely the unrestricted measure of shocks in each of the two sub-samples. The correlation between the two series is 0.99 in the first sub-period and 0.97 in the second (see Table 1.VII). On the other hand, the time series of shocks obtained under the restrictions associated with the PD indicator are weakly correlated with their unrestricted counterparts, especially in the post-1979 period. This weak correlation reflects frequent and sometimes important gaps with respect to the valid measures of government spending shocks. In particular, imposing the restrictions associated with the PD indicator would lead the econometrician to substantially underestimate the unexpected increase in public spending that occurred during the Vietnam-War period (mid-1960s) and to completely miss the one that followed September 11, 2001. These results are consistent with the test results discussed in the previous section and confirm that the primary deficit is a poor indicator of fiscal spending.

Regarding tax shocks, Figure 1.4 reveals that the restrictions associated with the CAT indicator do not occasion any major mis-measurement of tax innovations in the pre-1979 period: the correlation between the restricted and unrestricted series of innovations is 0.98 in this sub-period (see Table 1.VII). In the post-1979 period, however, these restrictions entail some important counterfactual implications, which explain their statistical rejection discussed in the previous section. For example, under these restrictions, one would mistakenly conclude that there were substantial exogenous tax cuts in 1994 and tax increases in 1999. The restrictions associated with the PD indicator, for their part, generate a measure of tax shocks that deviates markedly from the unrestricted one in both sub-samples, although the fit is much worse in the post-1979 period. This again confirms that the primary deficit is not an appropriate indicator of tax policy.

3.2.3 Are Fiscal Policy Shocks Anticipated ?

The SVAR approach has often been criticized on the ground that it may not be robust to fiscal foresight, i.e., the phenomenon that, due to legislative and implementation lags, future changes in fiscal policy are signaled to economic agents several months before they become effective.[13] To the extent that agents adjust their behavior in response to anticipated shocks, the resulting time series may have a non-invertible moving average component, such that it would be impossible to recover the true fiscal shocks from current and past variables.[14] Ramey (2011) presents suggestive evidence that the SVAR-based innovations miss the timing of the news and are in fact predictable. More specifically, she shows that the government spending shocks extracted from a standard SVAR (identified via a Cholesky decomposition) are Granger-caused by Ramey and Shapiro's (1998) war dates.[15]

In order to investigate whether this criticism also applies to our government spending shocks, we subject them to the Granger causality test performed by Ramey. More precisely, we regress government spending shocks on four lags of a dummy variable that represents the war dates, and test the joint significance of the regression coefficients. The results are reported in Panel A of Table 1.VIII. They indicate that the Ramey-Shapiro dates do not Granger-cause the (unrestricted) government spending shocks in any of the two periods.[16] Even when we consider the shocks implied by the restricted policy indicators of fiscal spending, we strongly reject the hypothesis that they are Granger-caused by the war dates.[17] We conclude that the SVAR government spending shocks correctly capture unexpected

[13] Leeper, Walker and Yang (2008) review the literature that reports reduced-form and anecdotal evidence on the extent of fiscal foresight.

[14] See Sims (1988), Hansen and Sargent (1991), Yang (2005), and Leeper, Walker and Yang (2008).

[15] Ramey (2011) adds 2001:3 to three episodes previously identified by Ramey and Shapiro (1950:3, 1965:1, 1980:1).

[16] This result is robust to using 1, 2 or 3 lags. We also considered the new military dates isolated by Ramey (2011) based on her reading of *Business Week* and the *New York Times*, and found no evidence that these dates Granger-cause the SVAR government spending shocks.

[17] The only exception occurs in the case of the PD indicator, for which government spending shocks are Granger-caused by the war dates in the 1979:3–2007:4 period.

changes in public expenditures. One might suspect that this is the case because the effects of fiscal foresight are being impounded into the price of bonds and by conditioning on this variable, we are able to capture the true conditioning set of agents. But we reach the same conclusion when we exclude the price of bonds from the system. This and the fact that the absence of Granger causality holds across several identification schemes suggest that Ramey's findings are most likely driven by the Korean-War episode, which is not covered by our sample period.[18]

To undertake an analogous check for tax shocks, we use the dates isolated by Romer and Romer (2010) to identify "exogenous" changes in tax policy based on presidential speeches and Congressional reports. In Romer and Romer's terminology, these exogenous changes correspond to legislated tax policy actions that are not taken for the purpose of offsetting factors that could affect output growth. Panel B of Table 1.VIII reports Granger-causality results for the SVAR tax shocks. These results clearly show that Romer and Romer's dates do not Granger-cause the SVAR tax shocks, irrespective of the time period and the specification.[19] This means that these shocks are not forecastable based on the dates of legislated exogenous tax changes.

Together, these findings suggest that the fiscal-foresight problem is not sufficiently severe to hinder the ability of the SVAR approach to correctly identify unanticipated fiscal policy shocks, at least conditional on the data used in this paper.[20] This could be due to the fact that economic agents do not behave in a forward-looking manner, either because they are myopic or because they are prevented from doing so (due, for example, to liquidity constraints). A more plausible explanation, however, is that an important fraction of fiscal policy shocks are in fact unanticipated.[21] A recent study by Mertens and Ravn (2009b) lends support to this conjecture. Using artificial data generated by a neoclassical model with anticipated and unanticipated fiscal shocks, these authors show that the SVAR approach

[18] We verified that once we extend the sample period to include the Korean War (that is, when we consider the period 1947:1–2007:4), we recover Ramey's result that the war dates Granger-cause the SVAR-based government spending shocks.

[19] There are only two dates for which Romer and Romer report simultaneously tax changes taken for exogenous and endogenous reasons. Excluding these two dates does not alter the outcome of our Granger-causality test.

[20] Perotti (2004) also finds little evidence that the SVAR fiscal innovations are predictable in a sample of 5 OECD countries. More specifically, he shows that these innovations are, in general, uncorrelated with the OECD forecasts of government spending and GDP growth.

[21] As emphasized by Perotti (2004), throughout a given fiscal year, there are often supplements to the Budget and other decisions by the governments that affect the outcome of fiscal policy. Moreover, Mertens and Ravn (2009b) point out that of the 70 changes in the tax bill identified by Romer and Romer (2008) as being exogenous, 32 took effect within 90 days of the date on which they were legislated. In their empirical analysis, Mertens and Ravn treat these tax changes as being unanticipated.

can successfully recover the true impulse responses to a unanticipated fiscal shocks provided that these shocks account for a relatively large fraction of the variance of fiscal variables. Mertens and Ravn also estimate the effects of unanticipated government spending shocks in the U.S. using an augmented SVAR procedure that is robust to the presence of anticipated effects and find very similar results to those obtained from a standard SVAR.[22]

3.3 Dynamic Effects of Fiscal Policy Shocks

In this section, we study the dynamic effects of unanticipated spending and tax shocks and their importance in accounting for aggregate fluctuations. We then contrast these results with those obtained upon imposing the identifying restrictions commonly used in the literature.

3.3.1 Dynamic Responses

Figures 1.5 and 1.6 report the dynamic responses of output, government spending, taxes, the price of bonds, and the quantity of bonds to a government spending shock and to a tax shock, respectively. In each case, the shock is normalized to its unconditional standard deviation, i.e., unity. Since the quantity of bonds is not included in the SVAR, its response is constructed residually using the government budget constraint. The figures also report (possibly asymmetric) 68% confidence intervals computed using the procedure developed by Sims and Zha (1999).

Government Spending Shock

The upper panels of Figure 1.5 show that, in both sub-periods, a positive government spending shock leads to a temporary increase in output. The shape and the magnitude of the output response differ sharply, however, across the two periods: in the pre-1979 period, the increase in output is largest on impact and is statistically significant only during the first three quarters following the shock. In the subsequent quarters, the response becomes statistically insignificant. In contrast, in the post-1979 period, the response of output is persistent, mostly statistically significant, and hump-shaped, reaching its maximum at around 6 quarters after the shock.

[22] The augmented SVAR procedure, however, requires imposing additional identifying restrictions.

As is common in the literature, we quantify the effects of government spending shocks on output by computing the associated multiplier, which is defined as the dollar change in output that results from a dollar increase in the exogenous component of public spending. Table 1.IX reports the value of the multiplier on impact, at the 4 quarter horizon and at the peak. In the 1960:1–1979:2 period, the spending multiplier is 0.93 on impact and barely exceeds 1 at the peak. The corresponding numbers for the 1979:3–2007:4 period are 1.34 and 2.66, respectively. These numbers indicate that fiscal spending appears to have become more effective in stimulating economic activity after 1979.

Figure 1.5 shows that taxes are initially essentially unresponsive to the government spending shock, suggesting that the increase in spending is mostly financed by debt. Since the price of bonds decreases on impact in the first sub-period and remains roughly constant in the second, the government budget constraint implies that the quantity of issued bonds must increase in both cases (to finance the increase in public spending), which is what the lower panels of Figures 1.5 show.

Tax shock

Figure 1.6 depicts the dynamic responses to a positive tax shock. The upper panels of this figure show notable differences in the response of output across the two sub-periods. In the pre-1979 period, output remains inertial for about three quarters after the shock before starting to fall in a persistent and statistically significant manner. After reaching a trough at around six quarters after the shock, output returns gradually to trend. This U-shaped pattern is much less apparent in the post-1979 period, where the unexpected increase in taxes leads to an immediate small increase in output followed by a very persistent, though statistically insignificant, decline.[23]

Table 1.IX reports the values of the tax multiplier, defined as the dollar increase in output resulting from a dollar cut in the exogenous component of taxes. The tax multiplier is essentially zero on impact in the 1960:1–1979:2 period and even negative in the 1979:3–2007:4 period. The maximum multiplier is larger in the former period than in the latter (0.84 versus 0.51), but it is less than 1 in both cases. Importantly, we find that the tax multiplier is generally smaller than the spending multiplier, consistent

[23]Dotsey (1994) shows that a positive response of output to an increase in taxes is not necessarily at odds with economic theory. In his model, higher capital-tax rates lead to an increase in output and investment when current government deficits are financed by future distortionary taxation.

with traditional Keynesian theory.[24] This result stands in contrast to that reported by Mountford and Uhlig (2009) who find that tax cuts are more effective than increases in government spending to boost the economy.

A second discrepancy in the results across the two periods concerns the response of government spending, which is positive in the pre-1979 period but negative after 1979. None of these responses, however, is statistically distinguishable from 0. Thus, our results provide little support for the so-called *starve-the-beast* hypothesis, which states that tax cuts should lead to a reduction in future government spending. Romer and Romer (2008) have recently emphasized the importance to test this hypothesis using *exogenous* measures of taxes to avoid biases due to inverse causation and omitted variables. Using the narrative records to isolate legislated tax changes that are unlikely to be correlated with other factors affecting government spending, they also find little evidence in favor of the starve-the-beast hypothesis.

The price of bonds also responds asymmetrically across the two sub-samples, rising significantly in the pre-1979 period and falling in the post-1979 period. In both cases, however, the initial increase in taxes is so large (relative to the response of government spending and the price of bonds) that the quantity of issued bonds falls after the shock.

3.3.2 Variance Decomposition

Figures 1.7 and 1.8 report the contribution (in percentage) of, respectively, government spending and tax shocks to the variance of the k-step ahead forecast error of output, government spending, taxes, the price and the quantity of bonds. The dashed lines delimit the confidence intervals.

Two main messages emerge from these figures. First, fiscal policy shocks explain a small fraction of output variability, especially at short horizons. Government spending and tax shock jointly explain less than 10 percent of the variance of output at horizons of less than four quarters in the pre-1979 period and less than 15 percent in the post-1979 period. Second, the relative importance of government spending and tax shocks in accounting for the variance of output has changed over time.

[24] The only exception occurs at the four-quarter horizon in the pre-1979 period.

While government spending shocks explain barely 5 percent of output variability at any given horizon before 1979, this fraction became larger than 10 percent in the post-1979 period. Conversely, the contribution of tax shocks to the variance of output fell from roughly 25 percent to less than 5 percent at the 20-quarter horizon.

Variance-decomposition results also indicate that government spending shocks account for more than 90 percent of the variance of the one-quarter ahead forecast error of fiscal spending regardless of the period considered. This result reflects the notion that spending is largely predetermined with respect to economic activity (recall that the restrictions associated with the G indicator are not rejected by the data). In contrast, the contribution of taxes to the variance of the one-quarter ahead forecast error of taxes is around 70 percent, which suggests that an important fraction of tax changes occur in response to other shocks to the economy.

Finally, the results show that tax shocks are much more important than government spending shocks in accounting for the variability of the price and the quantity of bonds. The contribution of tax shocks to the variance of the price of bonds is larger than 10 percent at any horizon before 1979 and exceeds 40 percent in the post-1979 period. Government spending shocks, on the other hand, explain less than 5 percent of this variance at almost any horizon (in both periods). For the quantity of bonds, the contribution of tax shocks is always larger than 40 percent at short horizons (and reaches 75 percent in the pre-1979 period), whereas that of government spending shocks never exceeds 10 percent.

3.3.3 Comparison with the Restricted Systems

It is instructive at this stage to assess the implications of imposing the various sets of identifying restrictions discussed in Section 2.2. More specifically, the purpose of this section is to determine whether these restrictions lead to significant departures from the impulse-response and variance-decomposition results discussed above.[25] Figure 1.9 superimposes on the unrestricted dynamic responses to a government spending shock those obtained using the Cholesky and Blanchard-Perotti identification schemes.[26]

[25] Since the PD indicator is found to be strongly rejected in the data in both sub-samples, we shall not discuss it any further.

[26] For the sake of remaining as faithful as possible to the original specification of Blanchard and Perotti (2002), the results based on their model are generated using a three-variable VAR that includes output, government spending and taxes.

The figure shows that the two restricted systems overestimate the effects of government spending shocks on output in the 1960:1–1979:2 period and underestimates it in the 1979:3–2007:4 period. This is in turn reflected in the values of the spending multiplier: Table 1.IX shows that under the Cholesky and Blanchard-Perotti identification schemes the multiplier is larger than 1 on impact in the pre-1979 period but less than 1 in the post-1979 period, which is the opposite of what the unrestricted system predicts. A similar message is conveyed by Figure 1.10, which compares the unrestricted and restricted variance-decomposition results for government spending shocks. The figure shows that the two restricted systems overstate the fraction of output variability explained by government spending shocks in the pre-1979 period, while the opposite scenario holds in the post-1979 period.

While Figures 1.9 and 1.10 show that the Cholesky and Blanchard-Perotti identification schemes lead to deviations from the unrestricted responses, these deviations are in general more pronounced in the post-1979 period. Given that the restrictions associated with the G indicator have been shown to be supported by the data (see Panel A of Table 1.VI), these deviations are likely due to the restrictions associated with the transmission mechanism of government spending shocks. In order to investigate this conjecture, we test these restrictions using a likelihood-ratio test. The results, shown in Panel B of Table 1.VI, indicate that, both under the Cholesky and Blanchard-Perotti schemes, these restrictions are strongly rejected by the data in the post-1979 period, but not in the pre-1979 period.[27]

Figure 1.11 contrasts the unrestricted responses to a tax shock with those implied by Blanchard and Perotti's specification. In both periods, the latter involves significant departures from the unrestricted dynamic responses. For example, the peak response of output to a tax shock is much larger in the restricted system, with a multiplier of 1.04 in the pre-1979 period and 0.78 in the post-1979 period, whereas the corresponding numbers are 0.84 and 0.51 in the unrestricted system. Figure 1.12 shows that Blanchard and Perotti's specification also implies that tax shocks account for a counterfactually large fraction of the variance of output at long horizons in the 1960:1–1979:2 period. A likelihood-ratio test of the restrictions associated with the transmission mechanism imposed by Blanchard and Perotti indicates that these restrictions are rejected at the 10 percent level in the pre-1979 period (see Panel

[27] Blanchard and Perotti (2002) estimate the average output elasticity of taxes, x, to be 2.08 based on data from 1947:1 to 1997:4. In our tests, however, we consider the values of 1.75 and 1.97 estimated by Perotti (2004) for the periods 1960:1–1997:4 and 1980:1–2001:4, respectively.

B of Table 1.VI). For the post-1979 period, on the other hand, these restrictions are not rejected by the data, which suggests that the discrepancy with respect to the unrestricted results are mainly due to the restrictions associated with the CAT indicator (which are strongly rejected in this period. See Panel A of Table 1.VI).

4 Extensions

Having analyzed the effects of fiscal policy shocks on aggregate output, we now investigate how these shocks affect private consumption and investment. This exercise is useful on two counts. First, it helps determine which type of private expenditure is more responsive to fiscal policy, thus allowing a better understanding of the channels through which fiscal instruments affect aggregate output. Second, the response of private consumption to unanticipated changes in government spending is useful to discriminate between competing views of fiscal policy: According to Keynesian theory, an increase in government spending should lead to an increase in consumption, whereas standard neoclassical models predict that public spending crowds out consumption due to a negative wealth effect (Barro and King, 1984 and Baxter and King, 1993). While most of existing studies using SVARs tend to corroborate the crowding-in effect, the magnitude of this effect is sensitive to identification. Furthermore, this effect may well vanish altogether if one relaxes the commonly used identifying assumptions.

To examine the implications of our unrestricted SVAR for consumption and investment, we extend system (7) by adding each of these two variables one at a time in a way that preserves specification (3)–(6) but that leaves the dynamics of the added variable unrestricted (as is the case for output). The implied dynamic responses of consumption to government spending and tax shocks are reported in Figures 1.13 and 1.14, respectively. The corresponding results for investment are shown in Figures 1.15 and 1.16. In order to gain further insights into the effects of fiscal policy shocks on the different categories of private spending, we also report results for durable and non-durable consumption and for residential and non-residential investment.

4.1 Consumption

In response to an unexpected increase in government spending, total consumption rises significantly in the two periods considered (Figure 1.13). In the pre-1979 period, the increase is large but short lived, becoming statistically insignificant within a year after the shock. A much more persistent pattern is observed in the post-1979 period, where the rise in consumption remains statistically significant for roughly three years. Looking at the response of durables and non-durables, one can see that the increase in total consumption before 1979 reflects almost exclusively the rise in the consumption of non-durable goods–the response of durable consumption being nil on impact and statistically insignificant at all horizons. In contrast, after 1979, both consumption categories contribute to the observed increase in total consumption. In sum, there is strong evidence of a crowding-in effect of public spending on private consumption, contrary to neoclassical theory.[28]

In the pre-1979 period, an unanticipated increase in taxes lowers total and non-durable consumption, but with a delay of several quarters (Figure 1.14). In both cases, the response is statistically significant in a window of 6 to 12 quarters after the shock and reaches its trough at around 9 quarters after the shock. Durable consumption, on the other hand, displays an oscillatory but statistically insignificant response. In the post-1979 period, tax shocks do not affect total consumption or its two categories in any significant manner. These findings corroborate our earlier conclusion that U.S. tax policy has become less effective after 1979.

4.2 Investment

According to the point estimates in Figure 1.15, a positive government spending shock decreases total investment in the pre-1979 period and raises it in the post-1979 period. Both responses, however, are statistically insignificant. While the response of non-residential investment is completely muted before 1979 and mostly statistically insignificant after 1979, residential investment reacts strongly but quite differently across the two periods: Before 1979, it begins to decline gradually until it reaches a significantly large trough at around 7 quarters after the shock, whereas it reacts positively after 1979, at least during the two quarters following the shock.

[28] Several explanations have been proposed to reconcile theory with data. Most of these explanations operate through consumer preferences (Bouakez and Rebei, 2007, Ravn, Schmitt-Grohé, and Uribe 2007, Monacelli and Perotti, 2008). Galí, López-Salido and Vallés (2007), on the other hand, propose a resolution that emphasizes the interaction of sticky prices, non-Ricardian consumers and a non-competitive labor market.

A significant sub-sample instability is also found in the response of total investment to a tax shock (Figure 1.16). After an initial increase during the two quarters following the shock, total investment falls significantly for several quarters in the pre-1979 period, while it continues to rise in a hump-shaped manner in the post-1979 period. These patterns are also apparent in the dynamic responses of non-residential investment. As for residential investment, we find that it exhibits a delayed decline in both sub-periods, reaching its trough between 1 and 2 years after the shock.

Overall, these results indicate that fiscal policy shocks affect aggregate output mainly through their effect on private consumption. Investment is essentially unresponsive to government spending shocks, and although it increases sharply following a tax cut, this effect only marginally affects output.[29]

5 Conclusion

The purpose of this paper was to estimate the macroeconomic effects of fiscal policy shocks in the U.S. using an alternative empirical methodology that relaxes the identifying restrictions commonly used in the SVAR literature. Identification is instead achieved by exploiting the conditional heteroscedasticity of the structural innovations. This approach avoids making arbitrary assumptions about the relevant policy indicator or its transmission mechanism.

Several important findings emerge from this study. First, based on historical data, increases in government spending are found to be more effective than tax cuts in stimulating U.S. economic activity. This conclusion supports the Keynesian view. Second, the dynamic effects of fiscal policy shocks and their relative importance to output fluctuations have changed significantly after 1979. Since this date is also believed to have marked an important shift in U.S. monetary policy, it would be interesting to investigate whether, and to what extent, the two phenomena are linked. Third, the crowding-in effect of public spending on private consumption documented in earlier SVAR-based studies is robust to relaxing conventional identifying assumptions. While a number of solutions have been proposed to reconcile this evidence with neoclassical theory, we believe more empirical work is needed to unravel the exact mechanism that gives rise to the positive covariance of public and private expenditures.

[29] Fatàs and Mihov (2001) also find no significant effect of government spending shocks on residential and non-residential investment. In contrast, Mountford and Uhlig (2009) report that both investment categories respond negatively and in a statistically significant manner to government spending and tax shocks.

6 References

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7 Appendix

7.1 Appendix B: Figures

Figure 1.1: Transformed Data

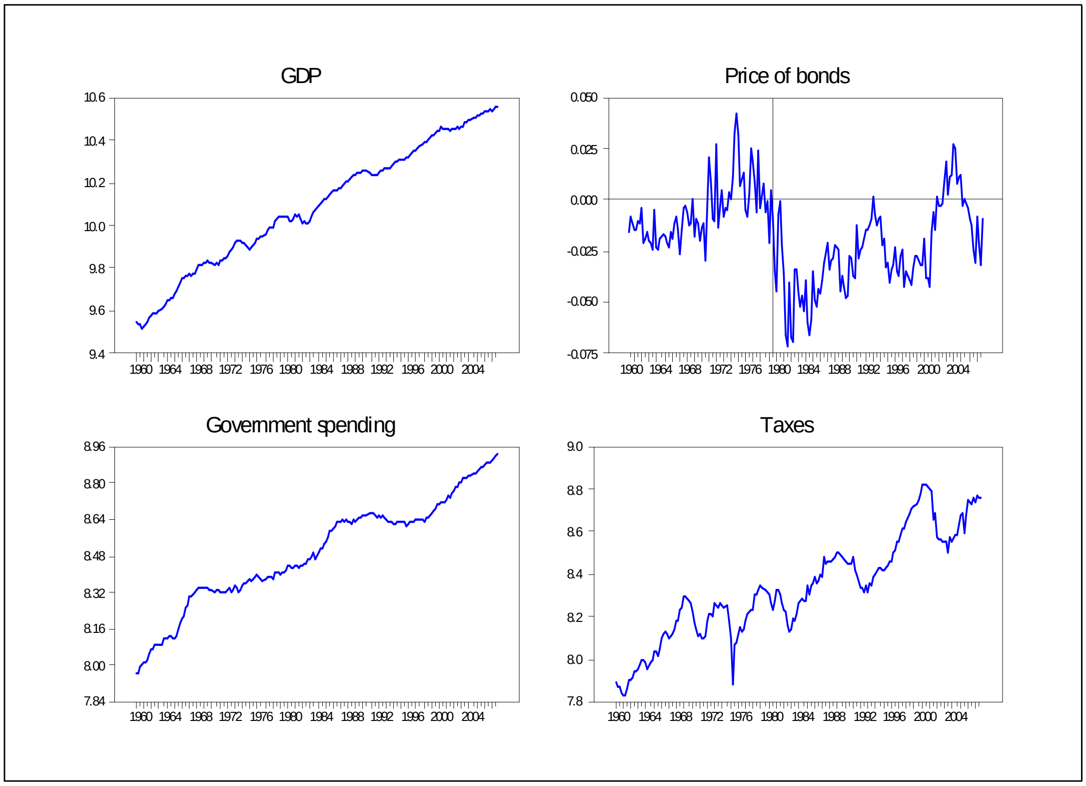


Figure 1.2: Conditional Variances of the Shocks

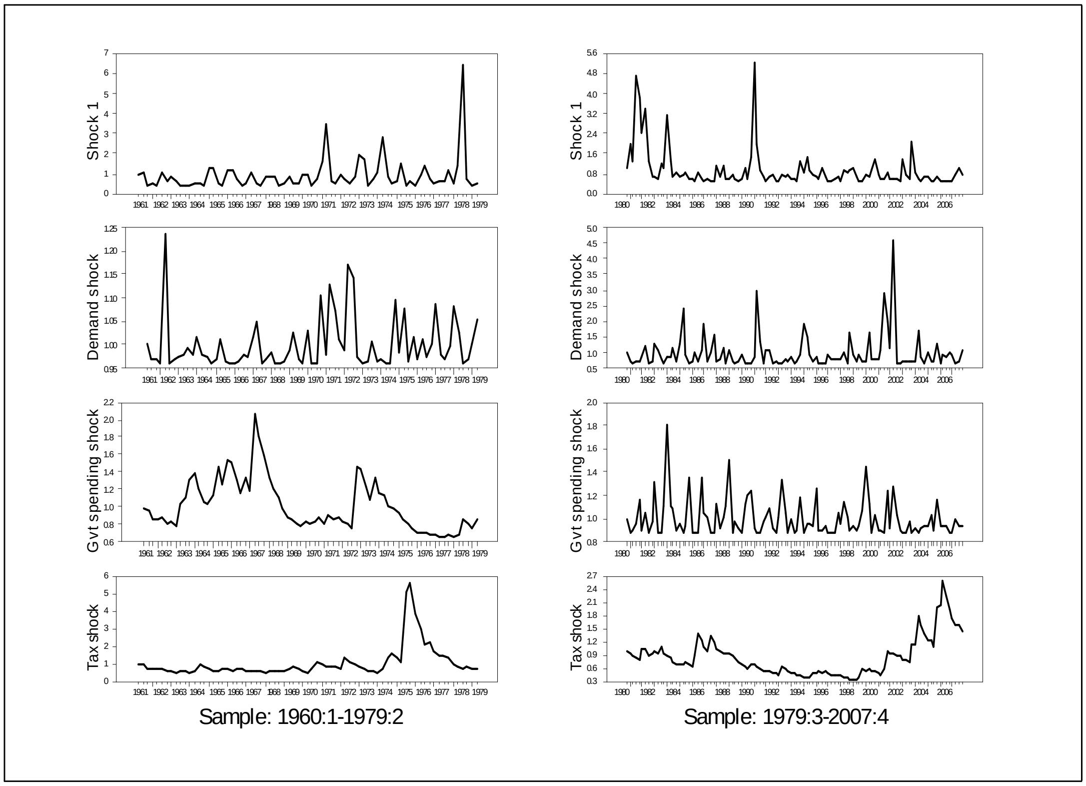
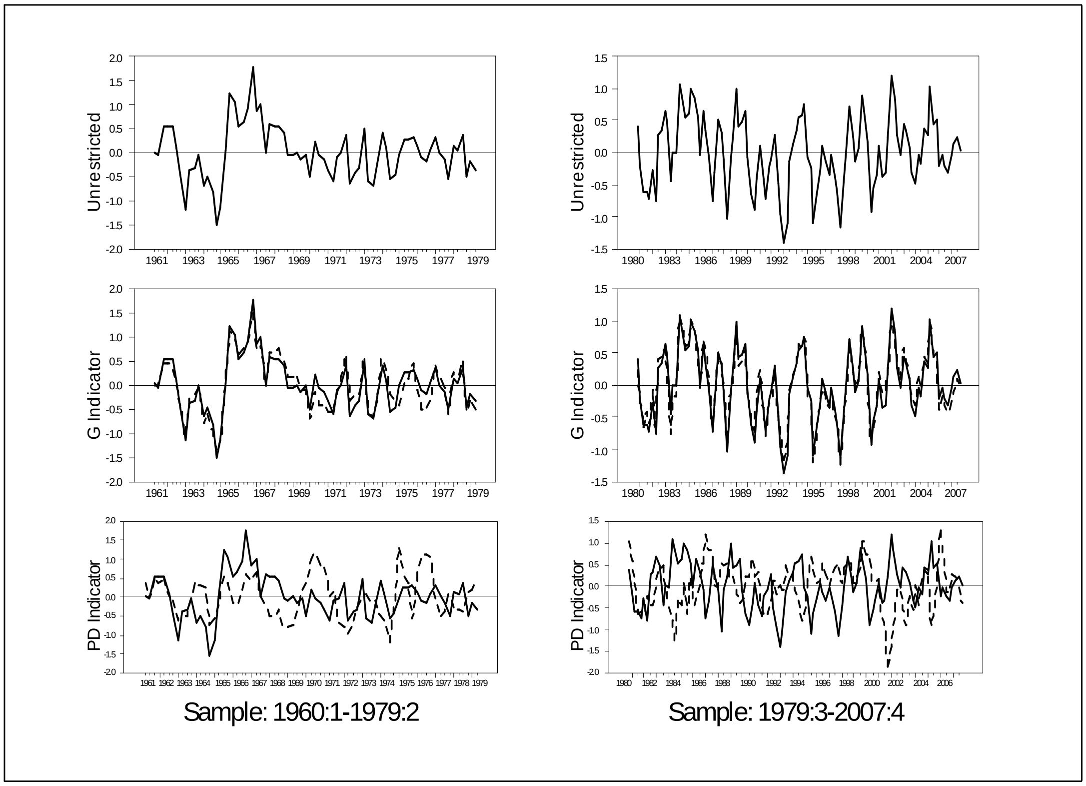
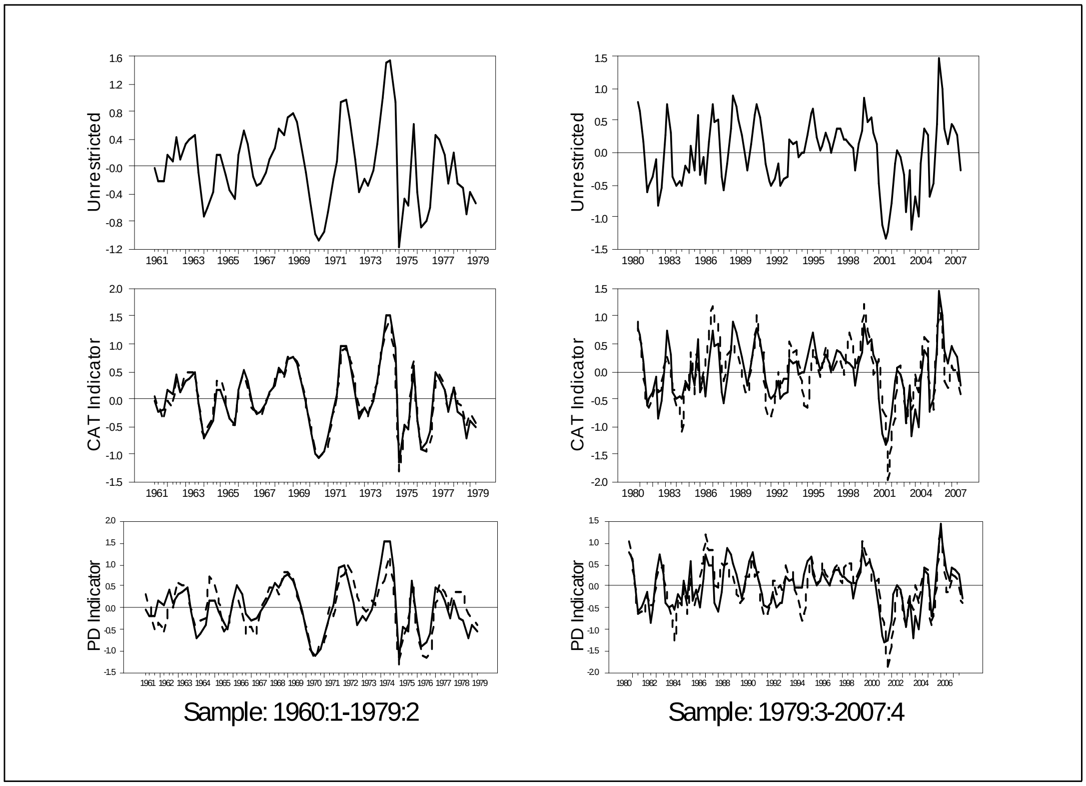


Figure 1.3: Government Spending Shocks



Solid lines: unrestricted measures, dashes: restricted measures.

Figure 1.4: Tax Shocks



Solid lines: unrestricted measures, dashes: restricted measures.

Figure 1.5: Dynamic Responses to a Government Spending Shock

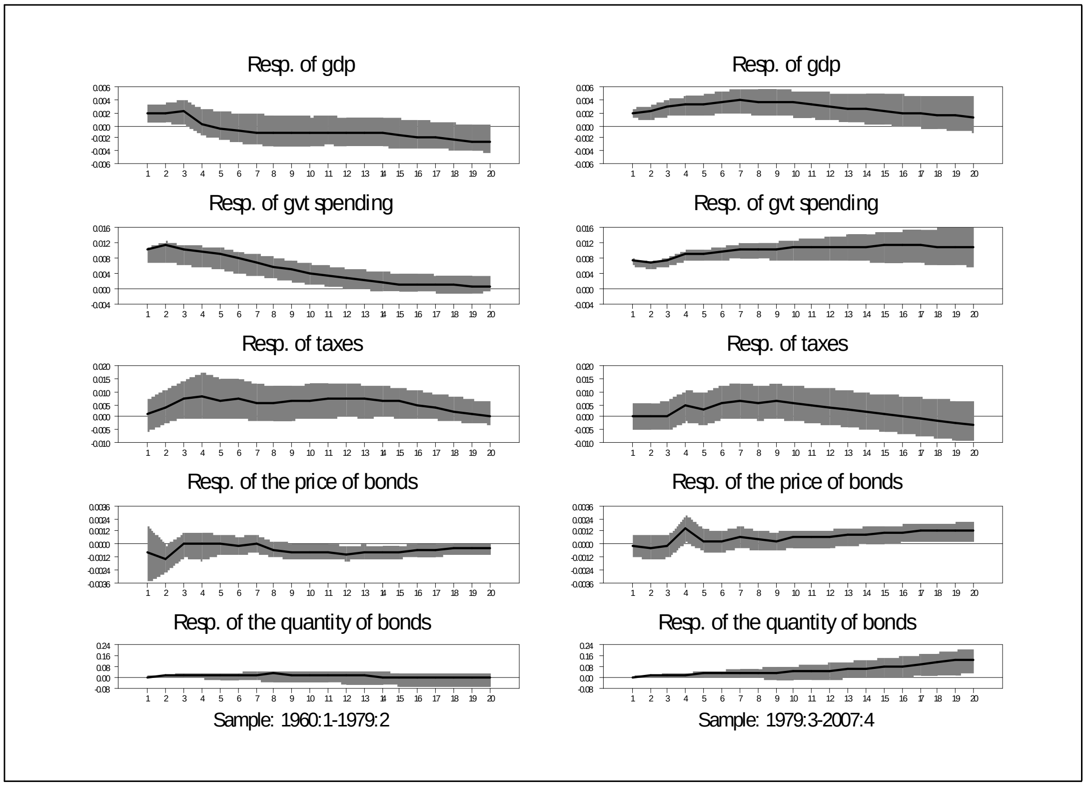


Figure 1.6: Dynamic Responses to a Tax Shock

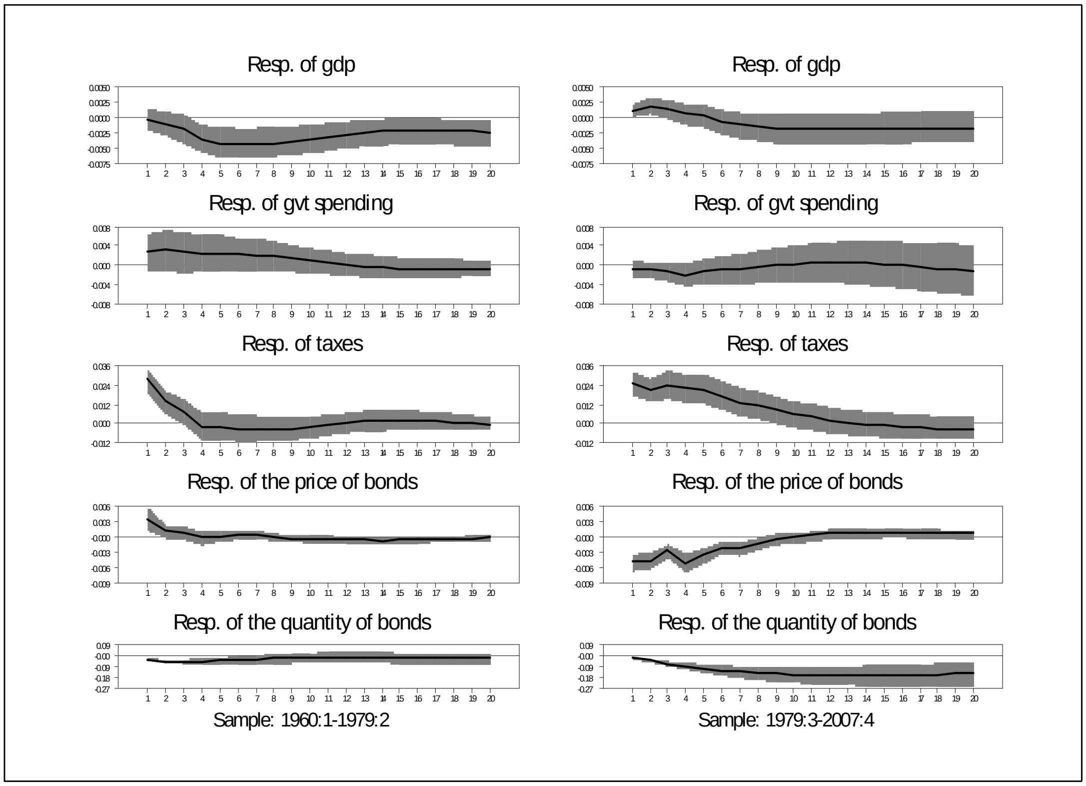


Figure 1.7: Contribution of Government Spending Shocks to the Conditional Variance of the K-Step Ahead Forecast Errors

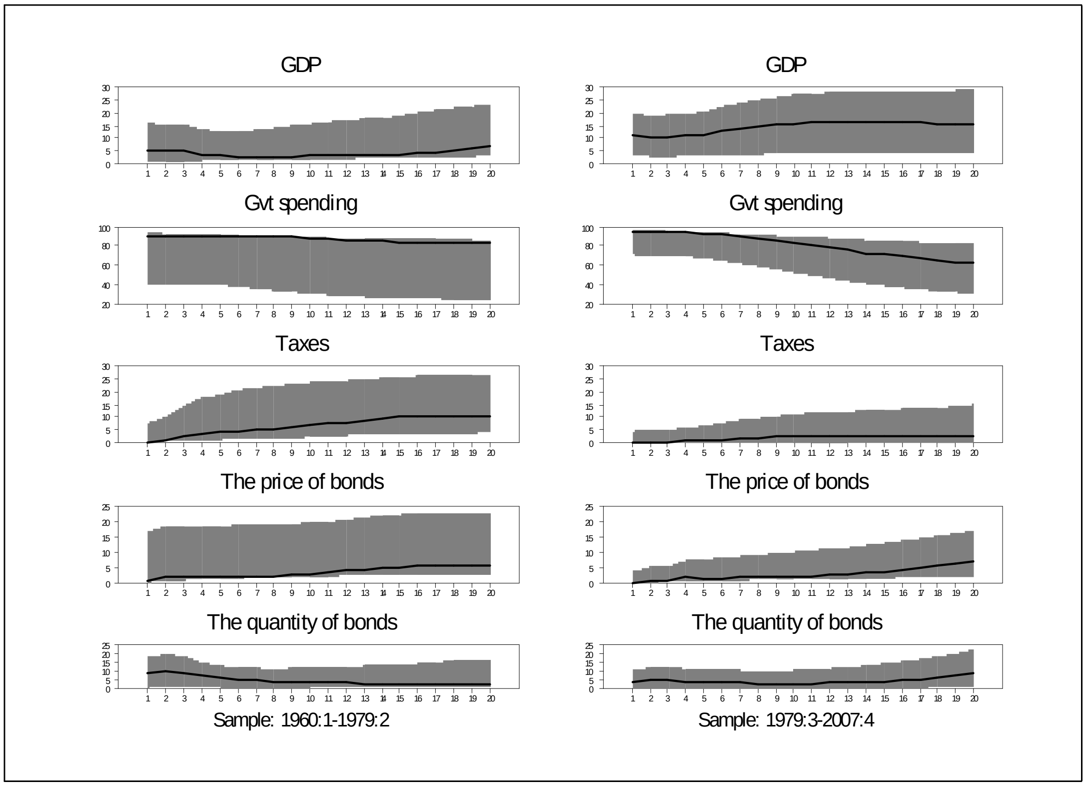


Figure 1.8: Contribution of Tax Shocks to the Conditional Variance of the K-Step Ahead Forecast Errors

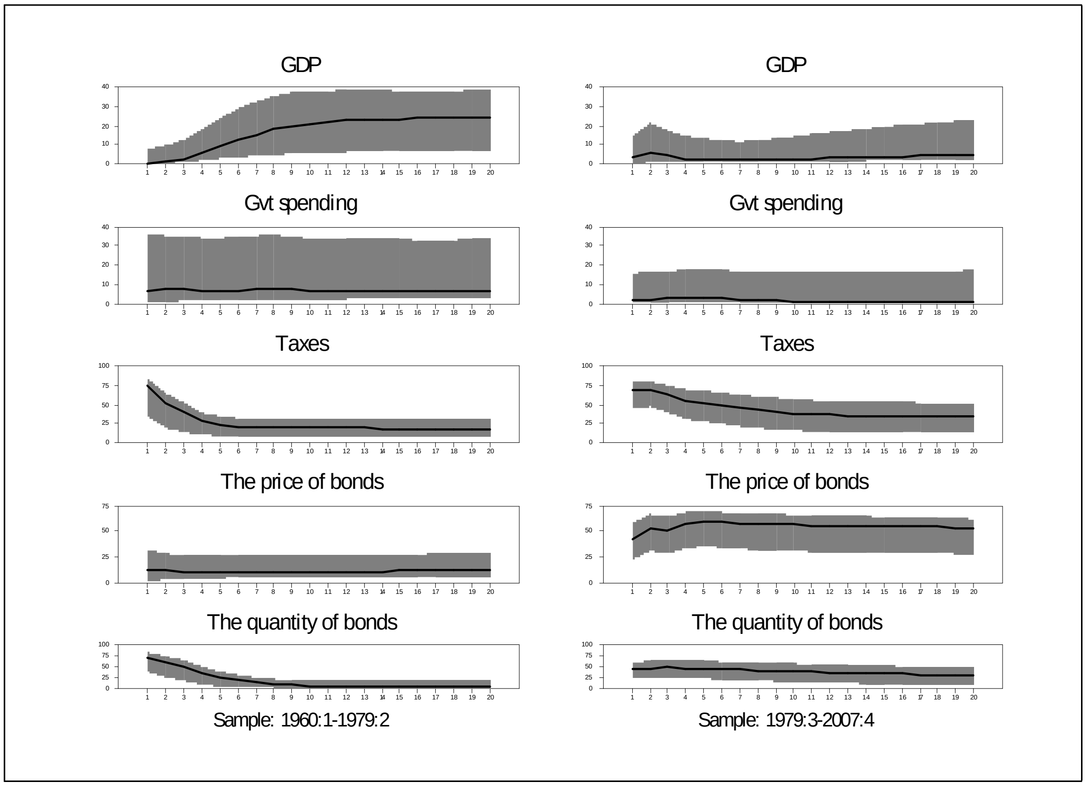
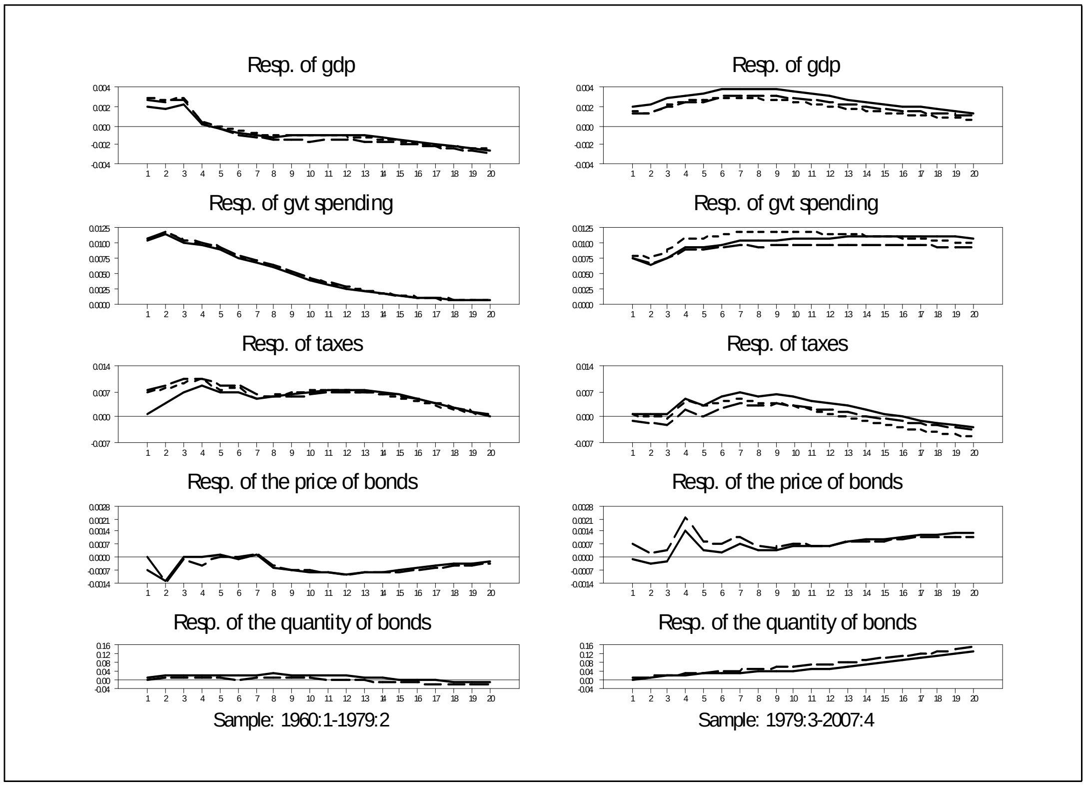


Figure 1.9: Dynamic Responses to a Government Spending Shock:

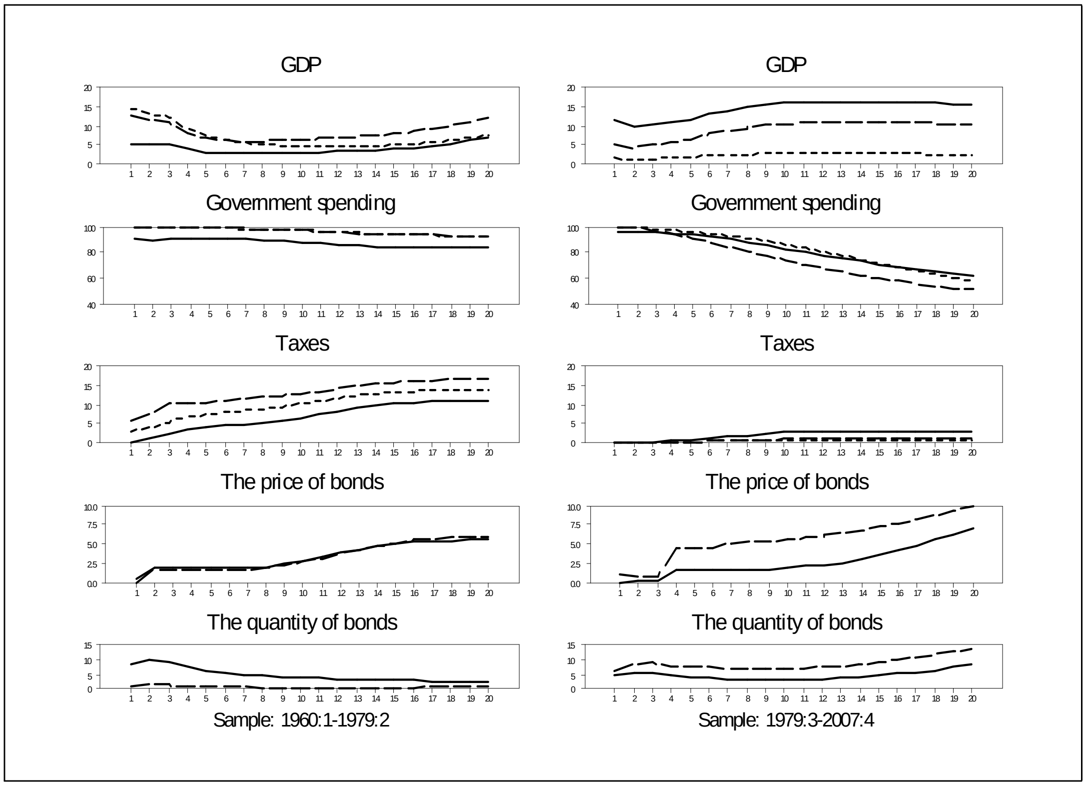
Unrestricted vs. Restricted Systems



Solid lines: unrestricted system, long dashes: Cholesky decomposition, short dashes: Blanchard-Perotti.

Figure 1.10: Contribution of Government Spending Shocks to the Conditional Variance

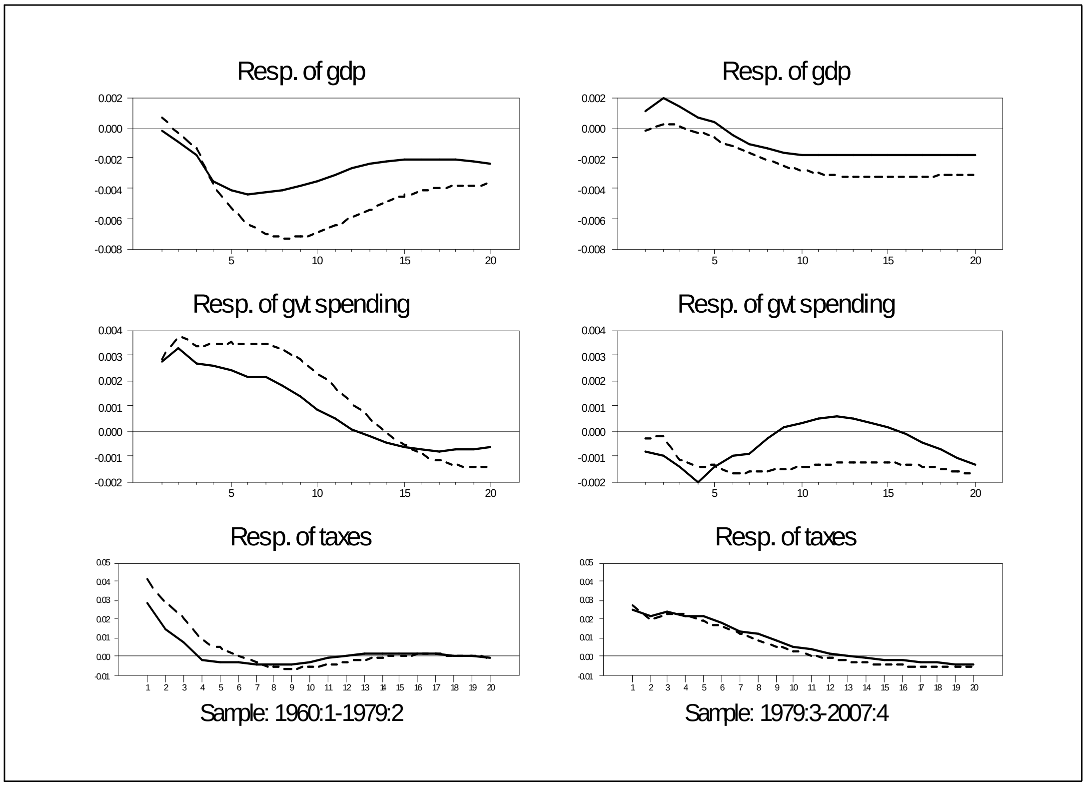
of the K-Step Ahead Forecast Errors: Unrestricted vs. Restricted Systems



Solid lines: unrestricted system, long dashes: Cholesky decomposition, short dashes: Blanchard-Perotti.

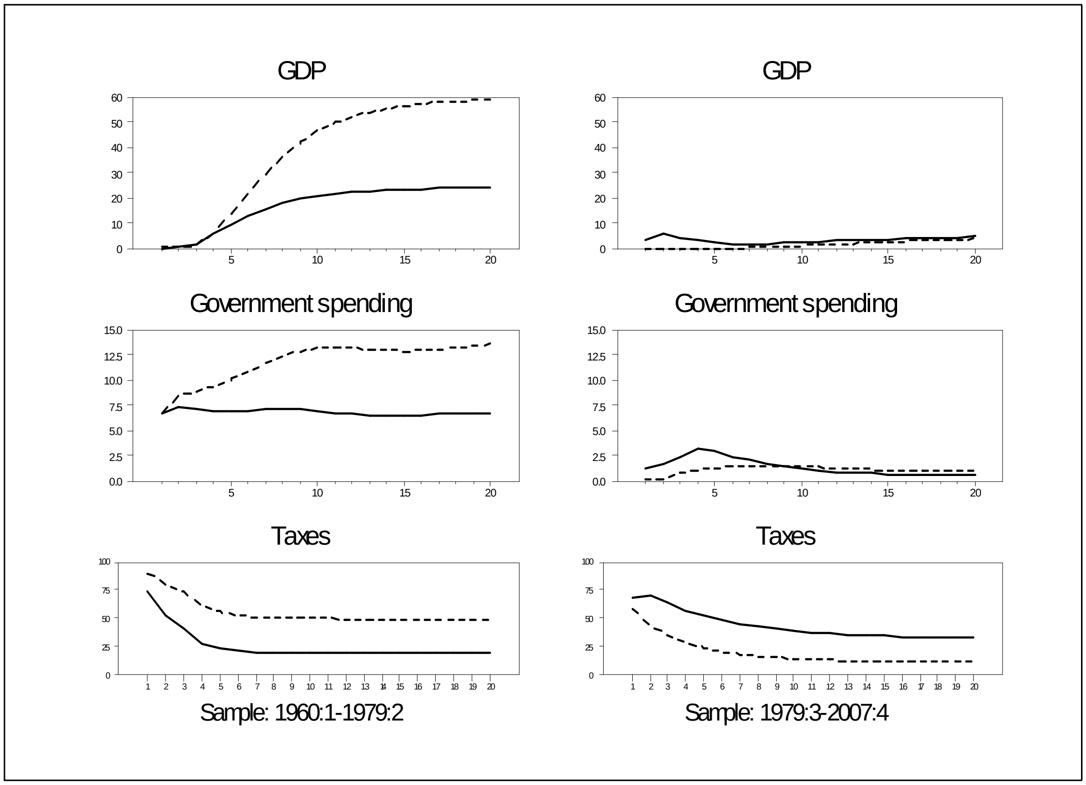
Figure 1.11: Dynamic Responses to a Tax Shock:

Unrestricted vs. Restricted Systems



Solid lines: unrestricted system, short dashes: Blanchard-Perotti.

Figure 1.12: Contribution of Tax Shocks to the Conditional Variance of the K-Step Ahead Forecast Errors: Unrestricted vs. Restricted Systems



Solid lines: unrestricted system, short dashes: Blanchard-Perotti

Figure 1.13: Dynamic Responses of Consumption to a Government Spending Shock

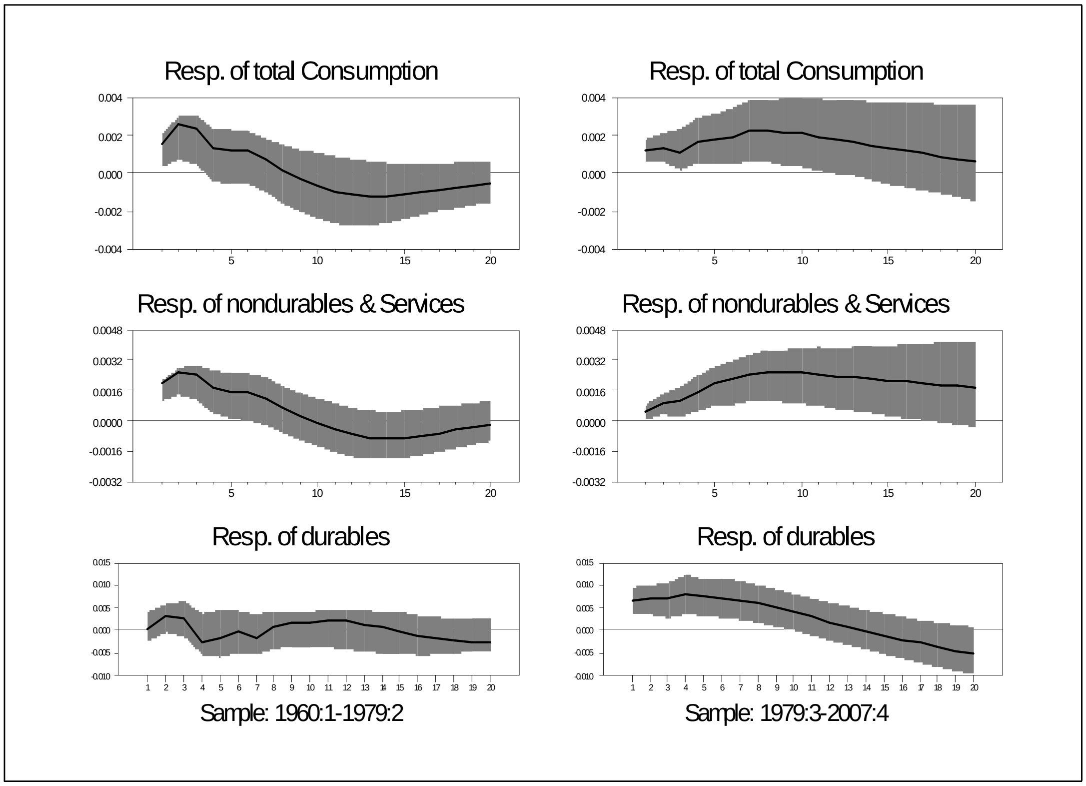


Figure 1.14: Dynamic Responses of Consumption to a Tax Shock

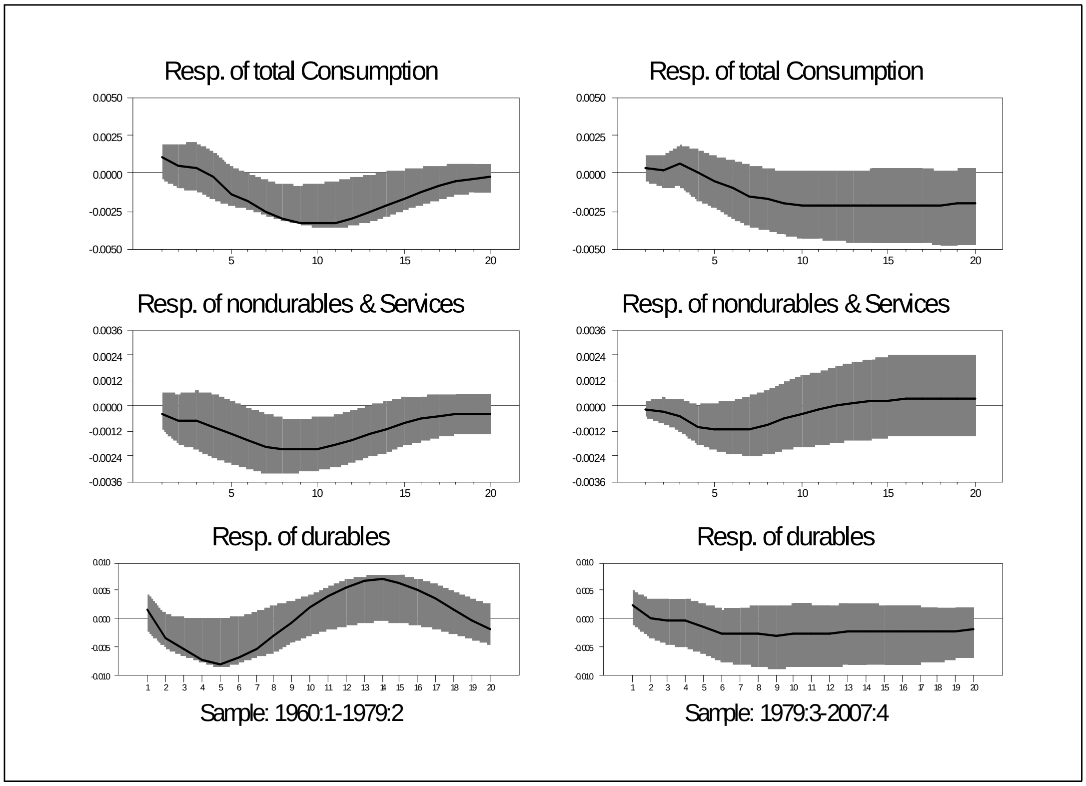


Figure 1.15: Dynamic Responses of Investment to a Government Spending Shock

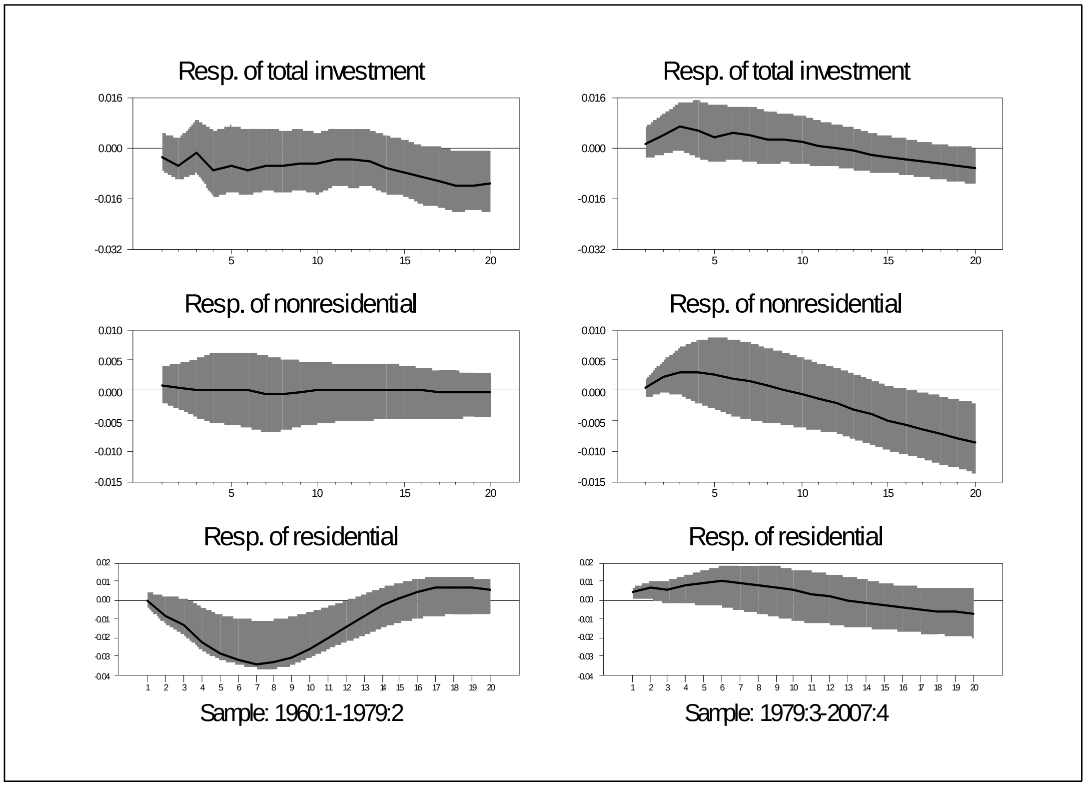
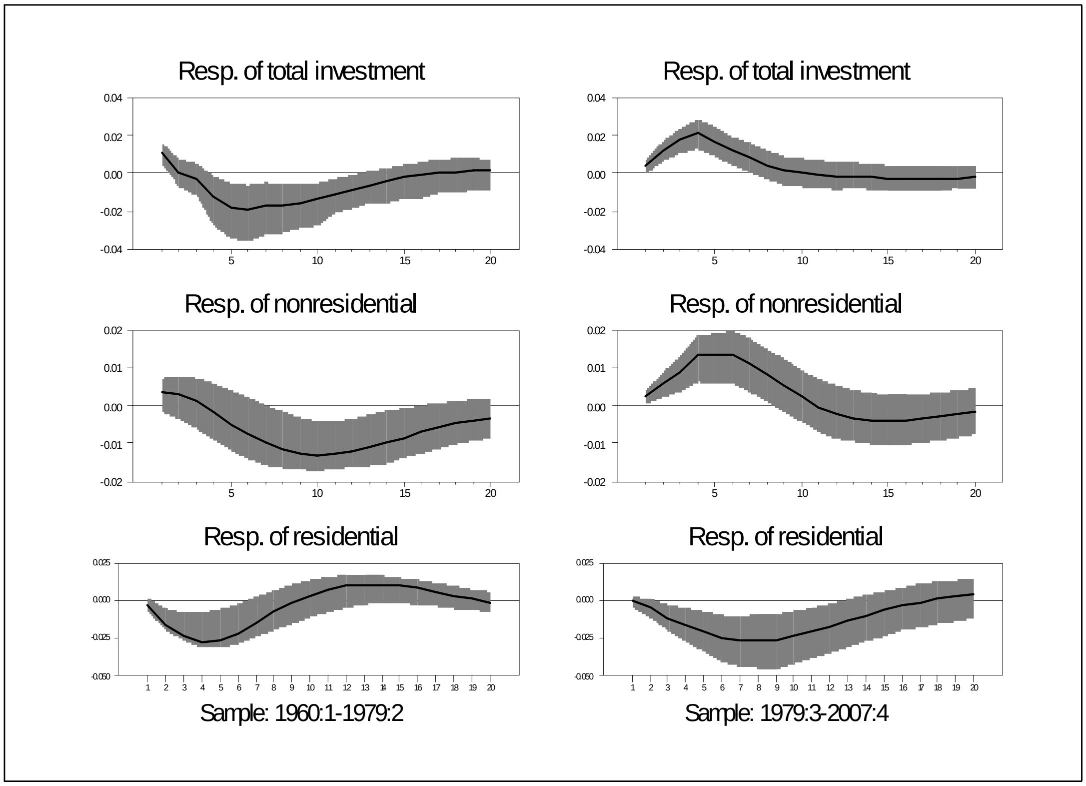


Figure 1.16: Dynamic Responses of Investment to a Tax Shock



7.2 Appendix B: Tables

Table 1.I. Multivariate ARCH Test for the VAR Residuals

Lag	1960:1–1979:2	1979:3–2007:4
1	0.000	0.000
2	0.000	0.000
4	0.000	0.000

Notes: Entries are p-values of the χ^2-distributed Lagrange multiplier test statistic.

Table 1.II. Multivariate ARCH Test for the Structural Shocks

Lag	1960:1–1979:2	1979:3–2007:4
1	0.000	0.001
2	0.461	0.000
4	0.000	0.307

Notes: Entries are p-values of the χ^2-distributed Lagrange multiplier test statistic.

Table 1.III. Specification Test Results

	Lag	1960:1–1979:2	1979:3–2007:4
$\epsilon^2_{1,t}$	1	0.579	0.534
	2	0.099	0.736
	4	0.144	0.358
$\epsilon^2_{d,t}$	1	0.972	0.592
	2	0.982	0.354
	4	0.698	0.166
$\epsilon^2_{g,t}$	1	0.544	0.968
	2	0.713	0.594
	4	0.722	0.313
$\epsilon^2_{\tau,t}$	1	0.852	0.642
	2	0.781	0.713
	4	0.679	0.219

Notes: Entries are the p-values associated with the McLeod-Li test statistic applied to the squared structural innovations relative to their conditional variances.

Table 1.IV. Estimates of the Structural Parameters

Parameter	Estimate 1960:1–1979:2	1979:3–2007:4
α	15.100	13.906
	(89.079)	(22.075)
β	−0.777	3.802
	(10.901)	(4.765)
η_g	0.224	−0.204
	(0.249)	(0.281)
η_τ	1.913	1.085
	(0.973)	(0.673)
θ_g	0.001	0.006
	(0.032)	(0.017)
θ_τ	−0.026	0.133
	(0.173)	(0.225)
ψ_g	0.097	−0.027
	(0.158)	(0.088)
ψ_τ	−0.293	−0.183
	(0.782)	(0.676)
σ_d	0.140	0.122
	(0.871)	(0.203)
σ_g	0.010	0.008
	(0.002)	(0.001)
σ_τ	0.029	0.024
	(0.007)	(0.008)

Note: Figures between parentheses are standard errors.

Table 1.V. Test of the Restriction: $a_{21} + a_{23} = -a_{24}$

	1960:1–1979:2	1979:3–2007:4
P-value	0.319	0.104

Note: P-values are those of the χ^2-distributed likelihood-ratio test statistic.

Table 1.VI. Tests of Commonly Used Identifying Restrictions

Type	Restrictions	1960:1–1979:2	1979:3–2007:4
A. Tests of Alternative Indicators of Fiscal Policy			
Spending Policy			
G	$\eta_g = \theta_g = \psi_g = 0$	0.293	0.670
PD	$\eta_g = \eta_\tau$, $\theta_g = \theta_\tau$, $\psi_G = 1$	0.063	0.004
Tax Policy			
CAT	$\theta_\tau = \psi_\tau = 0$	0.571	0.004
PD	$\eta_g = \eta_\tau$, $\theta_g = \theta_\tau$, $\psi_\tau = 1$	0.063	0.004
B. Tests of the Restrictions Associated with the Transmission of Fiscal Policy			
Spending Policy			
Cholesky	$a_{12} = a_{14} = \theta_\tau = 0$	0.155	0.001
Blanchard and Perotti	$a_{12} = \theta_\tau = 0$ and $\eta_\tau = x$	0.152	0.000
Tax Policy			
Blanchard and Perotti	$a_{12} = \theta_g = \eta_g = 0$ and $\eta_\tau = x$	0.093	0.149

Notes: Entries are p-values of the χ^2-distributed likelihood-ratio test statistics. x is fixed to 1.75 in 1960:1–1979:2 and 1.97 in 1979:3–2007:4.

Table 1.VII. Correlations Between the Unrestricted and Restricted Measures of Fiscal Policy Shocks

	Spending Policy		Tax Policy	
Policy indicator	1960:1–1979:2	1979:3–2007:4	1960:1–1979:2	1979:3–2007:4
G	0.988	0.967	–	–
	(0.005)	(0.011)	–	–
CAT	–	–	0.984	0.808
	–	–	(0.003)	(0.030)
PD	0.669	−0.115	0.841	0.807
	(0.151)	(0.195)	(0.032)	(0.030)

Note: Figures between parentheses are standard errors.

Table 1.VIII. Granger Causality Tests

	A. Do Ramey & Shapiro's Dates Cause SVAR-Based Government Spending Shocks?		B. Do Romer & Romer's Dates Cause SVAR-Based Tax Shocks?	
Policy indicator	1960:1–1979:2	1979:3–2007:4	1960:1–1979:2	1979:3–2007:4
Unrestricted	0.157	0.422	0.807	0.518
G	0.281	0.553	–	–
CAT	–	–	0.792	0.427
PD	0.565	0.030	0.468	0.515

Note: Entries are the p-values of the F-distributed statistic used to test the joint significance of the coefficients in a regression of the SVAR-based shocks on four lags of the dates.

Table 1.IX. Fiscal Multipliers

	1960:1–1979:2			1979:3–2007:4		
System	1 Q	4 Q	peak	1 Q	4 Q	peak
Spending Policy						
Unrestricted	0.927	0.037	1.028	1.342	2.151	2.656
Cholesky	1.192	0.121	1.194	0.887	1.645	2.232
Blanchard and Perotti	1.279	0.226	1.279	0.951	1.737	1.985
Tax Policy						
Unrestricted	0.039	0.682	0.843	−0.280	−0.161	0.509
Blanchard and Perotti	−0.082	0.525	1.036	0.044	0.066	0.780

Note: The multiplier is defined as the dollar change in output at a given horizon that results from a dollar increase (cut) in the exogenous component of government spending (taxes).

Essay 2. Fiscal Policy and External Adjustment: New Evidence

Abstract

Relatively little empirical evidence exists about countries' external adjustment to changes in fiscal policy and, in particular, to changes in taxes. This paper addresses this question by measuring the effects of tax and government spending shocks on the current account and the real exchange rate in a sample of four industrialized countries. Our analysis is based on a structural vector autoregression in which the interaction of fiscal variables and macroeconomic aggregates is left unrestricted. Identification is instead achieved by exploiting the heteroscedasticity of the structural disturbances. Three main findings emerge: (*i*) the data provide little support for the twin-deficit hypothesis, (*ii*) the estimated effects of unexpected tax cuts are generally inconsistent with the predictions of standard economic models, except for the US, and (*iii*) the puzzling real depreciation triggered by an expansionary public spending shock is substantially larger in magnitude than predicted by traditional identification approaches.

JEL classification: C32, E62, F41, H20, H50, H60.

Keywords: Government spending, Current account, Exchange rate, Taxes, Structural vector autoregression, Twin deficits.

1 Introduction

The latest financial crisis has revived interest in the macroeconomic effects of fiscal policy and its role as a stabilization tool, as nominal interest rates approached zero, leaving little room for monetary policy. However, while a large body of work has focused on assessing the effectiveness of tax and public spending policies in stimulating output and domestic absorption, relatively less effort has been devoted to studying the implications of those policies for countries' external adjustment and, by extension, for global imbalances. In particular, to our knowledge, only one paper, namely Kim and Roubini (2008), attempted to empirically evaluate the reaction of the current account and the real exchange rate to changes in taxes, and only a handful of studies attempted to measure the response of those two variables to changes in government spending (Corsetti and Müller 2006; Kim and Roubini 2008; Müller 2008; Monacelli and Perotti 2010; Enders, Müller and Scholl 2011). This is somewhat surprising given that the current account is commonly regarded as a barometer of a country's solvency, and that exchange-rate fluctuations critically affect a country's competitiveness on the world market and its trade balance.

Using structural vector autoregressions (SVARs) and focusing mostly on US data, the papers cited above find that unexpected tax cuts and increases in public spending unambiguously depreciate the real exchange rate. Kim and Roubini (2008) also find that a surprise tax cut worsens the budget deficit but improves the current account, a situation referred to as "twin divergence". On the other hand, no consensus has been reached regarding the effects of an unexpected increase in government spending on the current account, or whether it leads to twin divergence or twin deficits (i.e., positive comovement between the budget and external deficits).

Generally speaking, these findings are puzzling from a theoretical standpoint. A wide class of open-economy macro models indeed predict that an unexpected fiscal expansion should appreciate the currency in real terms and deteriorate the current account. In the case of a tax cut, the real appreciation occurs because there is a higher incentive to invest,[30] which raises the real interest rate. The rise in investment is typically larger than the increase in national saving, causing a current-account deficit. In the case of an increase in government spending, the appreciation results from the fact that

[30] This is the case as long as the tax cut is not lump sum.

public expenditures are relatively more intensive in domestically produced goods, which means that the increase in aggregate demand brought about by the increase in public spending will raise their relative price with respect to foreign goods. The rise in public spending also entails a negative wealth effect that induces households to borrow abroad to prevent a large drop in their consumption, thus worsening the current account.

The purpose of this paper is to provide new evidence on the effects of fiscal policy on changes in the net foreign position and on the real exchange rate in a sample of four industrialized countries, namely, Australia, Canada, the United Kingdom, and the United States. These four countries are known to have reliable non-interpolated quarterly data on fiscal variables. Our contribution to the empirical literature is threefold. First, we provide more comprehensive evidence on the response of the current account and the exchange rate to changes in taxes than Kim and Roubini, who focused exclusively on the US. Second, we use an estimation strategy that relaxes the identifying assumptions used in previous SVAR-based studies, which restrict the interaction of the variables of interest in a rather arbitrary way. Third, we document the implications of imposing these restrictions for the response of the current account and the exchange rate to fiscal shocks.

Our empirical strategy builds on that developed in our earlier work (Bouakez, Chihi, and Normandin 2010). More specifically, we identify fiscal-policy shocks by exploiting the conditional hetereoscedasticity of the shocks. When there is enough time variation in the conditional variances of the time series used in estimation, it becomes possible to identify the structural shocks and their effects without having to impose additional parametric restrictions, as would be the case under (the usually maintained assumption of) conditional homoscedasticity (see Sentana and Fiorentini 2001). Incidentally, several studies document that the macroeconomic time series that we use in our analysis display significant time-varying conditional volatilities.[31] In our framework, the matrix of contemporaneous interaction nests the parametric restrictions typically imposed in the literature, thereby allowing one to assess the bias resulting from such restrictions.[32]

[31] See, for example, Hsieh (1988, 1989), Engel and Hamilton (1990), Garcia and Perron (1996), Den Haan and Spear (1998), Engel and Kim (1999), Fountas and Karanasos (2007), Fernandez-Villaverde, Guerrón-Quintana, Kuester, and Rubio-Ramírez (2010), and Fernandez-Villaverde, Guerrón-Quintana, Rubio-Ramírez, and Uribe (2011).

[32] Leeper, Walker and Yang (2008) pointed out that the SVAR approach may not be robust to fiscal forsight– the phenomenon that, due to legislative and implementation lags, economic agents are likely to react to changes in taxes and governement spending several months before those changes actually take place. In the extreme case where all fiscal shocks

The empirical framework developed in our earlier paper casts fiscal policy in the context of a market for newly issued government bonds. The supply of bonds may or may not shift as a result of changes in taxes or public expenditures, depending on the government's implicit target. In turn, variations in taxes and public expenditures reflect both the automatic and systematic responses of these variables to changes in economic conditions, as well as fiscal-policy shocks. We extend this framework by assuming that the demand for government bonds originates not only domestically but also abroad, implying that the real exchange rate enters the bond-demand equation. We also include the current account among the vector used in estimation, while leaving its interaction with the remaining variables completely unrestricted.

Our results show important differences in the response of the current account to tax shocks across the four countries. While the current account remains essentially unresponsive to unexpected tax cuts in Australia and the UK, it improves in Canada and deteriorates in the US. In contrast, the primary budget deficit worsens in all cases, implying that the twin-deficit hypothesis (conditional on a tax shock) is supported only by US data. We also find that the real exchange rate remains essentially unchanged following the tax cut in Australia and the UK, but that it appreciates significantly and persistently in Canada and the US. These findings are novel and have not been previously reported in the empirical literature. Importantly, they are generally at odds with the predictions of standard economic models, except in the US. Finally, we show that imposing the restrictions commonly used to identify tax shocks leads to important mis-measurements of their effects. For example, the identification schemes proposed by Kim and Roubini (2008) or Monacelli and Perotti (2010) counterfactually imply that unexpected tax cuts lead to a twin divergence and to a real depreciation in the US.

Regarding the effects of government spending shocks, our results also reveal the absence of a clear pattern regarding the reaction of the current account. In response to an unexpected increase in public spending, the current account deteriorates in the UK, improves with a delay in Canada and the US, and remains unchanged in Australia. For its part, the budget deficit shrinks with a delay in Australia and

are anticipated, Leeper et al. show that the resulting time series may have a non-invertible moving average component, such that it would be impossible to recover the true fiscal shocks from current and past variables. In, Bouakez, Chihi, and Normandin (2010), however, we provide suggestive evidence that the fiscal foresight problem is not sufficiently severe to undermine the SVAR approach. This is likely due to the fact that empirical studies mostly use quarterly data and that an important fraction of the changes in fiscal policy are implemented within a quarter, as documented in Mertens and Ravn (2010).

the UK and worsens in Canada and the US. Again, these findings lend little support to the twin-deficit hypothesis. As for the real exchange rate, it depreciates significantly in all countries, except Canada, where it exhibits a muted and statistically insignificant response. Interestingly, our results indicate that the magnitude of the real depreciation triggered by an unexpected increase in public spending is larger than what is found using the commonly used approaches, making the "exchange rate puzzle" even worse.

The rest of the paper is organized as follows. Section 2 presents the empirical methodology, including the identification strategy, the estimation method, and the data. Section 3 discusses the estimation results and the dynamic effects of tax and government spending shocks. Section 4 evaluates the robustness of the results to alternative detrending methods and to an alternative sample period. Section 5 concludes.

2 Empirical Methodology

2.1 Specification and Identification

Assume that the data are represented by the following SVAR:

$$Az_t = \sum_{i=1}^{m} A_i z_{t-i} + \epsilon_t, \tag{1}$$

where z_t is a vector of variables that includes output (y_t) , the price of bonds (q_t), government spending (g_t), taxes (τ_t), the real exchange rate (s_t) defined as the relative price of a foreign basket in terms of the domestic basket, and the current account (x_t); and ϵ_t is a vector of mutually uncorrelated structural innovations, which include fiscal shocks. Denote by ν_t the vector of residuals (or statistical innovations) obtained by projecting z_t on its own lags. These residuals are linked to the structural innovations through

$$A\nu_t = \epsilon_t, \tag{2}$$

where $A \equiv [a_{i,j}]_{i,j=1,\ldots,6}$ is the matrix that captures the contemporaneous interaction among the variables included in z_t. We cast fiscal policy in the context of a market for newly issued bonds. More specifically, we assume the following structure:

$$\nu_{b,t}^d = -\alpha \nu_{q,t} + \beta(\nu_{y,t} - \nu_{\tau,t}) + \gamma \nu_{s,t} + \sigma_d \epsilon_{d,t}, \quad (3)$$

$$\nu_{p,t} \equiv \nu_{g,t} - \nu_{\tau,t} = \nu_{q,t} + \nu_{b,t}^s, \quad (4)$$

$$\nu_{g,t} = \eta_g \nu_{y,t} + \theta_g \sigma_d \epsilon_{d,t} + \psi_g \sigma_\tau \epsilon_{\tau,t} + \sigma_g \epsilon_{g,t}, \quad (5)$$

$$\nu_{\tau,t} = \eta_\tau \nu_{y,t} + \theta_\tau \sigma_d \epsilon_{d,t} + \psi_\tau \sigma_g \epsilon_{g,t} + \sigma_\tau \epsilon_{\tau,t}. \quad (6)$$

Equation (3) is the private sector's demand for newly issued government bonds (Treasury bills), expressed in innovation form. This formulation extends the one proposed in Bouakez, Chihi, and Normandin (2010) by assuming that the demand for bonds, $\nu_{b,t}^d$, depends not only on the price of bonds, $\nu_{q,t}$, and on disposable income, $\nu_{y,t} - \nu_{\tau,t}$, but also on the real exchange rate, $\nu_{s,t}$, in order to capture the portion of demand originating in the rest of the world. In this equation, $\epsilon_{d,t}$ represents a demand shock and σ_d is a scaling parameter. The parameter α measures the (absolute value of the) slope of the demand curve, and is assumed to be different from 1. The parameters β and γ are the elasticities of this demand to disposable income and to the real exchange rate, respectively, and both are assumed to be positive.

Equation (4) is (an approximation of) the government's budget constraint, and states that the innovation in the primary deficit, $\nu_{p,t}$, (i.e., the difference between government spending and taxes) must be equal to the innovation in the value of debt, with $\nu_{b,t}^s$ being the supply of bonds. Note that because this constraint is expressed in innovation form, it does not include the payment for bonds that mature in period t (since those bonds were issued in period $t-1$).[33] Equations (5) and (6) describe the procedures followed by the government to determine fiscal spending and taxes. The disturbances $\epsilon_{g,t}$ and $\epsilon_{\tau,t}$ are the fiscal shocks that we aim to identify. The former is a shock to government spending and the latter is a tax shock. The terms σ_g and σ_τ are scaling parameters. Equation (5) states that government spending may change in response to changes in output or to demand and tax shocks. Equation (6) has an analogous interpretation for taxes. In these equations, the parameters η_g and η_τ measure the automatic and systematic responses of, respectively, government spending and taxes to changes in output. In this respect, η_g and η_τ do not necessarily coincide with the elasticities of fiscal

[33] For simplicity, this equation also abstracts from seignorage revenues, which have historically been small in industrialized countries.

variables with respect to output estimated by Blanchard and Perotti (2002), which capture only the automatic adjustment of government spending and taxes.

Imposing equilibrium in the bonds market and solving for the structural innovations, ϵ_t, in terms of the residuals, ν_t, yield

$$\begin{pmatrix} a_{11} & a_{12} & a_{13} & a_{14} & a_{15} & a_{16} \\ -\frac{\beta}{\sigma_d} & \frac{\alpha-1}{\sigma_d} & \frac{1}{\sigma_d} & \frac{\beta-1}{\sigma_d} & -\frac{\gamma}{\sigma_d} & 0 \\ \frac{\psi_g(\eta_\tau-\beta\theta_\tau)-(\eta_g-\beta\theta_g)}{\sigma_g(1-\psi_g\psi_\tau)} & \frac{(1-\alpha)(\theta_g-\theta_\tau\psi_g)}{\sigma_g(1-\psi_g\psi_\tau)} & \frac{1-\theta_g+\theta_\tau\psi_g}{\sigma_g(1-\psi_g\psi_\tau)} & \frac{(1-\beta)(\theta_g-\theta_\tau\psi_g)-\psi_g}{\sigma_g(1-\psi_g\psi_\tau)} & \frac{\gamma(\theta_g-\theta_\tau\psi_g)}{\sigma_g(1-\psi_g\psi_\tau)} & 0 \\ \frac{\psi_\tau(\eta_g-\beta\theta_g)-(\eta_\tau-\beta\theta_\tau)}{\sigma_\tau(1-\psi_g\psi_\tau)} & \frac{(1-\alpha)(\theta_\tau-\theta_g\psi_\tau)}{\sigma_\tau(1-\psi_g\psi_\tau)} & \frac{\psi_\tau(\theta_g-1)-\theta_\tau}{\sigma_\tau(1-\psi_g\psi_\tau)} & \frac{1+(1-\beta)(\theta_\tau-\theta_g\psi_\tau)}{\sigma_\tau(1-\psi_g\psi_\tau)} & \frac{\gamma(\theta_\tau-\theta_g\psi_\tau)}{\sigma_\tau(1-\psi_g\psi_\tau)} & 0 \\ a_{51} & a_{52} & a_{53} & a_{54} & a_{55} & a_{56} \\ a_{61} & a_{62} & a_{63} & a_{64} & a_{65} & a_{66} \end{pmatrix}$$

$$\times \begin{pmatrix} v_{y,t} \\ v_{q,t} \\ v_{g,t} \\ v_{\tau,t} \\ v_{s,t} \\ v_{x,t} \end{pmatrix} = \begin{pmatrix} \epsilon_{1,t} \\ \epsilon_{d,t} \\ \epsilon_{g,t} \\ \epsilon_{\tau,t} \\ \epsilon_{5,t} \\ \epsilon_{6,t} \end{pmatrix}, \tag{7}$$

where a_{ij} $(i = 1,5,6,\ j = 1,...,6)$ are unconstrained parameters. This specification imposes the following restrictions: $a_{26} = 0$, $a_{36} = 0$, $a_{46} = 0$, $a_{24} = -(a_{21} + a_{23})$, $\frac{a_{32}}{a_{22}} = \frac{a_{35}}{a_{25}}$, and $\frac{a_{42}}{a_{22}} = \frac{a_{45}}{a_{25}}$.[34]

The conditional scedastic structure of system (7) is:

$$\Sigma_t = A^{-1}\Gamma_t A^{-1'}, \tag{8}$$

where $\Sigma_t = E_{t-1}(\nu_t\nu_t')$ is the (non-diagonal) conditional covariance matrix of the statistical innovations and $\Gamma_t = E_{t-1}(\epsilon_t\epsilon_t')$ is the (diagonal) conditional covariance matrix of the structural innovations. Without loss of generality, the unconditional variances of the structural innovations are normalized to unity $(I = E(\epsilon_t\epsilon_t'))$. The dynamics of the conditional variances of the structural innovations are determined by

$$\Gamma_t = (I - \Delta_1 - \Delta_2) + \Delta_1 \bullet (\epsilon_{t-1}\epsilon_{t-1}') + \Delta_2 \bullet \Gamma_{t-1}. \tag{9}$$

The operator $\bullet$ denotes the element-by-element matrix multiplication, while Δ_1 and Δ_2 are diagonal matrices of parameters. Equation (9) involves intercepts that are consistent with the normalization $I = E(\epsilon_t\epsilon_t')$. Also, (9) implies that all the structural innovations are conditionally homoscedastic if Δ_1

[34] Note that the last two restrictions imply the redundant restriction $\frac{a_{42}}{a_{32}} = \frac{a_{45}}{a_{35}}$.

and Δ_2 are null. On the other hand, some structural innovations display time-varying conditional variances characterized by univariate generalized autoregressive conditional heteroscedastic [GARCH(1,1)] processes if Δ_1 and Δ_2 — which contain the ARCH and GARCH coefficients, respectively — are positive semi-definite and $(I - \Delta_1 - \Delta_2)$ is positive definite. Finally, all the conditional variances follow GARCH(1,1) processes if Δ_1, Δ_2, and $(I - \Delta_1 - \Delta_2)$ are positive definite.

Under conditional heteroscedasticity, system (7) can be identified, allowing us to study the effects of fiscal policy shocks. The sufficient (rank) condition for identification states that the conditional variances of the structural innovations are linearly independent. That is, $\lambda = 0$ is the only solution to $\Gamma\lambda = 0$, such that $(\Gamma'\Gamma)$ is invertible — where Γ stacks by column the conditional volatilities associated with each structural innovation. The necessary (order) condition requires that the conditional variances of (at least) all but one structural innovations are time-varying. In practice, the rank and order conditions lead to similar conclusions, given that the conditional variances are parameterized by GARCH(1,1) processes (see Sentana and Fiorentini 2001). For further discussion of the intuition underlying identification through conditional heteroscedasticity, see Bouakez, Chihi, and Normandin (2010) (the first essay of the these).

2.2 Identification Under Homoscedasticity: Existing Approaches

Under conditional homoscedasticity, 15 restrictions need to be imposed on the matrix A in order to achieve identification. These restrictions constrain the contemporaneous interaction of the variables of interest in a way that reflects the econometrician's judgment about the process by which policy variables are determined and/or the manner in which they affect certain variables. Existing approaches to identify fiscal-policy shocks within SVARs can be grouped into the following four categories, depending on the resulting shape of the A matrix.

Recursive scheme

This scheme implies a system in which the matrix A is a lower triangular:

$$\begin{pmatrix} \tilde{a}_{11} & 0 & 0 & 0 & 0 & 0 \\ \tilde{a}_{21} & \tilde{a}_{22} & 0 & 0 & 0 & 0 \\ \tilde{a}_{31} & \tilde{a}_{32} & \tilde{a}_{33} & 0 & 0 & 0 \\ \tilde{a}_{41} & \tilde{a}_{42} & \tilde{a}_{43} & \tilde{a}_{44} & 0 & 0 \\ \tilde{a}_{51} & \tilde{a}_{52} & \tilde{a}_{53} & \tilde{a}_{54} & \tilde{a}_{55} & 0 \\ \tilde{a}_{61} & \tilde{a}_{62} & \tilde{a}_{63} & \tilde{a}_{64} & \tilde{a}_{65} & \tilde{a}_{66} \end{pmatrix} \begin{pmatrix} v_{g,t} \\ v_{y,t} \\ v_{\mathcal{T},t} \\ v_{x,t} \\ v_{q,t} \\ v_{s,t} \end{pmatrix} = \begin{pmatrix} \epsilon_{g,t} \\ \epsilon_{2,t} \\ \epsilon_{\mathcal{T},t} \\ \epsilon_{4,t} \\ \epsilon_{5,t} \\ \epsilon_{6,t} \end{pmatrix}. \tag{10}$$

In this specification, government spending is predetermined with respect to any other variable in the system and thus government spending shocks can be obtained simply by a Cholesky decomposition of the covariance matrix of the VAR residuals, where public spending is ranked first. This is the strategy employed by Kim and Roubini (2008), Corsetti and Müller (2006), and Müller (2008) to identify the effects of government spending shocks on the current account and the exchange rate. Among the three studies, only the one by Corsetti and Müller (2006) used data from multiple countries, namely Australia, Canada, the UK, and the US; the two others having focused exclusively on the US.

The system above also implies that output is predetermined with respect to taxes. Thus, following a tax shock, the initial response of output is nil by construction. On the other hand, taxes may respond contemporaneously to unexpected changes in output, reflecting the automatic and systematic responses of government revenue to changes in economic activity. This strategy of ordering output before taxes in a Cholesky decomposition has only been performed by Kim and Roubini, whereas the two other studies cited above did not study the effects of tax shocks on external variables.

Non-recursive scheme (KR)

Kim and Roubini (2008) consider an alternative identification scheme whereby government spending is still predetermined with respect to all the remaining variables, but where the contemporaneous interaction of output and taxes is left unrestricted. In order to obtain this additional degree of freedom, however, a parametric restriction must be imposed elsewhere in the system. Kim and Roubini achieve this requirement by setting $\tilde{a}_{31} = 0$, which yields

$$\begin{pmatrix} \tilde{a}_{11} & 0 & 0 & 0 & 0 & 0 \\ \tilde{a}_{21} & \tilde{a}_{22} & \tilde{a}_{23} & 0 & 0 & 0 \\ 0 & \tilde{a}_{32} & \tilde{a}_{33} & 0 & 0 & 0 \\ \tilde{a}_{41} & \tilde{a}_{42} & \tilde{a}_{43} & \tilde{a}_{44} & 0 & 0 \\ \tilde{a}_{51} & \tilde{a}_{52} & \tilde{a}_{53} & \tilde{a}_{54} & \tilde{a}_{55} & 0 \\ \tilde{a}_{61} & \tilde{a}_{62} & \tilde{a}_{63} & \tilde{a}_{64} & \tilde{a}_{65} & \tilde{a}_{66} \end{pmatrix} \begin{pmatrix} v_{g,t} \\ v_{y,t} \\ v_{\mathcal{T},t} \\ v_{x,t} \\ v_{q,t} \\ v_{s,t} \end{pmatrix} = \begin{pmatrix} \epsilon_{g,t} \\ \epsilon_{2,t} \\ \epsilon_{\mathcal{T},t} \\ \epsilon_{4,t} \\ \epsilon_{5,t} \\ \epsilon_{6,t} \end{pmatrix}.$$

Non-recursive scheme (MP)

Monacelli and Perotti (2010) also consider an alternative non-recursive scheme that does not impose any prior ordering between taxes and output, assuming that the two variables are simultaneously determined. However, in contrast to KR, they leave unrestricted the parameter $\tilde{a}_{31}$. Since such an assumption implies an additional parameter to estimate, Monacelli and Perotti follow the strategy originally proposed by Blanchard and Perotti (2002) of calibrating the elasticity of taxes with respect to output based on institutional information. More specifically, this elasticity measures the automatic adjustment of taxes to changes in output. In terms of our notation, such a specification can be written as

$$\begin{pmatrix} \tilde{a}_{11} & 0 & 0 & 0 & 0 & 0 \\ \tilde{a}_{21} & \tilde{a}_{22} & \tilde{a}_{23} & 0 & 0 & 0 \\ \tilde{a}_{31} & -\phi\tilde{a}_{33} & \tilde{a}_{33} & 0 & 0 & 0 \\ \tilde{a}_{41} & \tilde{a}_{42} & \tilde{a}_{43} & \tilde{a}_{44} & 0 & 0 \\ \tilde{a}_{51} & \tilde{a}_{52} & \tilde{a}_{53} & \tilde{a}_{54} & \tilde{a}_{55} & 0 \\ \tilde{a}_{61} & \tilde{a}_{62} & \tilde{a}_{63} & \tilde{a}_{64} & \tilde{a}_{65} & \tilde{a}_{66} \end{pmatrix} \begin{pmatrix} v_{g,t} \\ v_{y,t} \\ v_{\tau,t} \\ v_{x,t} \\ v_{q,t} \\ v_{s,t} \end{pmatrix} = \begin{pmatrix} \epsilon_{g,t} \\ \epsilon_{2,t} \\ \epsilon_{\tau,t} \\ \epsilon_{4,t} \\ \epsilon_{5,t} \\ \epsilon_{6,t} \end{pmatrix}, \tag{12}$$

where ϕ is the elasticity of taxes with respect to output. Monacelli and Perotti apply this scheme to measure the effects of government spending shocks on the current account and the exchange rate in Australia, Canada, the UK, and the US, though a special attention is paid to the latter country. It is worth emphasizing, however, that these responses are identical to those that would be obtained from the recursive or the KR schemes. Only in the case of tax shock would these three approaches imply different results.

Sign restrictions

An alternative identification strategy to pin down the effects of government spending shock is the so-called sign restriction approach, which identifies the elements of A such that the impulse responses of interest satisfy a number of shape and sign restrictions imposed by the econometrician. Enders, Müller, and Scholl (2011) apply this methodology to measure the effects of government spending shocks on the current account and the exchange rate. Their identification assumptions ensure that the following restrictions are satisfied in response to a positive government spending shock : (i) public spending increases during the first four quarters after the shock, (ii) the primary budget deficit increases for four quarters, (iii) output increases for two quarters, (iv) investment increases for six quarters, (v)

the nominal interest rate increases for four quarters, and (vi) inflation increases immediately after the shock. The response of the current account and the exchange rate, on the hand, are left unrestricted.

2.3 Estimation Method and Data

The elements of A, Δ_1, and Δ_2 are estimated using the following two-step procedure. We first estimate by ordinary least squares an $m-$order VAR that includes output, the price of bonds, the current account, the real exchange rate, government spending and taxes,[35] and extract the implied residuals, ν_t, for $t = m+1, ..., T$. For given values of the elements of the matrices A, Δ_1, and Δ_2, it is then possible to construct an estimate of the conditional covariance matrix Σ_t recursively, using equations (8) and (9) and the initialization $\Gamma_m = \epsilon_m \epsilon_m' = I$. Assuming that the residuals are conditionally normally distributed, the second step consists in selecting the elements of the matrices A, Δ_1, and Δ_2 that maximize the likelihood of the sample.

We use quarterly data covering the period 1973-1 to 2008-4. The analysis is performed for Australia, Canada, the UK, and the US. The choice of this sample of countries is mainly motivated by the availability of non-interpolated quarterly data on fiscal variables at the general government level. The series used in estimation are constructed as follows. Output is measured by real GDP. The price of bonds is measured by the inverse of the gross real return on 3-month treasury bills,[36] where the GDP deflator is used to deflate the gross nominal return. The current account is defined as the change in net foreign assets and is expressed as a fraction of GDP, and the exchange rate is measured by the real effective exchange rate, which is constructed such that an increase corresponds to a real depreciation. Government spending is defined as the sum of federal (defense and non-defense), state and local consumption and gross investment expenditures. Taxes are defined as total government receipts less net transfer payments. The spending and tax series are expressed in real terms using the GDP deflator. Output, government spending and taxes are divided by total population and all the series, except the current account to output ratio, are expressed in logarithm. The data sources and further details on the construction of the series are provided in the Appendix.

[35] The benchmark specification includes a constant, a quadratic trend, and four lags.
[36] We found the results to be robust when we measure the price of bonds using the return on 10-year treasury bonds.

3 Results

This section discusses the estimation and test results, as well as the dynamic responses to tax and government spending shocks implied by (7). It also compares these responses to those obtained by imposing the identifying restrictions commonly used in the literature.

3.1 Parameter Estimates and Specification Test

For each country, we estimate a 4-order VAR ($m = 4$). Table 2.I reports the p-values associated with the McLeod-Li test statistic applied to the squared VAR residuals. In the vast majority of cases, the test rejects the null hypothesis of absence of autocorrelation in the squared VAR residuals at 1, 2 and 4 lags. This result hints to the presence of conditional heteroscedasticity in the statistical innovations, which is likely to translate into time-varying conditional variances of the structural innovations.

Table 2.II presents the estimates of the GARCH(1,1) parameters. For each country, the estimates indicate that the conditional variances of (at least) five structural innovations are time-varying, and that the conditional variances of the structural innovations are linearly independent, thus satisfying the order (necessary) and rank (sufficient) conditions for the identification of system(7). The table also shows that government spending shocks exhibit a conditional volatility that is moderately persistent for Australia and Canada, but highly persistent for the UK and the US– where the persistence is measured by the sum of the ARCH and GARCH coefficients. On the other hand, the conditional volatility of tax shocks is highly persistent for all the countries except the US. A more telling representation of these conditional variances is provided by Figure 2.1. The figure shows important time variation in the conditional variances of both fiscal and non-fiscal shocks, which often display alternating episodes of high and low volatility. These results corroborate the findings of earlier studies that documented the presence of conditional volatility in the time series of output (Fountas and Karanasos 2007), the nominal interest rate (Garcia and Perron 1996; Den Haan and Spear 1998; Fernandèz-Villaverde et al. 2010), the exchange rate (Hsieh 1988, 1989, Engel and Hamilton 1990, Engel and Kim 1999), and fiscal variables (Fernandèz-Villaverde et al. 2011).

Does the GARCH(1,1) specification provide an adequate description of the process that governs the conditional variances of the structural innovations? To answer this question, we test whether there is any autocorrelation in the ratio of the squared structural innovations relative to their conditional variances. The Mcleod-Li test results, reported in Table 2.III, indicate that the null hypothesis of no autocorrelation cannot be rejected at any conventional level of significance for 1, 2 and 4 lags. This suggests that the GARCH(1,1) process is well specified.

Next, we turn to the estimates of the structural (bond-market) parameters, which we report in Table 2.IV. The estimates of α indicate that the slope of the demand for newly issued government bonds is negative and statistically significant for all countries. The estimates of β are positive and statistically significant in all cases, indicating a positive relation between the demand for bonds and disposable income. The elasticity of demand for bonds with respect to the real exchange rate, γ, is precisely estimated only for the US, but has the expected sign in all cases. The parameters measuring the automatic/systematic responses of government spending and taxes to output, η_g and η_τ respectively, are statistically significant for Canada and the US. The point estimates of η_τ for these two countries are substantially larger than the elasticity estimated by Blanchard and Perotti (2002) for the US, thus indicating that the systematic response of taxes to changes in output is quantitatively important. The parameters θ_g and θ_τ are mostly statistically significant, whereas the opposite is true for ψ_g and ψ_τ. Finally, the scaling factor of government spending shocks, σ_g, is smaller than that of tax shocks, σ_τ.

The parametric restrictions implied by our model, i.e., $a_{26} = 0$, $a_{36} = 0$, $a_{46} = 0$, $a_{24} = -(a_{21}+a_{23})$, $\frac{a_{32}}{a_{22}} = \frac{a_{35}}{a_{25}}$, and $\frac{a_{42}}{a_{22}} = \frac{a_{45}}{a_{25}}$, are tested using a Wald test. The p-values associated with the test statistic, reported in Table 2.V, indicate that these restrictions cannot be rejected at any conventional significance level for Australia, the UK, and the US. For Canada, these restrictions cannot be rejected only at the 4 percent (or lower) significance level. Since system (7) appears to be generally supported by the data, we henceforth refer to it as the unrestricted system and to its implications as the unrestricted ones.

3.2 Dynamic Effects of Tax Shocks

Figure 2.2 depicts the dynamic effects of an unexpected tax cut on output, the primary budget deficit, the current account and the real exchange rate. The first observation that emerges from this figure is that there is, in general, a similarity in results between Australia and the UK on the one hand, and Canada and the US on the other hand. Notwithstanding that tax cut is much less persistent in Canada and the US than in Australia and the UK, it leads to a persistent and statistically significant increase in output in the former countries, whereas in the latter the output response is muted on impact and mostly statistically insignificant. The negative tax shock deteriorates the primary budget deficit in all four countries, but the effect is larger and much more persistent in Australia and the UK than in Canada and the US.

In contrast, the response of the current account in the former two countries is flat and indistinguishable from zero. Hence, there is no evidence of twin deficits or twin divergence conditional on tax shocks for these two countries. On the other hand, the tax cut improves the current account in Canada, thus moving budget and external deficits in opposite directions–twin divergence. The opposite scenario occurs in the US, where the tax cut worsens both the budget deficit and the current account-twin deficits. Therefore, there is no overwhelming evidence that, in a response to a tax shock, budget and external deficits move in tandem. In addition, these results provide little support to the hypothesis that the likelihood and magnitude of twin deficits increase with the degree of openness of an economy (see Corsetti and Müller 2006). Finally, Figure 2.2 shows that the real exchange rate is unresponsive, in a statistical sense, to the tax cut in Australia and the UK, but that it appreciates significantly in Canada and the US, although in the latter case, the exchange rate response ceases to be significant six quarters after the shock. These results constitute the first novelty of the present paper, as no empirical evidence exists about the effects of tax shocks on external variables in countries other than the US. Importantly, we find that the US is an outlier inasmuch as it is the only case where the effects of unexpected tax cuts are generally consistent with the predictions of standard economic models.

How do these results compare with those obtained by imposing the identifying restrictions used in earlier studies? Answering this question enables one to assess whether or not and to what extent those restrictions are innocuous. To conserve space, we restrict the comparison to the case of the US. Figure 2.3 superimposes on the unrestricted responses obtained for the US those implied by the recursive identification scheme discussed in Section 2 and by the two non recursive schemes employed by Kim and Roubini (KR) and Monacelli and Perotti (MP).[37] In all cases, the system is estimated under the assumptions of conditional heteroscedasticity, so that any difference in results between the unrestricted and restricted systems would be solely attributed to the parametric restrictions on the coefficients of the matrix A. The figure shows that the three sets of identifying restrictions lead to important counterfactual implications. First, both the recursive and MP schemes severely understate the output response, predicting that it is essentially nil at all horizons, whereas the KR scheme implies that output actually falls in a response to a tax cut. Second, the three restricted systems imply that the unanticipated decrease in taxes worsens the budget primary deficit and improves the current account in the US, which contradicts the twin-deficit result obtained under the unrestricted specification. Finally, the tax cut leads to a real depreciation of the US dollar under the three alternative identification schemes, whereas the unrestricted system predicts a real appreciation.[38] These findings clearly show that imposing arbitrary parametric restrictions in order to achieve identification can lead to mistaken inference about a country's external adjustment to tax shocks.

3.3 Dynamic Effects of Government Spending Shocks

The impulse responses to an expected increase in government spending shock are illustrated in Figure 2.4. The shock is expansionary in all four countries, leading to a persistent and statistically significant increase in output, except in the US, where the positive effect on output becomes statistically insignificant five quarters after the shock. The increase in government spending deteriorates the primary budget deficit in Canada and the US, and improves it in Australia and the UK,[39] although in the latter case, the effect is mostly statistically insignificant. The current account remains unresponsive in Australia, improves in Canada and the US, and deteriorates in the UK. Thus, conditional on a government spending shock, there is stronger evidence of twin divergence than twin deficits. Again, we find little support for the hypothesis that twin deficits are more likely to occur in more open economies. Finally, Figure 2.4 indicates that the real exchange rate depreciates in a response to an unexpected increase in public spending, except in Canada, where the response is muted and statistically insignificant. This depreciation contradicts the predictions of standard open-economy models.

[37] These authors also focus on the US.

[38] The results obtained under the KR identification scheme are consistent with those reported in Kim and Roubini (2008), which are based on a shorter sample period.

[39] It is possible to obtain an improvement in the budget deficit following an expansionary public spending shock because our specification allows for an endogenous adjustment in taxes following such a shock, whereas earlier approaches restrict the initial response of taxes to be nil.

Figure 2.5 compares the results for the US with those obtained from the identification schemes used in existing studies, namely the recursive and sign-restriction approaches. Note that the dynamic responses to a government spending shock implied by the KR and MP are identical to those implied by the recursive approach, since all of these systems assume that government spending is predetermined with respect to any other variable and impose the same number of exclusion restrictions. In implementing the sign-restriction approach, we imposed the following restrictions on the dynamic responses to a positive government spending shock: (*i*) government spending increases for 4 quarters, (*ii*) the primary budget deficit (as a fraction of output) worsens for four quarters, (*iii*) output increases for two quarters, and (*iv*) the real price of bonds falls on impact.[40]

At short horizons, the results obtained from the recursive and sign-restriction approaches regarding the response of the budget deficit and the current account to a government spending shock are generally similar to those obtained from the unrestricted specification. All three approaches predict a worsening of the budget deficit and an improvement of the current account in the US in response to an expansionary spending shock. At longer horizons, however, the two alternative approaches underestimate the response of the current account. More important discrepancies exist when it comes to the response of the real exchange rate. While the recursive approach yields a real depreciation, the latter is much smaller in magnitude than that predicted by the unrestricted system, especially at short horizons (up to two years). The sign-restriction approach, on the other hand, predicts that the median exchange rate response is very small in magnitude and changes sign during the first 10 quarters after the shock, but that there is so much uncertainty about such a response, that one cannot in fact reject the hypothesis that it is actually nil. Together, the results imply that the "real exchange rate puzzle" is worse than one may think based on traditional approaches.

4 Robustness Analysis

We now study the robustness of the results to alternative detrending methods and to an alternative sample period. Recall that the benchmark results discussed so far were obtained from a system in which

[40] These restrictions are very similar to those imposed by Enders, Müller, and Scholl (2011), though not exactly the same. The reason is that our estimated system differs slightly from theirs. The dynamic responses we obtain using this approach are nonetheless remarkably similar to those reported by these authors.

variables are expressed as deviations from a quadratic trend, and which is estimated over the post-1973 period. In this section, we report results based on systems in which variables are expressed (*i*) in levels, (*ii*) as deviations from a linear trend, (*iii*) in first differences (except the current account and the real exchange rate, which are expressed in levels). We also estimate the system (with quadratically detrended data) for the post-1980 period, given that some studies suggest the presence of a structural break around the year 1980 (see Perotti 2005). We again restrict our attention to the US and report the results in Figure 2.6 for the case of a tax shock, and in Figure 2.7 for the case of a government spending shock. In the case where the data is expressed in first differences, the reported responses are those of the variables in levels and are obtained by cumulating the responses of the variables in first differences.

In general, the responses to a tax shock obtained under the alternative detrending methods are fairly similar to (and often statistically indistinguishable from) the benchmark responses, especially at short horizons. The only exceptions are the responses of output when the variables are expressed in levels and as deviations from a linear trend. On the other hand, the responses obtained for the post-1980 period are relatively smaller in magnitude than those pertaining to the entire sample period, although the wedge is generally not significantly large. An even stronger similarity in results between the benchmark and the alternative estimations is observed in the case of a government spending shock. The only notable difference concerns the response of the real exchange rate, which is smaller in magnitude in the post-1980 period than when the entire sample period is used in estimation.

To summarize, this robustness check confirms the message conveyed by the benchmark analysis regarding the adjustment of the US current account and exchange rate to fiscal-policy shocks: A surprise tax cut deteriorates the current account and appreciates the real exchange rate, whereas a surprise increase in public spending improves the current account with a delay and depreciates the real exchange rate.

5 Conclusion

This paper has investigated the effects of fiscal policy shocks on the current account and the exchange rate using an empirical methodology that relaxes the commonly used identifying assumptions, and which instead achieves identification by exploiting the conditional heteroscedasticity of the structural shocks within an SVAR.

Notwithstanding that the effects of fiscal policy shocks are not always consistent across the four countries included in our sample, we found some similarities between Australia and the UK on the one hand, and Canada and the US on the other hand. More importantly, we found little support for the twin-deficit hypothesis regardless of the underlying fiscal shock. We also found that the effects of unexpected tax cuts are generally at odds with standard economic theory, except for the US. Finally, our results indicate that unexpected increases in public spending depreciates the currency in real terms in all but one country (Canada). While this puzzling depreciation (from the perspective of standard open-economy models) has also been documented by other studies, our results indicate that those studies severely understate the magnitude of the exchange rate response, thus suggesting that the "exchange rate puzzle" is worse than one might think based on traditional identification approaches.

6 References

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7 Appendix

7.1 Appendix A: Data Construction and Sources

This appendix describes the data used in this paper. The sample covers the 1973-1 to 2008-4 period for Australia, Canada, the UK, and the US. For Australia and the UK, the data are taken from the International Financial Statistics (IFS) released by the International Monetary Funds, the Main Economic Indicators (MEI) and Economic Outlook (EO) released by the Organization for Economic Cooperation and Development, and from Datastream. Data for Canada are collected from the databases released by Statistics Canada (SC), while data for the US are taken from the National Income and Products Accounts (NIPA), the Federal Reserve Bank of Saint-Louis' Fred database (FRED), and the Federal Reserve Statistical Releases (FRSR).

Output is measured by the nominal GDP (sources: EO for Australia and the UK, SC for Canada, and NIPA for the US) normalized by the GDP deflator (sources: EO for Australia and the UK, SC for Canada, and NIPA for the US). The price of bonds is constructed as the inverse of the gross real return, where the GDP deflator is used to deflate the gross nominal return. The nominal return is measured by the 90 day commercial bill rate for Australia (source: MEI), the 3-month treasury bill rate for Canada (source: SC), the UK (source: IFS), and the US (source: FRED). Except for the US, the exchange rate is defined as the consumer price index-based real effective exchange rate (source: MEI). For the US, the exchange rate is measured by the trade-weighted real exchange rate index against major currencies (source: FRSR). The current account (sources: EO for Australia and the UK, SC for Canada, and NIPA for the US) is expressed as a percentage of GDP. Government expenditures are measured by the sum of consumption and gross investment expenditures of the general government (sources: EO for Australia and the UK, SC for Canada, and NIPA for the US) normalized by the GDP deflator. Taxes are defined as total receipts of the general government less net transfers (sources: EO for Australia and the UK, SC for Canada, and NIPA for the US) normalized by GDP deflator. Output, government spending and taxes are expressed in per capita terms by dividing them by total population (sources: Datastream for Australia and the UK, SC for Canada and FRED for the US). Output, government spending, taxes, the price of bonds and the exchange rate are expressed in logarithm.

7.2 Appendix B: Figures

Figure 2.1: Conditional Variances of the Structural Shocks

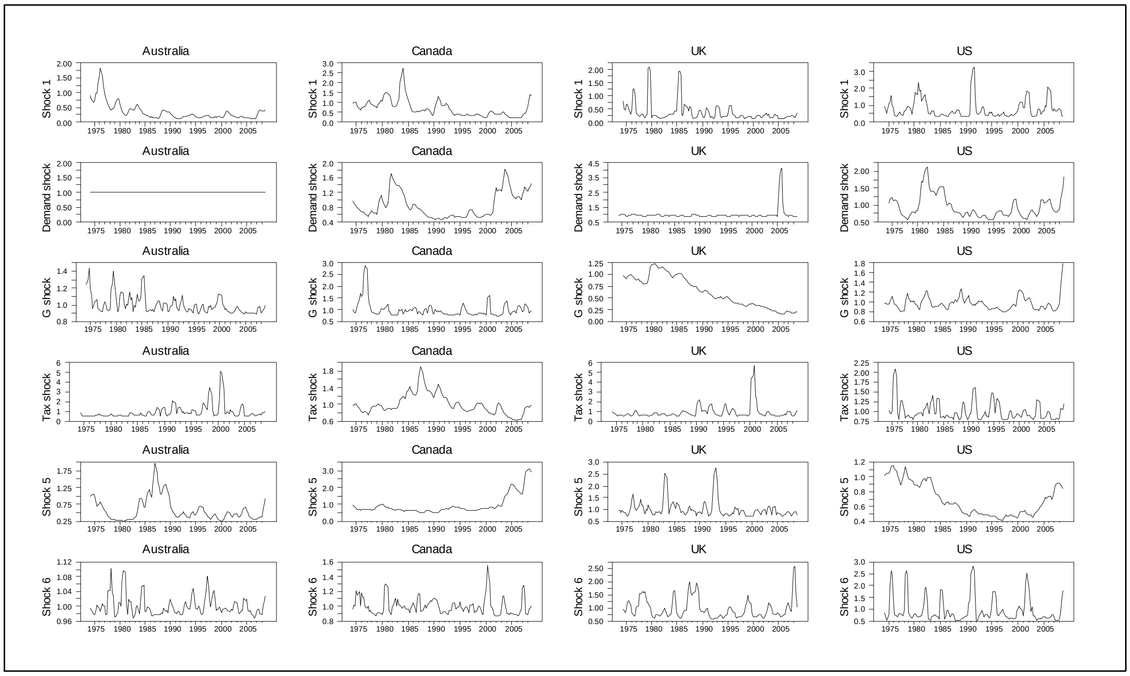
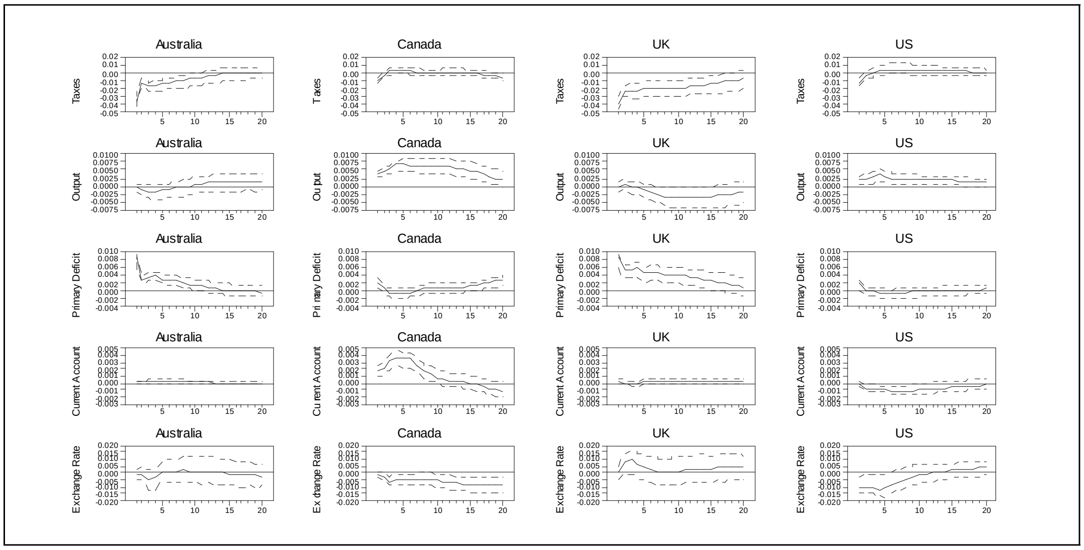
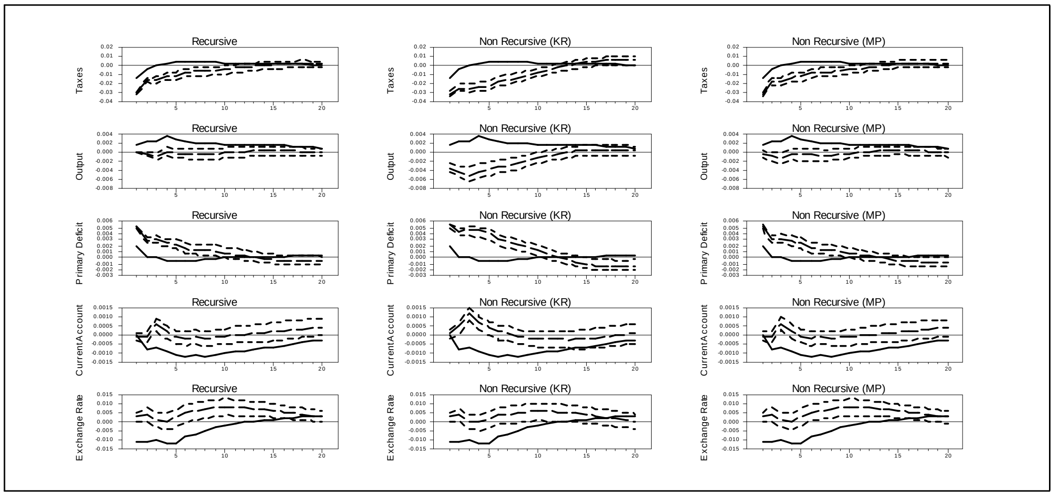


Figure 2.2: Unrestricted Dynamic Responses to a Negative Tax Shock



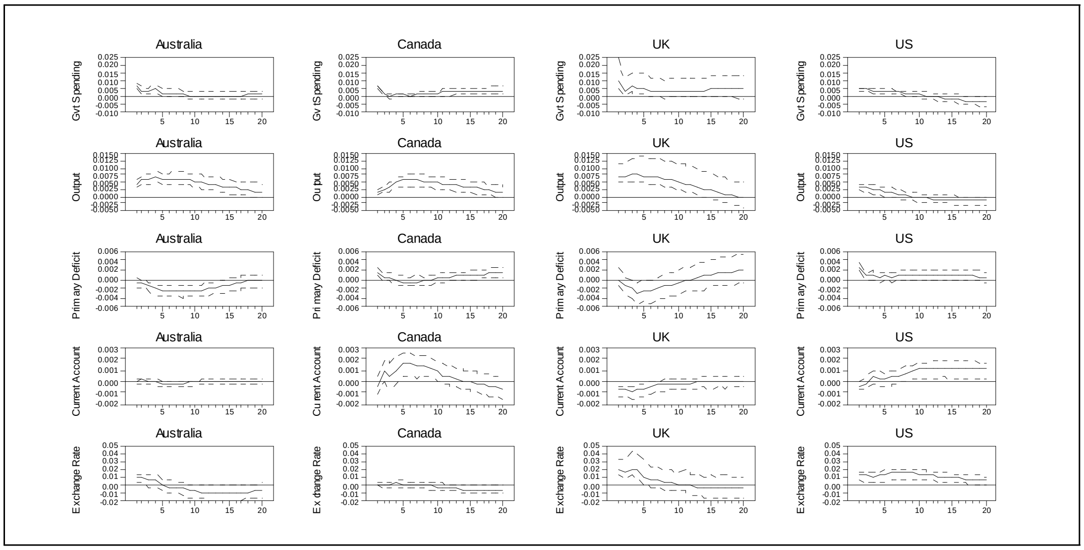
Notes: The solid lines correspond to the dynamic responses to a negative tax shock extracted from the unrestricted system for each country. The dotted lines are the 68 percent confidence intervals computed using the Sims-Zha (1999) Bayesian procedure.

Figure 2.3: Dynamic Responses to a Negative Tax Shock: Alternative Identification Schemes



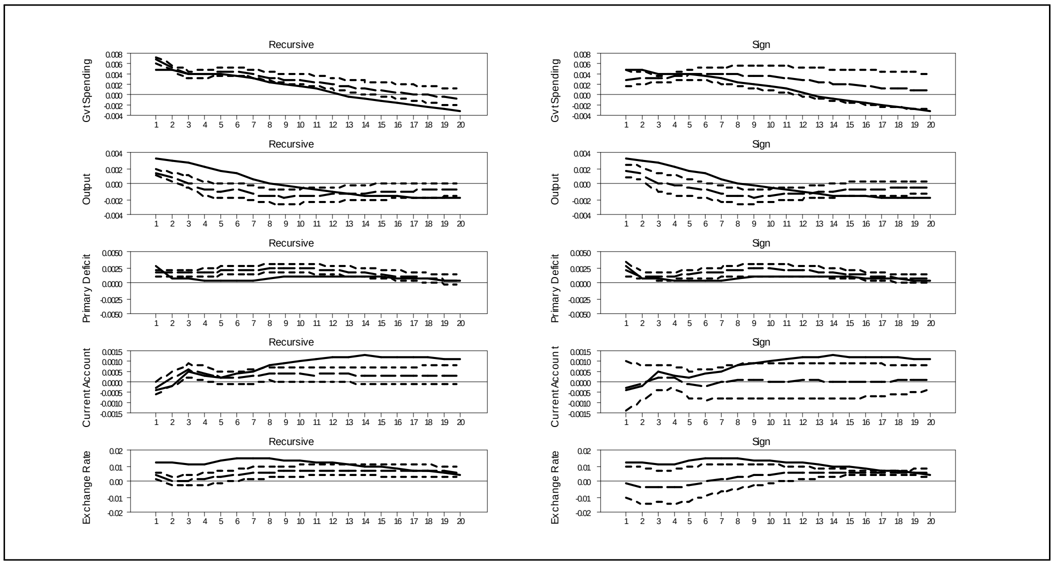
Notes: The solid (dashed) lines correspond to the dynamic responses to a negative tax shock extracted from the unrestricted (alternative) system for each country. The dotted lines are the 68 percent confidence intervals computed using the Sims-Zha (1999) Bayesian procedure.

Figure 2.4: Unrestricted Dynamic Responses to a Positive Government Spending Shock



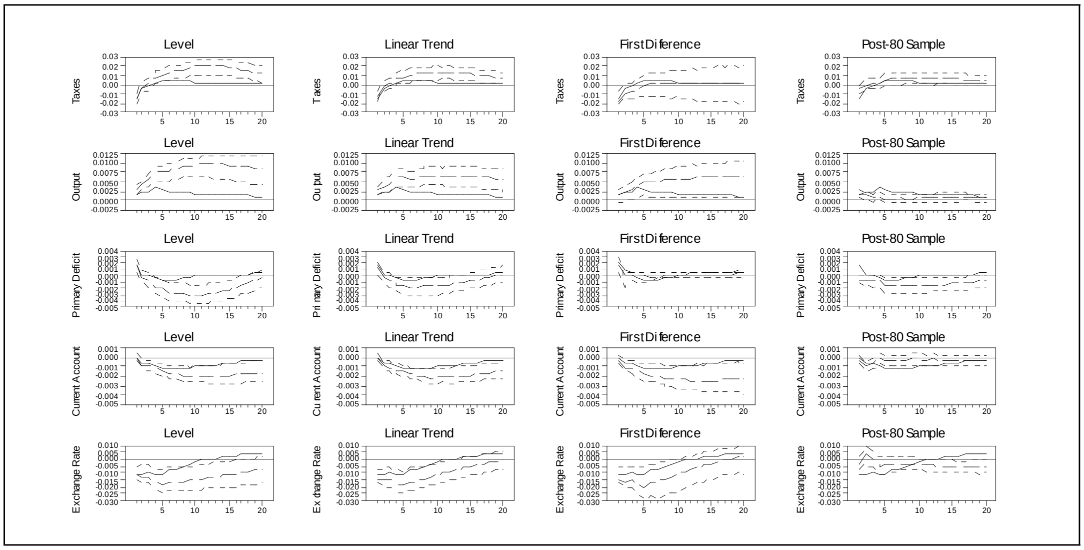
Notes: The solid lines correspond to the dynamic responses to a positive government spending shock extracted from the unrestricted system for each country. The dotted lines are the 68 percent confidence intervals computed using the Sims-Zha (1999) Bayesian procedure.

Figure 2.5: Dynamic Responses to a Government Spending Shock: Alternative Identification Schemes



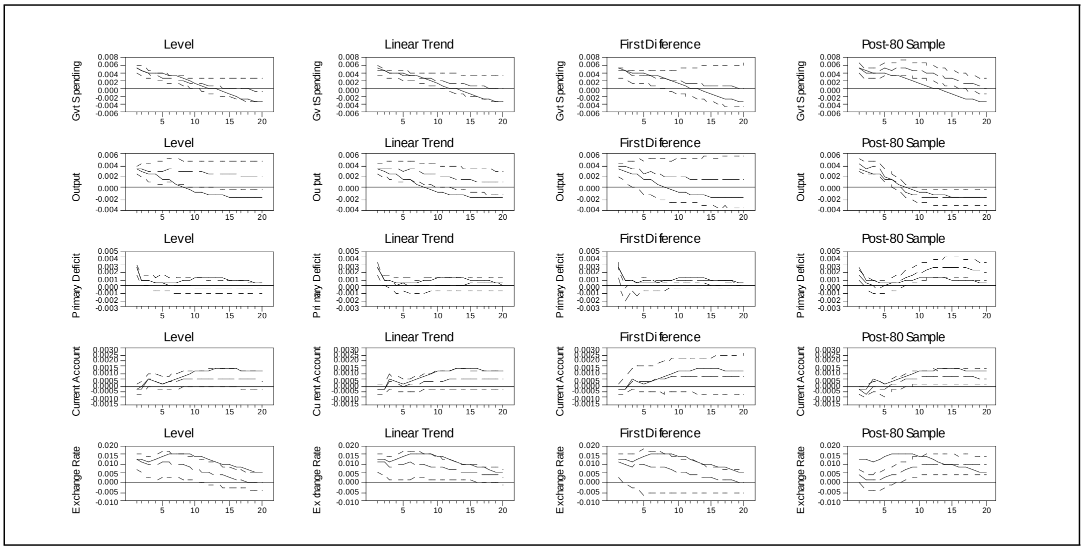
The solid (dashed) lines correspond to the dynamic responses to a positive government spending shock extracted from the unrestricted (alternative) system for each country. The dotted lines are the 68 percent confidence intervals computed using the Sims-Zha (1999) Bayesian procedure for the recursive case and the 68 percent intervals of the admissible dynamic responses for the sign-restriction case.

Figure 2.6: Dynamic Responses to a Negative Tax Shock: Robustness Analysis



The solid lines correspond to the dynamic responses extracted from the unrestricted system for the US. The long dashes correspond to the responses computed using alternative detrending methods and an alternative sample period. The short dashes are the 68 percent confidence intervals computed using the Sims-Zha (1999) Bayesian procedure.

Figure 2.7: Dynamic Responses to a Positive Government Spending Shock: Robustness Analysis



The solid lines correspond to the dynamic responses extracted from the unrestricted system for the US. The long dashes correspond to the responses computed using alternative detrending methods and an alternative sample period. The short dashes are the 68 percent confidence intervals computed using the Sims-Zha (1999) Bayesian procedure.

7.3 Appendix C: Tables

Table 2.I. Heteroscedasticity Test Results

	Lag	Australia	Canada	UK	US
$\nu_{y,t}^2$	1	0.043	0.988	0.003	0.025
	2	0.000	0.777	0.013	0.040
	4	0.000	0.623	0.032	0.159
$\nu_{q,t}^2$	1	0.129	0.725	0.017	0.862
	2	0.192	0.014	0.054	0.046
	4	0.027	0.083	0.212	0.012
$\nu_{g,t}^2$	1	0.038	0.055	0.021	0.077
	2	0.012	0.009	0.061	0.032
	4	0.029	0.021	0.226	0.075
$\nu_{\tau,t}^2$	1	0.008	0.014	0.481	0.003
	2	0.014	0.044	0.024	0.012
	4	0.059	0.038	0.000	0.045
$\nu_{s,t}^2$	1	0.680	0.111	0.014	0.459
	2	0.041	0.141	0.015	0.605
	4	0.029	0.000	0.053	0.406
$\nu_{x,t}^2$	1	0.578	0.050	0.326	0.000
	2	0.066	0.076	0.443	0.001
	4	0.055	0.151	0.636	0.003

Notes: Entries are the p-values associated with the McLeod-Li test statistic applied to the squared VAR residuals.

Table 2.II. Estimates of the GARCH(1,1) Parameters

	Australia	Canada	UK	US
$\epsilon_{1,t}$	0.293 (0.189)	0.250 (0.135)	0.881 (0.225)	0.596 (0.278)
	0.690 (0.194)	0.703 (0.152)	–	0.163 (0.202)
$\epsilon_{d,t}$	–	0.110 (0.079)	0.098 (0.288)	0.159 (0.162)
	–	0.848 (0.127)	0.032 (1.536)	0.721 (0.314)
$\epsilon_{g,t}$	0.116 (0.173)	0.273 (0.189)	0.063 (0.070)	0.095 (0.152)
	–	–	0.937 (0.083)	0.621 (0.798)
$\epsilon_{\tau,t}$	0.504 (0.194)	0.072 (0.101)	0.408 (0.236)	0.212 (0.172)
	–	0.856 (0.224)	0.135 (0.220)	–
$\epsilon_{5,t}$	0.244 (0.098)	0.089 (0.082)	0.305 (0.165)	0.062 (0.071)
	0.705 (0.122)	0.889 (0.124)	–	0.927 (0.085)
$\epsilon_{6,t}$	0.031 (0.140)	0.119 (0.145)	0.304 (0.184)	0.499 (0.286)
	–	0.084 (1.026)	0.318 (0.319)	–

Notes: Entries are the estimates (standard errors) of the parameters of the GARCH(1,1) processes. For each structural innovation, the first and second rows refer to the ARCH and GARCH coefficients, respectively. A dash (–) indicates that zero-restrictions are imposed to ensure that Δ_1 and Δ_2 are non-negative definite.

Table 2.III. Specification Test Results

	Lag	Australia	Canada	UK	US
$\epsilon_{1,t}^2$	1	0.804	0.774	0.484	0.407
	2	0.962	0.921	0.339	0.665
	4	0.928	0.996	0.626	0.718
$\epsilon_{d,t}^2$	1	0.916	0.782	0.908	0.966
	2	0.974	0.955	0.992	0.847
	4	0.961	0.971	0.999	0.590
$\epsilon_{g,t}^2$	1	0.974	0.704	0.821	0.499
	2	0.113	0.818	0.864	0.543
	4	0.171	0.722	0.857	0.681
$\epsilon_{\tau,t}^2$	1	0.826	0.662	0.724	0.995
	2	0.523	0.275	0.648	0.534
	4	0.367	0.279	0.794	0.583
$\epsilon_{5,t}^2$	1	0.145	0.595	0.531	0.238
	2	0.634	0.533	0.563	0.201
	4	0.589	0.487	0.374	0.483
$\epsilon_{6,t}^2$	1	0.959	0.848	0.802	0.786
	2	0.978	0.942	0.965	0.882
	4	0.929	0.748	0.993	0.591

Notes: Entries are the p-values associated with the McLeod-Li test statistic applied to the squared structural innovations relative to their conditional variances.

Table 2.IV. Estimates of the Structural Parameters

Parameter	Australia	Canada	UK	US
α	0.844 (0.186)	0.549 (0.243)	1.066 (0.195)	1.099 (0.168)
β	1.193 (0.120)	0.919 (0.132)	0.986 (0.294)	0.884 (0.111)
γ	0.073 (0.097)	0.012 (0.104)	0.056 (0.165)	0.267 (0.104)
η_g	0.484 (0.480)	−0.237 (0.741)	0.602 (1.230)	0.515 (0.258)
η_τ	1.134 (1.197)	12.038 (6.393)	1.152 (1.914)	5.783 (2.027)
θ_g	0.732 (0.159)	0.425 (0.192)	1.001 (0.153)	0.371 (0.179)
θ_τ	1.176 (0.192)	1.001 (0.726)	0.490 (0.320)	−0.076 (0.859)
ψ_g	−0.186 (0.114)	−0.037 (0.056)	0.017 (0.301)	0.153 (0.124)
ψ_τ	1.372 (1.818)	−2.571 (2.472)	0.428 (2.734)	−8.774 (6.874)
σ_d	0.019 (0.003)	0.018 (0.007)	0.027 (0.005)	0.011 (0.005)
σ_g	0.004 (0.005)	0.007 (0.002)	0.005 (0.017)	0.003 (0.002)
σ_τ	0.036 (0.007)	0.056 (0.027)	0.040 (0.009)	0.025 (0.012)

Notes: Numbers between parentheses are standard errors.

Table 2.V. Test of the Parametric Restrictions

	Australia	Canada	UK	US
P-value	0.619	0.040	0.948	0.612

Note: Entries are the p-values of the χ^2-distributed Wald test statistic associated with the restrictions $a_{26} = 0$, $a_{36} = 0$, $a_{46} = 0$, $a_{24} = -(a_{21} + a_{23})$, $\frac{a_{32}}{a_{22}} = \frac{a_{35}}{a_{25}}$, and $\frac{a_{42}}{a_{22}} = \frac{a_{45}}{a_{25}}$.

Essay 3. External and Budget Deficits in Some Developing Countries

Abstract

This paper documents and explains the positive comovement between the external and budget deficits of developing countries for which post-1960 time-series data are available. First, the estimates indicate that the empirical covariance between these deficits is always positive and is statistically significant for many cases. This is consistent with previous findings obtained from panel regressions. Second, the empirical covariance is close to that predicted from a tractable small open-economy, overlapping-generation model with heterogeneous goods. Also, the predicted covariance is induced by shocks which are closely related to internal conditions such as domestic resources and fiscal policies, and to a much lesser extent to external conditions such as the world interest rate, real exchange rate, and terms of trade. This analysis explaining the joint behavior of the external and budget deficits contrasts with earlier single-equation studies characterizing the individual behavior of either the external deficit or budget deficit.

JEL classification: E62, F32, F41.

Keywords: Covariance Decomposition, Dynamic Responses, Internal and External Conditions, Vector Autoregressions, Small-Open Economy, Overlapping-Generation Model with Heterogeneous Goods.

1 Introduction

The objectives of this paper are twofold. First, we document the comovement between the external and budget deficits for developing countries. Second, we explain this comovement from a model capturing the joint behavior of the external and budget deficits. It is worth stressing that these objectives are especially relevant for developing economies, since they are often perceived to face unsustainable financial positions at both country and government levels. The external deficit captures changes of foreign indebtedness, which provides information about the financial solvency of a country. The budget deficit measures changes of public indebtedness, which is related to the financial solvency of a government.

First, this paper highlights the existence of a positive comovement for all the 24 developing countries for which post-1960 data are available, as the external balance deteriorates when the budget deficit increases. Our analysis relies on the time-series of the external and budget deficits to extract the comovement that is specific to each country. Our primary measure of the comovement corresponds to the covariance between the external and budget deficits for each country. Empirically, the covariance is numerically positive for all countries and statistically significant for many cases. These results are robust to alternative measures of external deficits and exclusion of multivariate outliers of external and budget deficits.

Interestingly, our time-series findings are consistent with those obtained from panel regressions extracting the comovement that is common across countries. That is, the panel-regression results indicate that the estimated coefficient relating the external deficit to the budget deficit is always statistically positive (e.g. Calderon, Chong, and Zanforlin 2007; Gruber and Kamin 2007; Chinn and Prasad 2003; Calderon, Chong, and Loayza 2002). A similar relation is recovered from the estimated coefficients relating investment and/or saving to the budget deficit and the identity stating that the current account corresponds to national saving minus investment (Masson,Bayoumi, and Samiei 1998; Fry 1989, 1993).

Second, this paper explains the positive comovement between the external and budget deficits for 12 of the 24 initial developing countries for which all time-series are available for the estimation exercise.

To capture the joint behavior of these deficits, our analysis estimates a tractable small open-economy, overlapping-generation model with heterogeneous goods. The model offers the advantage of involving the external and internal conditions which are often considered for developing countries. The external conditions are the world interest rate, real exchange rate, and terms of trade, whereas the internal conditions are associated with domestic resources and fiscal policies, where these policies may reflect changes of the government's abilities to collect taxes due to corruption, black markets, or informal markets, for example. Also, the model captures different degrees of imperfectness of intergenerational linkages and financial markets, and as such potential liquidity constraints faced by developing countries.

The parameters of the model are estimated for each country such that the predicted second moments of the external and budget deficits, and in particular the covariance between these deficits, are close to their empirical counterparts. The predicted covariance is then decomposed into contributions measuring the portions attributable to shocks associated with each internal and external conditions. The contributions with large positive values provide information on the shocks corresponding to the prime determinants of the positive comovement between the external and budget deficits.

The results reveal that the contributions are almost always positive, so that most shocks induce a positive relation between the external and budget deficits for all countries. Also, the magnitude of the contributions indicates that the shocks associated with internal conditions, and especially domestic resources net of public absorptions, are the most important factors explaining the positive comovement between the external and budget deficits for most countries. Moreover, simulation exercises suggest that the predominance of the shocks related to internal conditions, and, in particular, government-expenditure shocks, occurs when the degree of imperfectness of intergenerational linkages or financial markets is large. Finally, a robustness analysis confirms that the contributions can be viewed as providing a lower bound of the importance of internal conditions in the determination of the positive comovement between the external and budget deficits.

Importantly, our findings contrast with those obtained from single-equation analyses characterizing either the individual behavior of the external deficit or budget deficit, rather than their joint behavior. That is, it seems difficult from the empirical literature based on single-equation techniques to identify which factors induce a positive comovement between the external and budget deficits. For example, an increase of output implies that the external deficit sometimes increases significantly (e.g. Calderon, Chong, and Zanforlin 2007; Calderon, Chong, and Loayza 2002) and sometimes does not (e.g. Chinn and Prasad 2003), whereas the budget deficit and the governments' borrowing possibilities are not significantly affected (e.g. Combes and Saadi-Sedik 2006; Roubini 1991; Berg and Sachs 1988). Also, an increase of the world interest rate induces the external deficit to significantly decreases (e.g. Calderon, Chong, and Loayza 2002), while the probability of rescheduling the public debt statistically increases such that the budget deficit may increase (e.g. Berg and Sachs 1988). Finally, an improvement of the terms of trade implies that the external deficit significantly decreases (e.g. Calderon, Chong, and Zanforlin 2007; Chinn and Prasad 2003; Calderon, Chong, and Loayza 2002), whereas the budget deficit significantly increases (e.g. Combes and Saadi-Sedik 2006).

The rest of the paper is organized as follows. Section 2 documents the empirical comovement between the external and budget deficits for several developing countries. Section 3 presents the model used to explain the joint behavior of the external and budget deficits. Section 4 elaborates the empirical method used to decompose the predicted covariance between the external and budget deficits. Section 5 reports the empirical results. Section 6 concludes.

2 Empirical Regularities

This section documents the comovement between the external and budget deficits for 24 developing countries. Our sample includes annual observations covering the longest period since the post-1960 era for 12 countries in Africa, 7 countries in the Americas, 4 countries in Asia, and 1 country in Oceania. The selections of the countries, frequency, and time periods are dictated by the availability of the data. In particular, there are several missing values for the budget deficit for many developing countries. The data are fully described in the Data Appendix.

Figure 3.1 displays the external and budget deficits. The external deficit refers to the negative of the ratio of nominal current account to nominal gross domestic product (GDP). The budget deficit corresponds to the ratio of nominal budget deficit to nominal GDP. Visual inspection of the plots suggests the existence of a positive comovement between the external and budget balances. For many countries, the external and budget positions seem to move in the same direction for most time periods. In general, these movements translate into both external and budget deficits over prolonged horizons. In some cases, however, these movements lead to external balances alternating between deficits and surpluses over time with persistent budget deficits, as for Nigeria and South Africa.

Table 3.I reports two statistics summarizing the comovements between the external and budget deficits. The first statistic is the empirical covariance between the external and budget deficits. Later on, the covariance will prove useful to perform a decomposition allowing the identification of the main factors explaining the comovement between the external and budget deficits. The second statistic is the empirical correlation between the external and budget deficits. Correlations are frequently used in business-cycle studies to document comovements between variables (e.g. Mendoza 1995). The empirical covariance and correlation are computed from the linearly detrended external and budget deficits. This transformation is applied to achieve the stationarity conditions, as suggested by computing the augmented Dickey-Fuller test on autoregressive processes including up to six lags.

The statistics indicate the existence of a positive comovement between the external and budget deficits. For example, the empirical covariance between the external and budget deficits is numerically positive for all countries. It is statistically significant at the 1% level for 11 countries, at the 10% level for 6 additional countries, and at the 25% level for 3 more countries. Likewise, the empirical correlation between the external and budget deficits is numerically positive for all countries. The correlation averages to 0.472 across all countries; it ranges from a low of 0.007 in Burundi to a high of 0.848 in Sri Lanka; and it is larger than 0.75 for 5 countries, between 0.50 and 0.75 for 6 countries, and between 0.25 and 0.50 for 7 countries. The correlation is statistically significant at the 1% level for 15 countries, at the 10% level for 4 additional countries, and at the 25% level for 2 more countries. Similar empirical covariance and correlation are obtained when the external and budget deficits are detrended from the Hodrick-Prescott filter. For briefness, these results are not reported, but are available upon request.

Tables 3.II and 3.III verify the robustness of the results by using alternative measures and methods. For the alternative measures, the external deficit is defined as (1) the ratio of the change in net external debt to GDP, and (2) the ratio of the change in net external public debt to GDP. For the alternative methods, the multivariate outliers of external deficit (measured by the negative of the current account to GDP ratio) and budget deficit are removed by using (1) the Spearman rank correlation, (2) the Kendall rank correlation, and (3) the correlation in the Mahalanobis distance associated with the 97.5% quantile of the Chi-squared distribution. For all cases, the empirical correlation is computed from the linearly detrended external and budget deficits.

For the alternative measures, a positive comovement between the external and budget deficits is confirmed for most cases. Exceptionally, the empirical correlations obtained under the first and second measures of external deficit are negative for only 2 and 3 countries, but are not significantly different from zero. For the remaining cases, the correlations for the first and second measures average to 0.417 and 0.432; and range between 0.002 and 0.758, and between 0.045 and 0.758. Also, the correlations are statistically significant at the 10% level for 19 and 18 countries. These findings suggest that the positive comovement between the external and budget deficits does not only exist under broad definitions which include the change in net external debt as well as net foreign direct investments, but also for narrow measures based exclusively on changes in net external debt.

For the alternative methods, the empirical correlations between the external and budget deficits are also positive for almost all countries. For these cases, the correlations for the first, second, and third methods average to 0.458, 0.321, and 0.485; and range between 0.044 and 0.831, from 0.006 to 0.663, and between 0.007 and 0.848. Moreover, the correlations are statistically significant at the 10% level for 17, 17, and 18 countries. These results indicate that the positive comovement between the external and budget deficits is not only observed in periods of crises (such as the well-known twin crisis affecting jointly the banking sector and the balance of payment of a country), but also during normal episodes.

Overall, the statistics reveal the existence of a positive comovement between the external and budget deficits for many developing countries. This is consistent with previous panel-regression findings, where the estimated coefficient relating the external deficit to the budget deficit is statistically positive for developing countries (e.g. Calderon, Chong, and Zanforlin 2007; Gruber and Kamin 2007; Chinn and

Prasad 2003; Calderon, Chong, and Loayza 2002). However, this contrasts with the heterogeneous results documented for industrial countries. For example, the correlation between the external and budget deficits is numerically positive for only half of a large set of industrial countries (e.g. Boileau and Normandin 2008). Also, the estimated coefficient relating the external deficit to the budget deficit is statistically positive for only half of the OECD countries (e.g. Piersanti 2000), and is no longer significant for panels of industrial countries (e.g. Chinn and Prasad 2003).

3 The Economic Environment

This section presents the economic environment explaining the joint behavior of the external and budget deficits. This environment relies on a small open-economy, overlapping-generation model with heterogeneous goods. It is worth stressing that the model is appealing for developing countries in many dimensions. In particular, it involves the external conditions which are relevant for developing economies, such as the world interest rate, real exchange rate, and terms of trade. Also, the model relates the external and budget deficits to the internal conditions associated with domestic resources and fiscal policies, where these policies may reflect, among other things, changes in the abilities of governments of developing countries to collect taxes due to corruption, black markets, or informal markets. Furthermore, the economic environment captures different degrees of imperfectness of intergenerational linkages and financial markets, and as such liquidity constraints faced by many developing countries.

Admittedly, the tractability of the model requires some simplifying assumptions. For example, it characterizes normal economic episodes (rather than extreme phenomena, such as twin crises) and defines the external deficit as the change in net external public debt (rather than using a broad definition which also includes net foreign direct investments). Note, however, that these features are consistent with the empirical results indicating that the positive comovement between the external and budget deficits is robust to the exclusion of multivariate outliers of deficits and to narrow definitions of the external deficit (see Tables 3.II and 3.III). Also, the economic environment assumes that output is exogenous, and as such the model omits the adjustments of external and budget deficits related to investment decisions. Nevertheless, the model focuses on the adjustments induced by the behavior of consumers, whom are likely to be subject to important, and perhaps the most severe, liquidity constraints.

Specifically, the model assumes that each domestic consumer born at time s solves the following problem in period t:

$$\max E_t \sum_{j=0}^{\infty} \beta^j (1-\rho)^j \frac{C_{s,t+j}^{1-\gamma}}{1-\gamma}, \tag{1.1}$$

$$s.t.\ C_{s,t} = \left[\omega^{\frac{1}{\xi}} \left(C_{s,t}^T\right)^{\frac{\xi-1}{\xi}} + (1-\omega)^{\frac{1}{\xi}} \left(C_{s,t}^N\right)^{\frac{\xi-1}{\xi}}\right]^{\frac{\xi}{\xi-1}}, \tag{1.2}$$

$$C_{s,t}^T = \left[\varpi^{\frac{1}{\zeta}} \left(C_{s,t}^H\right)^{\frac{\zeta-1}{\zeta}} + (1-\varpi)^{\frac{1}{\zeta}} \left(C_{s,t}^F\right)^{\frac{\zeta-1}{\zeta}}\right]^{\frac{\zeta}{\zeta-1}}, \tag{1.3}$$

$$(1-\rho)\left(B_{s,t+1} + F_{s,t+1}\right) = (1+r_t)\left(B_{s,t} + F_{s,t}\right) + Y_{s,t} - T_{s,t} - P_t C_{s,t}. \tag{1.4}$$

Equation (1.1) specifies the utility function in terms of private consumption of a composite good. Equation (1.2) defines this consumption in terms of tradable and non-tradable goods. Equation (1.3) expresses tradable consumption in terms of home and foreign tradable goods. Equation (1.4) depicts the intertemporal budget constraint of the consumer.

All variables are measured in terms of home tradable goods. Specifically, $C_{s,t}$ denotes an index of private consumption, $C_{s,t}^N$ is the consumption of non-tradable goods, $C_{s,t}^T$ is the consumption of tradable goods, $C_{s,t}^F$ is the consumption of foreign tradable goods, and $C_{s,t}^H$ is the consumption of home tradable goods. P_t, P_t^N, P_t^T, P_t^F, and P_t^H are the corresponding price indices, with the normalization $P_t^H = 1$. $B_{s,t}$ is the purchase of one-period bonds issued by the domestic government, $F_{s,t}$ is the purchase of one-period bonds issued by the foreign government, r_t is the world interest rate on one-period bonds, $T_{s,t}$ is lump-sum taxes, and $Y_{s,t} = Y_{s,t}^H + P_t^N Y_{s,t}^N$ is the value of output, where $Y_{s,t}^H$ and $Y_{s,t}^N$ are resources of home tradable and non-tradable goods. The term E_t represents the expectation operator conditional on the information available in period t.

Also, the parameter β corresponds to the discount factor, γ is the reciprocal of the elasticity of intertemporal substitution of consumption, ξ is the elasticity of substitution between tradable and non-tradable goods, ζ is the elasticity of substitution between home and foreign tradable goods, ω is the weight of tradable goods in total consumption, and ϖ is the weight of home tradable goods in total tradable consumption. The parameter ρ is the probability of being dead next period, or equivalently, the death and birth rates when the population is constant (e.g. Blanchard 1985). Consequently, $\rho = 0$ indicates that the domestic economy is described by an infinitely-lived representative consumer model, so that agents fully smooth their consumption. Conversely, $\rho = 1$ implies that the domestic environment is represented by a sequence of static economies in which each cohort is fully replaced in the subsequent period by a different cohort, such that agents consume only their current income. The parameter ρ may be related to the imperfectness of intergenerational linkages. In this context, a large value of ρ indicates that consumers are not altruistic, so that agents prefer a consumption profile which is not fully smoothed. Alternatively, ρ may be related to the degree of imperfectness of financial markets. In this case, a large value of ρ indicates that consumers experience difficulties in selling or buying bonds, so that agents are unable to fully smooth consumption through time.

The domestic public sector is described as:

$$\left(B_{t+1} + B^*_{t+1}\right) = (1 + r_t)\left(B_t + B^*_t\right) + P_t G_t - T_t, \tag{2.1}$$

$$= \left(B_t + B^*_t\right) + D_t. \tag{2.2}$$

Equations (2.1) and (2.2) correspond to the intertemporal budget constraint of the government. The variables without the subscript s refer to aggregate variables. In particular, B^*_t is the aggregate foreign purchases of one-period domestic bonds, G_t is the government expenditures on goods, and D_t is the budget deficit including the debt service.

The external deficit of the domestic economy is measured as the negative of the current account. The current account is:

$$Z_t = (F_{t+1} - F_t) - \left(B^*_{t+1} - B^*_t\right), \tag{3}$$

where equation (3) defines the current account as the change ofnet foreign asset positions.

The model (1.1)–(3) is solved from an analytical approximation, which is fully described in a technical appendix available from the authors. In brief, the individual consumption function is derived, first, from the Euler equation associated with (1.1) and (1.4) and the distributional assumption of log normality (e.g. Campbell and Mankiw 1989), and, second, from the expected integrated budget constraint associated with (1.4) which is linearized around the means (e.g. Campbell and Deaton 1989). Then, the aggregate consumption function is derived from the individual consumption function and the assumptions that all consumers alive in a given time period face identical taxes and have the same tradable and non-tradable outputs (e.g. Gali 1990). The current account function is derived from the definition (3) , the aggregate budget constraints associated with (1.4) and (2.1) , and the aggregate consumption function. To highlight the relation between the external and budget deficits, the current account function is rewritten by substituting aggregate taxes from the expected integrated budget constraint associated with (2.1) and (2.2) , which is linearized around the means (e.g. Normandin 1999). Finally, the consumer price indices associated with (1.2) and (1.3) are log-linearized around the means of exchange rate and terms of trade. The exchange rate is defined as $q_t = \left(\frac{P_t^N}{P_t^T}\right)$. The terms of trade correspond to $\tau_t = \left(\frac{P_t^H}{P_t^F}\right)$.

The analytical approximation takes the following form:

$$\mathbf{x}_{1,t} = \mathbf{\Theta}_{11}\mathbf{x}_{1,t-1} + \mathbf{\Theta}_{12}\mathbf{x}_{2,t}, \tag{4}$$

or more explicitly,

$$\begin{pmatrix} \mathbf{P}_{t+1} \\ z_t \end{pmatrix} = \begin{pmatrix} \mathbf{\Theta}_{pp} & \mathbf{0} \\ \mathbf{\Theta}_{zp} & 0 \end{pmatrix} \begin{pmatrix} \mathbf{P}_t \\ z_{t-1} \end{pmatrix} + \begin{pmatrix} \mathbf{\Theta}_{pf} & \mathbf{\Theta}_{pa} \\ \mathbf{\Theta}_{zf} & 1 \end{pmatrix} \begin{pmatrix} \mathbf{f}_t \\ a_t \end{pmatrix},$$

with

$$a_t = \mathbf{\Theta}_{af} E_t \sum_{j=1}^{\infty} \lambda^j \mathbf{f}_{t+j}, \tag{5}$$

where expression (4) corresponds to the rules for the predetermined and nonpredetermined variables. Equation (5) represents the purely forward-looking component of the rules.

All variables are demeaned. The predetermined variables are $\mathbf{P}_t = \left((f_t - b_t^*) \quad (b_t + b_t^*)\right)'$, where $(f_t - b_t^*) = \frac{(F_t - B_t^*)}{Y_{t-1}}$ and $(b_t + b_t^*) = \frac{(B_t + B_t^*)}{Y_{t-1}}$. The nonpredetermined variable is $z_t = \frac{Z_t}{Y_t}$. The forcing variables are $\mathbf{f}_t = (r_t \ \Delta\log\tau_t \ \Delta\log q_t \ \Delta\log Y_t \ \log g_t \ d_t)'$, where Δ is the first difference operator, $g_t = \frac{P_t G_t}{Y_t}$, and $d_t = \frac{D_t}{Y_t}$. The forcing variables include the typical exogenous stochastic variables for small open economies, such as developing countries. These variables reveal information on the external conditions related to interest rate (r_t) , terms of trade $(\Delta\log\tau_t)$, and exchange rate $(\Delta\log q_t)$, as well as on the internal conditions related to domestic resources $(\Delta\log Y_t)$, net of public absorptions $(\log g_t)$, and fiscal policies (d_t). Specifically, the budget deficit provides information on taxes, since government expenditures and debt service are given (that is, g_t and r_t are exogenous, and $(b_t + b_t^*)$ is predetermined). For convenience, a_t is termed the adjusted current account.

Table 3.IV relates the coefficients of the rules to the parameters of the model and the means of the variables. These coefficients reveal that the rules are static when the probability of death is unity ($\rho = 1$, so that $\lambda = 0$). In this case, the current account is exclusively affected by the contemporaneous output and budget deficit (see the nonzero elements of $\mathbf{\Theta}_{zf}$). First, the current account improves following an increase of output, through a positive wealth effect. Second, the current account deteriorates following an increase of budget deficit, since it reflects a tax cut which leads to an increase of consumption (including that of foreign tradable goods). This translates into a positive relation between the external and budget deficits. As explained above, this relation can be due to a non-altruistic behavior associated with imperfect intergenerational linkages or to liquidity constraints related to imperfect financial markets.

In contrast, the rules are dynamic when the probability of death is smaller than one ($0 \leq \rho < 1$, so that $0 < \lambda < 1$). In this case, the current account is affected by all expected future forcing variables (see the elements of $\mathbf{\Theta}_{af}$). First, the current account deteriorates in response to an expected increase of output and an expected decrease of government expenditures, since this expected increase of resources, net of public absorption, induces an increase of current consumption (including that of foreign tradable goods). Second, when the elasticity of intertemporal substitution of consumption exceeds one, then the current account may deteriorate in response to an expected decrease of interest rate, an expected appreciation of exchange rate, and an expected deterioration of terms of trade, through the substitution effects associated with an increase of the price of future consumption relative to that of current consumption, an increase of the price of future non-tradable goods relative to that of future tradable goods, and an increase of the price of future foreign tradable goods relative to that of future home tradable goods. Finally, a positive probability of death implies that the current account deteriorates in response to an expected increase of the budget deficit, whereas a zero probability of death implies that the current account is unaffected because the contemporaneous consumption is unaltered while private saving increases to reimburse the budget deficit induced by a tax cut. Hence, a zero probability of death implies that there is no relation between the external and budget deficits.

Note that the forward-looking component (5) involves the expectations of future forcing variables. These expectations are constructed from the process:

$$\mathbf{w}_t = \mathbf{\Phi}\mathbf{w}_{t-1} + \mathbf{u}_t. \tag{6}$$

Here, $\mathbf{w}_t = \begin{pmatrix} \mathbf{f}_t' & a_t \end{pmatrix}'$ includes all actual forcing variables and the actual adjusted current account, $\mathbf{u}_t$ contains the innovations, $\mathbf{\Phi}$ incorporates the feedback coefficients, and $\mathbf{\Omega}_u = E\left[\mathbf{u}_t\mathbf{u}_t'\right]$ captures the covariances of the innovations. As usual, the forcing variables summarize the information contained in the history of these variables that helps forecast future forcing variables. Also, it can be shown that the adjusted current account captures the additional information contained in hidden variables that improves the forecasts of future forcing variables (e.g. Boileau and Normandin 2002).

The analytical approximation is completed by constructing the expectations of future forcing variables from the process (6) and by substituting these terms in the forward-looking component (5). This yields:

$$\mathbf{x}_{2,t} = \mathbf{\Theta}_{22}\mathbf{x}_{2,t-1} + \mathbf{\Theta}_{2u}\mathbf{u}_t. \tag{7}$$

Here, $\mathbf{x}_{2,t} = \mathbf{\Theta}_{2u}\mathbf{w}_t$ includes all actual forcing variables and the predicted adjusted current account, $\mathbf{\Theta}_{2u}\mathbf{u}_t$ contains the innovations, and $\mathbf{\Theta}_{22} = \mathbf{\Theta}_{2u}\mathbf{\Phi}\mathbf{\Theta}_{2u}^{-1}$ incorporates the feedback coefficients. Also, $\mathbf{\Theta}_{2u} = \begin{pmatrix} \mathbf{e}_1 \; ... \; \mathbf{e}_6 \; \Upsilon' \end{pmatrix}'$, $\mathbf{e}_k$ contains the value one for the kth element and zero elsewhere, $\Upsilon = \mathbf{\Theta}_{af}\mathbf{e}\mathbf{\Phi}\lambda\left[\mathbf{I} - \mathbf{\Phi}\lambda\right]^{-1}$, $\mathbf{e} = (\mathbf{e}_1 \; ... \; \mathbf{e}_6)'$, and $\mathbf{I}$ is the identitymatrix. Note that some of the feedback coefficients of (7) reflect the dynamic interactions between the budget deficit and the forcing variables related to internal and external conditions. These responses of the budget deficit and the responses of the current account (discussed above) may lead to a positive relation between the external and budget deficits.

Finally, the processes (4) and (7) are stacked to form the first-order restricted vector autoregression (VAR):

$$\mathbf{x}_t = \mathbf{\Theta}_x \mathbf{x}_{t-1} + \mathbf{\Theta}_u \mathbf{u}_t, \tag{8}$$

or

$$\begin{pmatrix} \mathbf{x}_{1,t} \\ \mathbf{x}_{2,t} \end{pmatrix} = \begin{pmatrix} \mathbf{\Theta}_{11} & \mathbf{\Theta}_{12}\mathbf{\Theta}_{22} \\ \mathbf{0} & \mathbf{\Theta}_{22} \end{pmatrix} \begin{pmatrix} \mathbf{x}_{1,t-1} \\ \mathbf{x}_{2,t-1} \end{pmatrix} + \begin{pmatrix} \mathbf{\Theta}_{12}\mathbf{\Theta}_{2u} \\ \mathbf{\Theta}_{2u} \end{pmatrix} \mathbf{u}_t.$$

The representation (8) imposes restrictions that reflect the structure of the model. This representation will prove useful to extract the predictions of the model, including those stating which are the key factors that induce a positive relation between the external and budget deficits.

Throughout our analysis, the predictions of the model will also be confronted to their empirical counterparts. For this purpose, these counterparts will be measured from the first-order unrestricted VAR:

$$\mathbf{s}_t = \mathbf{\Theta}_s \mathbf{s}_{t-1} + \mathbf{v}_t. \tag{9}$$

Here, $\mathbf{s}_t = \left(\mathbf{f}_t' \quad z_t\right)'$ includes all actual forcing variables as well as the actual current account, $\mathbf{v}_t$ are the innovations, $\mathbf{\Theta}_s$ are the feedback coefficients, and $\mathbf{\Omega}_v = E\left[\mathbf{v}_t \mathbf{v}_t'\right]$ are the covariances of the innovations. Note that the specification (9) does not place any restrictions, unlike the representation (8).

4 Estimation Method

This section elaborates the estimation method designed to evaluate the parameters of the restricted VAR (8) and to identify the main determinants of the positive comovement between the external and budget deficits. Ideally, the method should jointly estimate all the parameters of system (8). In practice, however, this exercise is difficult to perform given the large number of parameters to estimate relative to the number of observations. Specifically, there is a total of 92 parameters which include (i) the means μ_y, μ_q, μ_τ, μ_r, μ_g, μ_d, $\mu_{(f-b^*)}$, μ_c, and μ_t associated with the variables $\Delta \log Y_t$, $\log q_t$, $\log \tau_t$, r_t, $\log g_t$, d_t, $(f_t - b_t^*)$, $\log c_t = \log\left(\frac{P_t C_t}{Y_t}\right)$, and $\log t_t = \log\left(\frac{T_t}{Y_t}\right)$, (ii) the feedback coefficients and innovation covariances incorporated in $\mathbf{\Phi}$ and $\mathbf{\Omega}_u$ of the process (6), and (iii) the preference parameters ρ, γ, ξ, ζ, ω, and ϖ involved in the agent's problem (1.1) – (1.4). In contrast, the samples for our different countries include between 24 and 44 annual observations per variable.

To circumvent this problem, we apply the following multi-step estimation procedure. The first step uses the sample estimates to evaluate the means. The second step estimates the feedback coefficients and innovation covariances by Ordinary Least Squares (OLS). For this purpose, the variables involved in the process (6) are obtained by using the actual data for forcing variables, current account, and predetermined variables; by constructing the adjusted current account as $a_t = z_t - \mathbf{\Theta}_{zp}\mathbf{p}_t - \mathbf{\Theta}_{zf}\mathbf{f}_t$ (see the last equation of system (4)); and by fixing the means to their estimates and the preference parameter ρ to a given value (see Table 3.IV).

The third step estimates the preference parameters ρ, γ, ξ and ζ by the Generalized Method of Moments (GMM). To do so, the means, feedback coefficients, and innovation covariances are fixed to their estimates and the following second moment conditions are exploited:

$$E\left[\left(cov + z_t d_t\right)\mathbf{w}_{t-1}\right] = 0, \tag{10.1}$$

$$E\left[\left(var_z - z_t^2\right)\mathbf{w}_{t-1}\right] = 0, \tag{10.2}$$

$$E\left[\left(var_d - d_t^2\right)\mathbf{w}_{t-1}\right] = 0. \tag{10.3}$$

Here, cov is set to the empirical covariance (reported in Table 3.I), while var_z and var_d are fixed to the empirical variances of the current account and budget deficit. Also, $\mathbf{w}_{t-1}$ is the vector of instruments, while z_t and d_t are substituted by the values predicted by the restricted VAR (8).

$$z_t = \mathbf{e}_3^{'}\mathbf{x}_t = \mathbf{e}_3^{'}\left(\mathbf{\Theta}_x\mathbf{x}_{t-1} + \mathbf{\Theta}_u\mathbf{u}_t\right), \tag{11.1}$$

$$d_t = \mathbf{e}_9^{'}\mathbf{x}_t = \mathbf{e}_9^{'}\left(\mathbf{\Theta}_x\mathbf{x}_{t-1} + \mathbf{\Theta}_u\mathbf{u}_t\right). \tag{11.2}$$

To gain intuition, note that equation (10.1) implies that the GMM estimates select values for the preference parameters involved in the nonlinear regression $cov = -z_t d_t + \epsilon_t$ such that the error term ϵ_t is orthogonal to lagged information, where z_t and d_t are given by $(11.1) - (11.2)$. Taking expectations implies that $E\left(cov\right) = -E\left(z_t d_t\right) + E\left(\epsilon_t\right)$ or $cov = -\sigma_{zd}$ provided that the error term is centered on zero, so that the estimated preference parameters ensure that the predicted covariance between the external and budget deficit, $-\sigma_{zd}$, is close to the empirical covariance, cov, shown in Table 3.I. Analogously, expressions (10.2) and (10.3) yield GMM estimates implying that the predicted variances of the current account and budget deficit are close to their empirical counterparts. Moreover, the predicted covariance and variances imply, by construction, that the predicted correlation between the external and budget deficits is similar to the empirical one (reported in Table 3.I). Hence, the three conditions $(10.1) - (10.3)$ allow one to predict the moments that are used to document the empirical regularities describing the comovement between the external and budget deficits. In practice, the GMM estimates are obtained for the preference parameters ρ and γ and for the composite parameters $\epsilon_q = 1 + \left(\frac{1-\omega}{\omega}\right)e^{(1-\xi)\mu_q}$ and $\epsilon_\tau = 1 + \left(\frac{1-\varpi}{\varpi}\right)e^{(1-\zeta)\mu_\tau}$, given that the preference parameters ω, ϖ, ξ, and ζ are not individually identified.

The last step of the estimation procedure consists in repeating the second and third steps. More explicitly, the estimate of ρ obtained in the third step is used in the second step to reestimate the feedback coefficients $\mathbf{\Phi}$ and innovation covariances $\mathbf{\Omega}_u$. These new estimates of $\mathbf{\Phi}$ and $\mathbf{\Omega}_u$ are then used in the third step to update the estimate of ρ. These iterations are done until the estimates of the preference parameter ρ and those in $\mathbf{\Phi}$ and $\mathbf{\Omega}_u$ converge to fix points.

The empirical method then uses the estimates of parameters to evaluate and decompose the predicted covariance between the external and budget deficits. This exercise is relevant since the estimation procedure ensures that the predicted covariance is close to the empirical one. Specifically, the predicted covariances between the variables governed by the restricted VAR (8) are given by:

$$\mathbf{\Sigma}_x = \sum_{j=0}^{\infty} \mathbf{\Theta}_x^j \mathbf{\Theta}_u \mathbf{\Omega}_u \mathbf{\Theta}_u^{'} \mathbf{\Theta}_x^{j'}, \tag{12.1}$$

$$= \sum_{j=0}^{\infty} \mathbf{\Psi}_{x,j} \mathbf{\Psi}_{x,j}^{'}. \tag{12.2}$$

Here, $\mathbf{\Sigma}_x = E\left[\mathbf{x}_t \mathbf{x}_t^{'}\right]$ is the predicted covariance matrix, $\mathbf{\Omega}_u = \mathbf{\Lambda}_u \mathbf{\Lambda}_u^{'}$ is the covariance matrix of the innovations of the process (6), $\mathbf{\Psi}_{x,j} = \mathbf{\Theta}_x^j \mathbf{\Theta}_u \mathbf{\Lambda}_u$ is the matrix summarizing the predicted dynamic responses of various variables j periods after each orthogonal shock, and $\mathbf{\Lambda}_u$ is a lower triangular matrix transforming the innovations into orthogonal shocks. In practice, $\mathbf{\Lambda}_u$ is obtained from a Choleski factorization associated with a given ordering of the variables contained in $\mathbf{w}_t = \begin{pmatrix} \mathbf{f}_t^{'} & a_t \end{pmatrix}^{'}$. As will be clear later on, different orderings are used to verify the robustness of the results.

From expression (12.2), the predicted covariance between the external and budget deficits is decomposed as follows:

$$-\sigma_{zd} = \psi_{x,r} + \psi_{x,\tau} + \psi_{x,q} + \psi_{x,y} + \psi_{x,g} + \psi_{x,d} + \psi_{x,a}. \tag{13}$$

The component $\psi_{x,r} = \sum_{j=0}^{\infty} \left(-\mathbf{e}_3^{'} \mathbf{\Psi}_{x,j} \mathbf{e}_4\right) \left(-\mathbf{e}_9^{'} \mathbf{\Psi}_{x,j} \mathbf{e}_4\right)$ represents the portion of the predicted covariance between the external and budget deficits which is attributable to the interest-rate shock. In practice, this portion is computed by evaluating the sum over 100 years. Also, the portion involves the predicted dynamic responses of external deficit $\left[i.e.\ \left(-\mathbf{e}_3^{'} \mathbf{\Psi}_{x,j} \mathbf{e}_4\right)\right]$ and budget deficit $\left[i.e.\ \left(-\mathbf{e}_9^{'} \mathbf{\Psi}_{x,j} \mathbf{e}_4\right)\right]$ to an interest-rate shock. The other terms in (13) are defined in an analogous way. The components $\psi_{x,\tau}$, $\psi_{x,q}$, $\psi_{x,y}$, $\psi_{x,g}$, $\psi_{x,d}$ and $\psi_{x,a}$ measure the contributions to the predicted covariance of the terms-of-trade shock, exchange-rate shock, output shock, government-expenditure shock, tax shock, and other shocks. The predicted contribution with the largest positive value provides information on the shock corresponding to the prime determinant of the positive comovement between the external and budget deficits.

Finally, the predicted contributions of each shock are compared to the empirical ones. To do so, the empirical contributions are computed by using the OLS estimates of the unrestricted VAR (9) to evaluate the following expression:

$$\boldsymbol{\Sigma}_s = \sum_{j=0}^{\infty} \boldsymbol{\Theta}_s^j \boldsymbol{\Omega}_v \boldsymbol{\Theta}_s^{j'} \tag{14.1}$$

$$= \sum_{j=0}^{\infty} \boldsymbol{\Psi}_{s,j} \boldsymbol{\Psi}_{s,j}^{'}. \tag{14.2}$$

Note that $\boldsymbol{\Sigma}_s = E\left[\mathbf{s}_t \mathbf{s}_t^{'}\right]$ is the empirical covariance matrix and $\boldsymbol{\Omega}_v = \boldsymbol{\Lambda}_v \boldsymbol{\Lambda}_v^{'}$ is the covariance matrix of the innovations associated with the process (9). Also, $\boldsymbol{\Psi}_{s,j} = \boldsymbol{\Theta}_s^j \boldsymbol{\Lambda}_v$ contains the empirical dynamic responses and $\boldsymbol{\Lambda}_v$ is a lower triangular matrix associated with a Choleski factorization for a given ordering of the variables included in $\mathbf{s}_t = \begin{pmatrix} \mathbf{f}_t^{'} & z_t \end{pmatrix}^{'}$. Again, different orderings will be used to gauge the robustness of the results.

5 Results

This section applies the empirical method just described for 12 of our 24 initial countries. The selection of the countries relies on the availability of the data required for the estimation exercise. In particular, the series on public debt are often missing. The data are fully described in the Data Appendix.

The estimates of the means, feedback coefficients, and innovation covariances used to construct the restricted VAR (8) are available upon request. Empirically, the estimates of the feedback coefficients and innovation covariances are obtained from a version of the first-order process (6), which includes constant terms and a linear trend. Note that higher lag structures sometimes lead to perfect collinearity and are never selected by the Bayesian information criterion.

Table 3.V reports the estimates of the preference and composite parameters, which are also used to construct the restricted VAR(8). The estimates systematically display the expected signs and the appropriate magnitudes. Specifically, the estimates indicate that the probability of death, ρ, is always between zero and one. Also, the probability of death averages to 0.517 across all countries; and it ranges from a low of 0.083 in Tunisia to a high of 0.863 in Sierra Leone. As explained above, large values of ρ may reflect non-altruistic behavior associated with imperfect intergenerational linkages

or liquidity constraints related to imperfect financial markets. Interestingly, previous findings detect strong liquidity constraints for many developing countries (e.g. Haque and Montiel 1989). Among the countries which are common to our sample, severe liquidity constraints are documented for India and Nigeria, but not for Morocco. Our estimates accord with these results, that is, the estimates of the probability of death are larger for India and Nigeria than for Morocco.

The estimates imply that the elasticity of intertemporal substitution of consumption, $\frac{1}{\gamma}$, is always between zero and one. Also, the elasticity averages to 0.594 across all countries; and it ranges from a low of 0.385 in Mauritius to a high of 0.932 in Pakistan. These estimates are consistent with previous findings, where the elasticity is smaller than unity for almost all selected developing countries (e.g. Ogaki, Ostry, and Reinhart 1996; Ostry and Reinhart 1992; Giovannini 1985).

As expected, the estimates reveal that the composite parameters, ϵ_q and ϵ_τ, are always larger than one. Also, the elasticities of substitution between tradable and non-tradable goods, ξ, and between home and foreign tradable goods, ζ, are always positive — where these elasticities are recovered from the definitions $\epsilon_q = 1 + \left(\frac{1-\omega}{\omega}\right) e^{(1-\xi)\mu_q}$ and $\epsilon_\tau = 1 + \left(\frac{1-\varpi}{\varpi}\right) e^{(1-\zeta)\mu_\tau}$ and, given values of the weights ω and ϖ, and the estimated values of the means μ_q and μ_τ. The elasticities ξ and ζ uniformly decline as the weights of tradable goods in total consumption, ω, and of home tradable goods in total tradable consumption, ϖ, increase. For example, ξ and ζ average to 1.425 and 1.588 when $\omega = \varpi = 0.3$; whereas ξ and ζ average to 1.067 and 1.210 when $\omega = \varpi = 0.7$. Overall, these values are consistent with previous findings, where the estimates of the elastiticity of substitution between imported and non-tradable goods are about 1.2 and the estimates of the weight are around 0.8 across selected developing countries (e.g. Ostry and Reinhart 1992). In these studies, however, the tradable goods are usually not decomposed into imported and non-imported goods.

Table 3.VI evaluates the fit of the model. In particular, the overidentification restrictions associated with the three moment conditions $(10.1) - (10.3)$ are tested. Empirically, these restrictions are refuted at the 5% level for only 4 countries. The analysis is pursued by performing the challenging exercise testing that the empirical and predicted covariances and correlations between the external and budget deficits are the same. The empirical moments are computed directly from the raw data (as in Table

3.I) or from the first-order unrestricted VAR (9), which includes constant terms and a linear trend. The predicted moments are calculated from the restricted VAR (8). Importantly, the empirical and predicted moments display similar values for most countries. Also, the restrictions of identical empirical and predicted moments are rejected at the 5% level for 5 and 6 countries, when the empirical covariances and correlations are generated from the raw data and from the unrestricted VAR. Although not reported for briefness, the empirical and predicted standard deviations of the external deficit are not statistically different for almost all countries, whether the empirical standard deviations are computed from the raw data or the unrestricted VAR. Similar results hold when the empirical standard deviations of the budget deficit are confronted to the predicted ones.

Turning to the central objective of our analysis, Table 3.VII reports the decompositions of the empirical and predicted covariances between the external and budget deficits. The empirical and predicted contributions of each shock to the covariances are computed from expressions (14.1) and (12.1), by using the unrestricted VAR (9) and restricted VAR (8) with Choleski factorizations associated with the following benchmark orderings: $\mathbf{s}_t = \begin{pmatrix} \mathbf{f}_t^{'} & z_t \end{pmatrix}^{'}$ and $\mathbf{w}_t = \begin{pmatrix} \mathbf{f}_t^{'} & a_t \end{pmatrix}^{'}$, where $\mathbf{f}_t = (r_t \; \Delta\log\tau_t \; \Delta\log q_t \; \Delta\log Y_t \; \log g_t \; d_t \,)^{'}$. These orderings place the variables related to the external conditions before those associated with the internal conditions. This implies that the external conditions are more predetermined than the internal ones, as expected for small open economies such as developing countries. As such, the shocks related to the internal conditions capture the portions that are orthogonal to the external variables. Also, the shocks related to $\log g_t$ correspond to shocks on the level of government expenditures, rather than on the ratio of government expenditures to output, given that they capture the portions that are orthogonal to output. Likewise, the shocks related to d_t are shocks on taxes, since they measure the portions that are orthogonal to output, government expenditures, and interest rates. Finally, the shocks affecting z_t and a_t capture any other shocks than those already associated with $\mathbf{f}_t$, because they are the portions that are orthogonal to forcing variables.

The estimates of the empirical contributions are most of the time positive (see Table 3.VII). This implies that shocks often induce a positive relation between the external and budget deficits, translating into a positive empirical covariance between the two deficits. Also, the empirical contributions display different sizes across the various shocks. For the internal conditions, the empirical contributions related to the government-expenditure shock exhibit the largest numerical values for 7 out of 12 countries (i.e. Costa Rica, Honduras, Mauritius, Morocco, Pakistan, Sierra Leone, and Sri Lanka), the portions attributable to the output shock are the most important components for 2 countries (i.e. Nigeria and South Africa), whereas the contributions of the budget-deficit shock associated with a tax change never display the largest values. For the external conditions, the contributions related to the interest-rate shock reach the largest numerical values for only 1 country (i.e. India), and the portions attributable to the exchange-rate shock are the most important components for only 2 countries (i.e. Malaysia and Tunisia), while the contributions of the terms-of-trade shock never exhibit the largest values.

Importantly, the estimates of the predicted contributions lead to similar findings (see Table 3.VII). Specifically, the predicted contributions are almost always positive, so that most shocks generate a positive relation between the external and budget deficits. Also, the sizes of the predicted contributions indicate that, for the internal conditions, the government-expenditure and output shocks represent the most important factors for 6 countries (i.e. Costa Rica, Mauritius, Morocco, Sierra Leone, Sri Lanka, and Tunisia) and 3 countries (i.e. Honduras, Nigeria, and South Africa), whereas, for the external conditions, the interest-rate and terms-of-trade shocks correspond to the dominant factors for only 1 country (i.e. India) and 2 countries (i.e. Malaysia and Pakistan).

In sum, the empirical and predicted contributions reveal that the shocks associated with the internal conditions, and especially the domestic resources net of public absorptions, are the most important factors explaining the positive comovement between the external and budget deficits for 9 out of 12 countries, whereas the external conditions represent the prime factors for only 3 countries. Similar results are obtained by selecting the empirical and predicted contributions with the largest positive, statistically significant (rather than numerical), values. Overall, these results contrast with the findings obtained from single-equation analyses, where it seems difficult to explain the positive relation between the external and budget deficits (e.g. Calderon, Chong, and Zanforlin 2007; Combes and Saadi-Sedik 2006; Chinn and Prasad 2003; Calderon, Chong, and Loayza 2002; Roubini 1991; Berg and Sachs 1988).

At this point, it is instructive to provide some intuitions behind the covariance-decomposition results presented so far, and, in particular, to understand why the positive relation between the external and budget deficits is predominantly explained by government-expenditure shocks. For this purpose, Figure 3.2 displays, for Sri Lanka, the empirical and predicted dynamic responses of external and budget deficits following positive, one standard deviation, shocks. The case of Sri Lanka is selected because it constitutes an example of the countries where government-expenditure shocks correspond to the most important explanatory factor. Also, the empirical and predicted responses are computed from the unrestricted VAR (9) and restricted VAR (8) with the benchmark orderings.

For each shock, the empirical and predicted responses of external deficit are similar, at all horizons. Likewise, the empirical and predicted responses of budget deficit are almost identical. Given the covariance decompositions (14.1) and (12.1), these responses imply that the empirical and predicted contributions of each shock are fairly close. Focusing now on the restricted VAR (8), the various shocks produce predicted responses of external and budget deficits which are almost always of the same sign, for each horizon. This implies that most shocks induce a positive relation between the external and budget deficits. Moreover, the various shocks induce predicted responses of external and budget deficits which display different magnitudes. Specifically, a government-expenditure shock leads to the largest (in absolute values) response of external deficit and to a pronounced response of budget deficit. This result implies that the predicted contribution of the government-expenditure shocks is the largest.

The analysis is next deepen by performing simulation exercises to illustrate the extent to which the signs and sizes of the contributions, and especially those of government-expenditure shocks, are affected under different values of some key parameters of the model. To this end, Table 3.VIII decomposes the simulated covariances between the external and budget deficits. The simulated contributions are computed by evaluating the restricted VAR (8) with the benchmark ordering from various sets of parametrizations. Each set alters the value of only one of the preference parameters ρ, γ, ξ, and ζ at a time, but fixes the weights ω and ϖ to 0.7 and sets all the other parameters to the estimated values for Sri Lanka.

The simulations reveal that the contributions of the shocks are systematically positive, whether the parametrizations involve small or large values of the preference parameters. Also, the magnitudes of the contributions are fairly similar across the various parametrizations. More importantly, the contributions of government-expenditure shocks are almost always the largest, confirming that this type of shocks constitutes the prime explanatory factor of the positive relation between the external and budget deficits. Exceptionally, the contributions of government-expenditure shocks become smaller than the contributions of interest-rate shocks, and to a lesser extent to those of terms-of-trade shocks, when the probability of death, ρ, is small.

Figure 3.3 shows that this exceptional case occurs because a positive interest-rate shock leads to a larger (in absolute values) response of external deficit over time when ρ is small. Recall that this type of shock induces a decrease of the price of future consumption relative to that of current consumption, which, in turn, implies an immediate increase of private saving and a decrease of external deficit (see Section 3). Moreover, the increase of private saving is magnified as the probability of death decreases, given that small values of ρ may reflect more altruistic behavior or less severe liquidity constraints.

Likewise, Figure 3.3 indicates that a terms-of-trade shock yields a larger (in absolute values) response of external deficit over several horizons when ρ is small. In contrast, a government-expenditure shock mainly generates a larger response of external deficit at impact, but not over longer horizons, when ρ is small. However, equation (12.1) implies that selecting small values of ρ has a negligible effect on the contributions of government-expenditure shocks because this component involves the product of the response of external deficit and the response of budget deficit, while the latter is null at impact.

Finally, Table 3.IX verifies the robustness of the covariance-decomposition results. To do so, the empirical contributions(14.1) and predicted contributions (12.1) are computed from the unrestricted VAR (9) and restricted VAR (8) with the alternative orderings $\mathbf{s}_t = \left(\mathbf{f}_t' \;\; z_t\right)'$ and $\mathbf{w}_t = \left(\; \mathbf{f}_t' \quad a_t \;\right)'$, where $\mathbf{f}_t = \left(\Delta \log Y_t \;\; \log g_t \;\; d_t \;\; r_t \;\; \Delta \log \tau_t \;\; \Delta \log q_t\right)'$. These orderings imply that the internal conditions are more predetermined than the external ones, unlike the benchmark orderings. Importantly, this case may be relevant for some developing economies which heavily rely on natural resources. This is because the endowment of these resources is crucially affected by exogenous, and possibly highly predetermined, factors. Examples of such factors are weather conditions affecting crops for agricultural economies (e.g. Costa Rica), geological conditions affecting the mining industry (e.g. South Africa), and reserves of oil affecting the petrolium industry (e.g. Nigeria). All these exogenous factors are captured by shocks to domestic resources, net of public absorption, in our alternative orderings.

For the alternative orderings, the magnitudes of the empirical contributions indicate that, for the internal conditions, the government-expenditure, output, and budget-deficit shocks are the dominant factors for 4 countries (i.e. Costa Rica, Morocco, Pakistan, and Sierra Leone), 5 countries (i.e. Honduras, Malaysia, Mauritius, Nigeria, and South Africa), and 2 countries (i.e. India and Sri Lanka), while, for the external conditions, the exchange-rate shock is the prime factor for only 1 country (i.e. Tunisia). Similarly, the predicted contributions imply that the government-expenditure, output, and budget-deficit shocks are the dominant factors for 9 countries (i.e. Costa Rica, Honduras, Mauritius, Morocco, Nigeria, Pakistan, Sierra Leone, Sri Lanka, and Tunisia), 2 countries (i.e. Malaysia and South Africa), and 1 country (i.e. India). Consequently, the empirical and predicted contributions confirm that the internal conditions, and in particular the domestic resources net of public absorptions, are the most important factors explaining the positive relation between the external and budget deficits for all countries, except perhaps for Tunisia. Interestingly, these contributions reinforce the conclusion reached from the benchmark orderings. Moreover, the results suggest that the contributions obtained from the benchmark orderings can be viewed as providing a lower bound of the importance of the internal conditions in the determination of the positive comovement between the external and budget deficits.

6 Conclusion

This paper documented and explained the positive relation between the external and budget deficits of developing countries for which post-1960 data are available. First, we highlighted the existence of a positive comovement between the external and budget deficits. Our analysis relies on time-series of the external and budget deficits to extract the comovement that is specific to each country. Empirically, the empirical covariance between the external and budget deficits is always positive and is statistically significant for many cases. This is consistent with previous findings obtained from panel regressions extracting the comovement that is common across countries.

Second, we explained the joint behavior of the external and budget deficits from a small open-economy, overlapping-generation model with heterogeneous goods. Our model offers the advantage of relating the external and budget deficits to the external and internal conditions usually considered for developing economies. The model further captures different degrees of imperfectness of intergenerational linkages and of financial markets, and as such potential liquidity constraints faced by developing countries.

The model is estimated for each country such that the predicted second moments of the external and budget deficits, and in particular the covariance between these deficits, are close to their empirical counterparts. The predicted covariance is then decomposed into contributions measuring the portion attributable to each shock. The size of the contributions indicates that the shocks associated with the internal conditions, and especially the domestic resources net of public absorptions, are the most important factors explaining the positive comovement between the external and budget deficits for most countries. This contrasts with previous findings obtained from single-equation analyses characterizing the individual behavior of either the external deficit or budget deficit.

Finally, the robustness of our results could be verified by extending our tractable model to take into account production and investment decisions as well as to include fiscal and monetary rules. Such robustness analyses are left for future work.

7 References

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8 Appendix

8.1 Appendix A: Data Construction and Sources

This appendix describes the data which are mainly collected from the International Financial Statistics (IFS) released by the International Monetary Fund, and the World Development Indicators (WDI) published by the World Bank. The annual data cover the longest period since the post-1960 era for our selected developing economies. The selections of the countries, frequency, and time periods are dictated by the availability of the data, and in particular of the budget deficit and public debt. Table 3.I lists the countries and time periods for which the external and budget deficits are available. Table 3.V lists the countries for which all variables required for the estimation exercise are available.

The variables that are central to our analysis are the external and budget deficits. The external deficit refers to the negative of the nominal current account in U.S. dollars (source: IFS) converted in domestic currencies with the appropriate nominal exchange rate (source: IFS), divided by the nominal gross domestic product in domestic currencies (source: IFS). To verify the robustness of the empirical regularities, the external deficit is alternatively measured as the change in nominal net external debt or the change in nominal net external public debt, normalized by the nominal gross domestic product. For these cases, the nominal external debts in U.S. dollars (source: WDI) are converted in domestic currencies by using the nominal exchange rate. For the budget deficit, it is always measured as the nominal budget deficit in local currencies (source: IFS), normalized by the nominal gross domestic product.

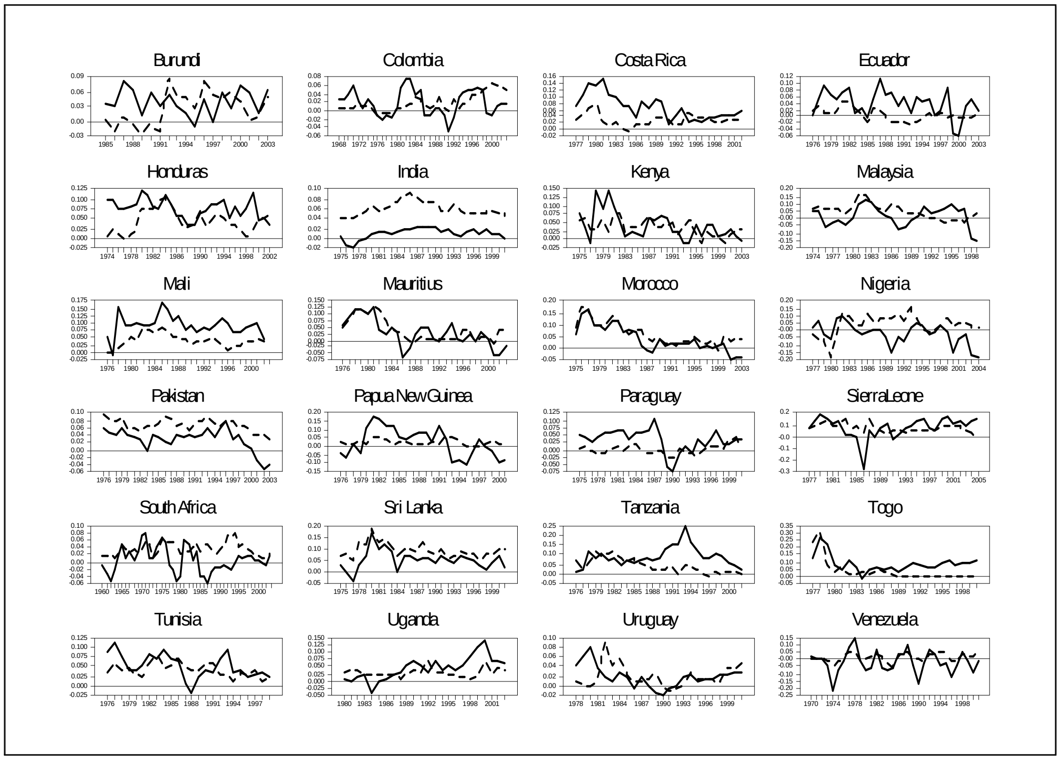
The predetermined variables are the public debt and net foreign assets. The public debt is measured by the nominal foreign and domestic public debts of central governments net of guaranteed loans in domestic currencies (source: IFS), divided by the nominal gross domestic product. The net foreign assets correspond to the nominal net foreign assets in U.S. dollars (source: Lane and Milesi-Ferretti 2007) expressed in domestic currencies, normalized by the nominal gross domestic product. Exceptionally, for Sierra Leone the nominal net foreign assets in U.S. dollars is taken from the WDI.

The forcing variables are the world interest rate, terms of trade, exchange rate, output, and government expenditures. The world interest rate is proxied by the nominal yield on three-month US treasury bills (source: IFS) minus the expected inflation, constructed as the one-step-ahead forecasts of an ARMA(1,1) process for the annual growth rate of the US consumer price index (source: IFS) (e.g. Uribe and Yue 2006). The terms of trade are measured from the ratio of export prices to import prices (source: WDI). The exchange rate corresponds to the real effective exchange rate (source: IFS). Exceptionally, for Honduras, Mauritius, and Sri Lanka the real effective exchange rate is taken from Cashin, Cespedes, and Sahay (2004), while for India the real effective exchange rate is published by the Reserve Bank of India. Output is obtained from the nominal gross domestic product, divided by the consumer price index. The government expenditures are computed as the nominal government expenditures on services, consumption goods, and investment goods in domestic currencies (source: IFS), normalized by the nominal gross domestic product.

The additional variables required for estimation purposes are taxes and private consumption, normalized by the nominal gross domestic product. Taxes correspond to the ratio of nominal tax revenues in domestic currencies (source: IFS) to nominal gross domestic product. Private consumption is obtained from the nominal household expenditures on services, non-durable goods, and durable goods in domestic currencies (source: IFS).

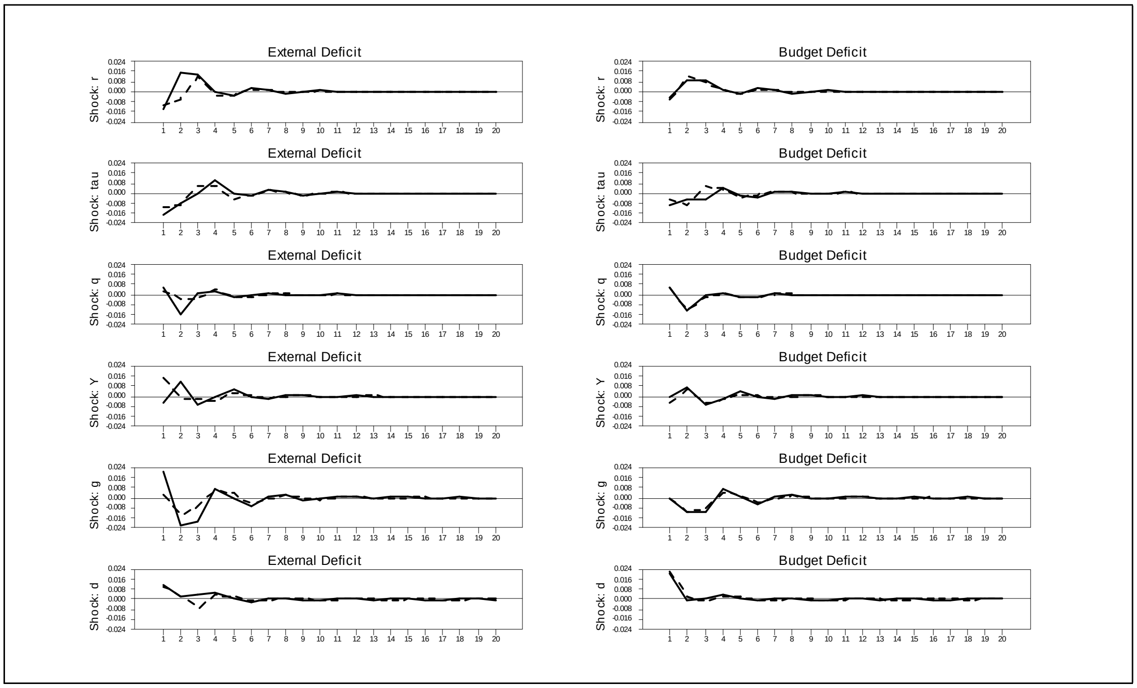
8.2 Appendix B: Figures

Figure 3.1: Transformed Data



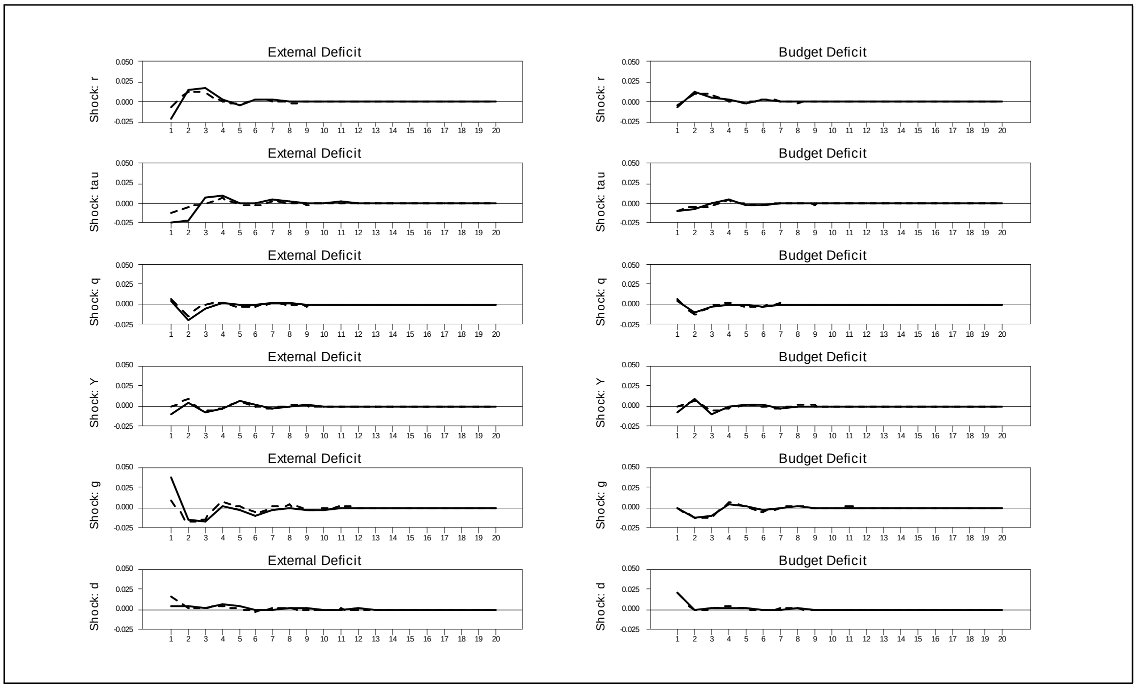
The solid (dashed) lines correspond to the external deficit (budget deficit).

Figure 3.2: Empirical and Predicted Dynamic Responses



The solid (dashed) lines correspond to the predicted (empirical) responses of external and budget deficits, obtained by estimating the restricted (unrestricted) VAR with the benchmark ordering for Sri Lanka.

Figure 3.3: Simulated Dynamic Responses



The solid (dashed) lines correspond to the simulated responses of external and budget deficits, obtained by evaluating the restricted VAR with the benchmark ordering from a parametrization involving $\rho = 0.2$ ($\rho = 0.8$).

8.3 Appendix C: Tables

Table 3.I. Empirical Regularities: Basic Statistics

Country	*cov*	*corr*	Country	*cov*	*corr*
Burundi [1985−2003]	0.061 (0.968)	0.007 (0.968)	Pakistan [1976−2003]	10.570 (0.005)	0.804 (0.000)
Colombia [1968−2003]	0.885 (0.269)	0.202 (0.262)	Papua New Guinea [1976−2001]	7.169 (0.092)	0.392 (0.038)
Costa Rica [1977−2002]	7.871 (0.016)	0.624 (0.000)	Paraguay [1975−2001]	1.832 (0.247)	0.285 (0.179)
Ecuador [1976−2003]	3.077 (0.025)	0.355 (0.003)	Sierra Leone [1977−2005]	19.250 (0.010)	0.408 (0.000)
Honduras [1974−2002]	8.013 (0.000)	0.520 (0.000)	South Africa [1960−2003]	1.621 (0.116)	0.215 (0.079)
India [1975−2001]	1.291 (0.015)	0.442 (0.007)	Sri Lanka [1975−2001]	11.810 (0.005)	0.848 (0.000)
Kenya [1975−2003]	7.884 (0.000)	0.514 (0.000)	Tanzania [1976−2002]	12.540 (0.003)	0.489 (0.000)
Malaysia [1974−1999]	9.938 (0.269)	0.241 (0.191)	Togo [1977−2000]	45.340 (0.085)	0.826 (0.000)
Mali [1976−2003]	11.480 (0.000)	0.764 (0.000)	Tunisia [1976−1999]	6.079 (0.002)	0.606 (0.000)
Mauritius [1976−2003]	19.940 (0.000)	0.741 (0.000)	Uganda [1980−2003]	0.125 (0.899)	0.027 (0.899)
Morocco [1975−2003]	15.480 (0.008)	0.823 (0.000)	Uruguay [1978−2001]	1.365 (0.044)	0.234 (0.064)
Nigeria [1977−2004]	11.600 (0.131)	0.276 (0.072)	Venezuela [1970−2001]	14.190 (0.000)	0.686 (0.000)

Note. *cov* and *corr* refer to the GMM estimates of the empirical covariance (multiplied by 10000) and correlation between the external and budget deficits. Numbers in parentheses are the p-values associated with the t test that the estimates are null, where this test involves Newey-West standard errors. Entries in brackets represent the sample periods.

Table 3.II. Empirical Regularities: Statistics for Alternative Measures

Country	$corr_1$	$corr_2$	Country	$corr_1$	$corr_2$
Burundi	0.002 (0.992)	0.045 (0.790)	Pakistan	0.716 (0.000)	0.704 (0.000)
Colombia	−0.167 (0.300)	−0.104 (0.482)	PapuaNewGuinea	0.379 (0.054)	0.398 (0.016)
Costa Rica	0.157 (0.050)	0.147 (0.074)	Paraguay	−0.254 (0.145)	−0.166 (0.371)
Ecuador	0.307 (0.028)	0.397 (0.001)	SierraLeone	0.098 (0.567)	0.102 (0.551)
Honduras	0.246 (0.009)	0.307 (0.000)	SouthAfrica	0.266 (0.090)	0.155 (0.568)
India	0.618 (0.000)	0.559 (0.000)	SriLanka	0.751 (0.000)	0.681 (0.000)
Kenya	0.490 (0.000)	0.519 (0.000)	Tanzania	0.399 (0.003)	0.288 (0.048)
Malaysia	0.554 (0.000)	0.680 (0.000)	Togo	0.436 (0.067)	0.441 (0.059)
Mali	0.341 (0.046)	0.376 (0.032)	Tunisia	0.758 (0.000)	0.758 (0.000)
Mauritius	0.715 (0.000)	0.592 (0.000)	Uganda	0.548 (0.000)	0.529 (0.000)
Morocco	0.683 (0.000)	0.689 (0.000)	Uruguay	0.310 (0.097)	0.327 (0.090)
Nigeria	0.389 (0.000)	0.372 (0.000)	Venezuela	0.012 (0.947)	-0.006 (0.969)

Note. $corr_i$ refer to the estimates of the empirical correlation between the change in net external debt and budget deficit ($i = 1$), and between the change in net external public debt and budget deficit ($i = 2$). Numbers in parentheses are the p-values associated with the test that the estimates are null.

Table 3.III. Empirical Regularities: Statistics for Alternative Methods

Country	$corr_1$	$corr_2$	$corr_3$	Country	$corr_1$	$corr_2$	$corr_3$
Burundi	−0.044 (0.859)	0.006 (0.971)	0.007 (0.968)	Pakistan	0.598 (0.001)	0.476 (0.000)	0.777 (0.000)
Colombia	0.193 (0.260)	0.127 (0.276)	0.208 (0.192)	PapuaNewGuinea	0.335 (0.080)	0.237 (0.077)	0.392 (0.038)
Costa Rica	0.490 (0.010)	0.354 (0.011)	0.553 (0.000)	Paraguay	0.148 (0.464)	0.134 (0.327)	0.189 (0.422)
Ecuador	0.375 (0.048)	0.254 (0.059)	0.355 (0.003)	SierraLeone	0.399 (0.031)	0.295 (0.025)	0.472 (0.001)
Honduras	0.613 (0.000)	0.430 (0.001)	0.520 (0.000)	SouthAfrica	0.178 (0.248)	0.123 (0.239)	0.215 (0.079)
India	0.471 (0.012)	0.362 (0.008)	0.540 (0.001)	SriLanka	0.831 (0.000)	0.663 (0.000)	0.848 (0.000)
Kenya	0.471 (0.009)	0.320 (0.015)	0.661 (0.000)	Tanzania	0.527 (0.001)	0.379 (0.006)	0.575 (0.000)
Malaysia	0.142 (0.491)	0.083 (0.552)	0.241 (0.191)	Togo	0.596 (0.002)	0.435 (0.003)	0.342 (0.027)
Mali	0.816 (0.000)	0.612 (0.000)	0.836 (0.000)	Tunisia	0.659 (0.000)	0.464 (0.002)	0.606 (0.000)
Mauritius	0.575 (0.001)	0.434 (0.001)	0.741 (0.000)	Uganda	0.044 (0.839)	0.014 (0.923)	−0.144 (0.434)
Morocco	0.650 (0.000)	0.502 (0.000)	0.823 (0.000)	Uruguay	0.515 (0.009)	0.319 (0.029)	0.381 (0.007)
Nigeria	0.202 (0.304)	0.164 (0.221)	0.217 (0.232)	Venezuela	0.714 (0.000)	0.524 (0.000)	0.669 (0.000)

Note. $corr_i$ refer to the estimates of the empirical Spearman rank correlation between the external and budget deficits ($i = 1$), the Kendall rank correlation ($i = 2$), and the correlation in the Mahalanobis distance associated with the 97.5% quantile of the $\chi^2(2)$ distribution ($i = 3$). Numbers in parentheses are the p-values associated with the tests that the estimates are null.

Table 3.IV. Economic Environment: Coefficients of Rules

$$\mathbf{\Theta}_{pp} = \begin{pmatrix} (1-\rho) & -\rho\left(\frac{\nu+\rho}{\mu_r+\rho}\right) \\ 0 & \left(\frac{1}{1+\mu_y}\right) \end{pmatrix} \qquad \mathbf{\Theta}_{zp} = \begin{pmatrix} \left[\left(\frac{\mu_r}{1+\mu_y}\right) - (\nu+\rho)\right] \\ -\rho\left(\frac{\nu+\rho}{\mu_r+\rho}\right) \end{pmatrix}'$$

$$\mathbf{\Theta}_{pf} = \begin{pmatrix} (1-\rho)\,\mu_{(f-b^*)} & 0 \\ 0 & 0 \\ 0 & 0 \\ -(1-\rho)\,\mu_{(f-b^*)} & 0 \\ e^{\mu_g}\lambda & 0 \\ -\rho\left(1+\mu_y\right)\left(\frac{\nu+\rho}{\mu_r+\rho}\right) & 1 \end{pmatrix}' \qquad \mathbf{\Theta}_{pa} = \begin{pmatrix} 1 \\ 0 \end{pmatrix}$$

$$\mathbf{\Theta}_{zf} = \begin{pmatrix} (1-\rho)\,\mu_{(f-b^*)} \\ 0 \\ 0 \\ \left(\nu+\rho-\frac{\mu_r}{1+\mu_y}\right)\mu_{(f-b^*)} \\ -e^{\mu_g}\lambda \\ -\rho\left(\frac{\nu+\rho}{\mu_r+\rho}\right)\left(1+\mu_y\right) \end{pmatrix}' \qquad \mathbf{\Theta}_{af} = \begin{pmatrix} (1-e^{\mu_t}) - e^{\mu_c}\left(1-\frac{1}{\gamma}\right) \\ -\left[e^{\mu_g}+e^{\mu_c}\left(1-\frac{1}{\gamma}\right)\right]\left(1-\frac{1}{\epsilon_\tau}\right) \\ \left[e^{\mu_g}+e^{\mu_c}\left(1-\frac{1}{\gamma}\right)\right]\left(1-\frac{1}{\epsilon_q}\right) \\ -(1-e^{\mu_t}) \\ e^{\mu_g}\left(\frac{\nu+\rho}{1+\nu}\right) \\ -\rho\left(\frac{\nu+\rho}{\mu_r+\rho}\right)\left(1+\mu_y\right) \end{pmatrix}'$$

$$\lambda = \left(\frac{1-\rho}{1+\nu}\right) \qquad \nu = \mu_r - \mu_y > 0$$

$$\epsilon_q = 1 + \left(\frac{1-\omega}{\omega}\right)e^{(1-\xi)\mu_q} \qquad \epsilon_\tau = 1 + \left(\frac{1-\varpi}{\varpi}\right)e^{(1-\zeta)\mu_\tau}$$

Note. μ_y, μ_q, μ_τ, μ_r, μ_g, μ_d, $\mu_{(f-b^*)}$, μ_c, and μ_t are the means of $\Delta\log Y_t$, $\log q_t$, $\log\tau_t$, r_t, $\log g_t$, d_t, $(f_t - b_t^*)$, $\log c_t = \log\left(\frac{P_tC_t}{Y_t}\right)$, and $\log t_t = \log\left(\frac{T_t}{Y_t}\right)$. Also, γ is the reciprocal of the elasticity of intertemporal substitution of consumption, ξ is the elasticity of substitution between tradable and non-tradable goods, ζ is the elasticity of substitution between home and foreign tradable goods, ω is the weight of tradable goods in total consumption, ϖ is the weight of home tradable goods in total tradable consumption, and ρ is the probability of death.

Table 3.V. Results: Estimates of the Preference Parameters[41]

Country	ρ	γ	ϵ_q	ϵ_τ	ω	ξ	ϖ	ζ
Costa Rica	0.427 (0.002)	2.103 (0.029)	1.711 (0.043)	1.186 (0.000)	0.3	1.261	0.3	1.560
					0.7	0.889	0.7	1.185
Honduras	0.613 (0.001)	1.523 (0.003)	1.166 (0.000)	1.032 (0.000)	0.3	1.554	0.3	1.942
					0.7	1.199	0.7	1.569
India	0.716 (0.001)	1.465 (0.001)	1.308 (0.012)	1.020 (0.000)	0.3	1.447	0.3	1.997
					0.7	1.073	0.7	1.642
Malaysia	0.391 (0.463)	1.364 (0.566)	1.013 (0.037)	1.010 (0.145)	0.3	2.064	0.3	2.301
					0.7	1.714	0.7	1.929
Mauritius	0.378 (0.002)	2.597 (0.000)	2.478 (0.007)	1.608 (0.051)	0.3	1.099	0.3	1.296
					0.7	0.732	0.7	0.923
Morocco	0.580 (0.000)	2.236 (0.002)	1.240 (0.003)	1.129 (0.000)	0.3	1.501	0.3	1.624
					0.7	1.127	0.7	1.258
Nigeria	0.846 (0.000)	1.225 (0.416)	1.400 (0.056)	1.201 (0.247)	0.3	1.349	0.3	1.560
					0.7	1.014	0.7	1.173
Pakistan	0.272 (0.035)	1.073 (0.015)	1.106 (0.323)	2.333 (0.446)	0.3	1.641	0.3	1.118
					0.7	1.290	0.7	0.761
Sierra Leone	0.863 (0.000)	2.024 (0.006)	3.131 (0.231)	2.346 (0.483)	0.3	1.019	0.3	1.158
					0.7	0.669	0.7	0.671
South Africa	0.489 (0.000)	1.398 (0.015)	1.656 (0.000)	1.615 (0.003)	0.3	1.263	0.3	1.283
					0.7	0.912	0.7	0.923
Sri Lanka	0.548 (0.000)	1.440 (0.000)	1.150 (0.000)	1.061 (0.000)	0.3	1.581	0.3	1.811
					0.7	1.222	0.7	1.434
Tunisia	0.083 (0.743)	1.754 (0.017)	1.520 (0.121)	1.341 (0.276)	0.3	1.320	0.3	1.413
					0.7	0.959	0.7	1.049

[41] Note. ρ, γ, ϵ_q, and ϵ_τ are the GMM estimates of the preference and composite parameters. ξ and ζ are obtained from the definitions $\epsilon_q = 1 + \left(\frac{1-\omega}{\omega}\right) e^{(1-\xi)\mu_q}$ and $\epsilon_\tau = 1 + \left(\frac{1-\varpi}{\varpi}\right) e^{(1-\zeta)\mu_\tau}$, given some values of the weights ω and ϖ, and the estimated values of the means μ_q and μ_τ. Numbers in parentheses are the p-values associated with the t test that the estimates are null, where this test involves Newey-West standard errors.

Table 3.VI. Results: Empirical and Predicted Moments[42]

Country	J	cov	$-\sigma_{zd}$	$cov = -\sigma_{zd}$	$corr$	$-\rho_{zd}$	$corr = -\rho_{zd}$
Costa Rica	[0.085]	7.871	7.383	(0.753)	0.624	0.686	(0.296)
		8.695		(0.397)	0.763		(0.194)
Honduras	[0.249]	8.013	12.70	(0.024)	0.520	0.854	(0.000)
		7.177		(0.008)	0.491		(0.000)
India	[0.215]	1.291	1.989	(0.038)	0.442	0.452	(0.777)
		0.882		(0.001)	0.403		(0.166)
Malaysia	[0.011]	9.938	11.80	(0.669)	0.241	0.451	(0.000)
		10.70		(0.800)	0.295		(0.000)
Mauritius	[0.740]	19.94	14.10	(0.000)	0.741	0.883	(0.237)
		6.080		(0.000)	0.328		(0.000)
Morocco	[0.137]	15.48	17.30	(0.342)	0.823	0.671	(0.083)
		12.60		(0.014)	0.764		(0.288)
Nigeria	[0.304]	11.60	18.10	(0.262)	0.276	0.489	(0.000)
		12.30		(0.317)	0.295		(0.000)
Pakistan	[0.015]	10.57	7.042	(0.364)	0.804	0.869	(0.303)
		12.10		(0.193)	0.907		(0.547)
Sierra Leone	[0.082]	19.25	16.70	(0.737)	0.408	0.427	(0.848)
		18.60		(0.803)	0.427		(1.000)
South Africa	[0.049]	1.621	3.194	(0.000)	0.215	0.639	(0.000)
		2.418		(0.000)	0.351		(0.000)
Sri Lanka	[0.049]	11.81	19.50	(0.006)	0.848	0.830	(0.753)
		8.935		(0.000)	0.738		(0.108)
Tunisia	[0.144]	6.079	3.571	(0.057)	0.606	0.391	(0.002)
		8.315		(0.004)	0.904		(0.000)

[42]Note. J refers to the J-statistics, and entries in brackets are the p-values associated with the χ^2 test that these statistics are null. cov and $-\sigma_{zd}$ are the estimates of the empirical and predicted covariances (multiplied by 10000) between the external and budget deficits. $corr$ and $-\rho_{zd}$ are the empirical and predicted correlations between the external and budget deficits. The empirical moments are computed as in Table 3.I (first line) and from the unrestricted VAR (second line). Entries in parentheses are the p-values associated with the χ^2 test that the empirical and predicted moments are identical. The tests take into account the uncertainty related to the estimates of the preference and composite parameters by using the δ-method.

Table 3.VII. Results: Decompositions of the Empirical and Predicted Covariances

Country	ψ_r	ψ_τ	ψ_q	ψ_y	ψ_g	ψ_d	ψ_z or ψ_a
Costa Rica	7.775	-9.597^{***}	-0.515	-7.450	11.640^{***}	7.072^{***}	-0.235
	1.421^{***}	0.369^{***}	0.145	0.495	4.585^{***}	0.285	0.082
Honduras	0.097	-0.401	0.499	3.321	3.500^{**}	-0.301	0.462
	1.324^{***}	0.901^{***}	2.475^{***}	4.219^{***}	3.164^{***}	0.864^{***}	0.322^{***}
India	0.521^{***}	0.003	-0.081^{***}	0.039	-0.013	0.323	0.089
	0.706^{***}	0.305^{***}	0.116^{***}	0.066^{*}	0.177^{**}	0.563^{***}	0.055^{***}
Malaysia	1.514	2.464	4.257^{**}	0.539	0.309	1.493^{***}	0.137
	2.357^{*}	3.788^{***}	1.459^{***}	1.687	1.518^{***}	0.501^{**}	0.518^{***}
Mauritius	-1.294	-4.737	0.904	3.144	7.219^{**}	0.142	0.702
	0.550^{***}	2.945^{***}	2.349^{***}	1.034^{***}	4.536^{***}	1.377^{*}	1.289^{***}
Morocco	1.251	1.143	-1.325	2.063	3.576^{*}	3.443^{*}	2.425
	0.475^{***}	0.773^{***}	0.070	1.267^{**}	10.200^{***}	4.102^{**}	0.472
Nigeria	10.48	5.458	-1.827	15.780^{***}	-22.980^{***}	7.919^{*}	-2.534
	0.529^{**}	1.311^{***}	0.255	7.867^{**}	4.239^{***}	0.359	3.492^{***}
Pakistan	1.131	4.157	-0.341	-0.085	7.235	-0.118	0.153
	0.783	2.891^{*}	0.128	0.113	2.420^{*}	0.502	0.205
Sierra Leone	2.487	-0.479	4.850^{***}	3.071	5.932^{**}	2.158	0.620
	2.217^{***}	1.040^{***}	0.515^{***}	3.935	6.781^{*}	0.812^{***}	1.376^{***}
South Africa	-0.430	0.810	0.262	2.339^{*}	1.072	-1.494^{***}	-0.142
	0.255^{***}	0.437^{***}	-0.014	1.157^{***}	0.550^{***}	0.852^{***}	-0.042^{***}
Sri Lanka	0.964	2.230	0.665	-0.648	2.713^{**}	2.678^{***}	0.332
	3.368^{***}	2.684^{***}	2.336^{***}	1.615^{***}	5.532^{***}	2.586^{***}	1.389^{***}
Tunisia	-4.257	1.972	19.93^{*}	-1.560	-1.441	-6.207^{*}	-0.123
	-0.419^{**}	0.106^{***}	1.481^{***}	0.104^{*}	1.928^{**}	0.372^{*}	-0.001

Note. $\psi^{'}s$ are the estimates of the components (multiplied by 10000) of the empirical covariance (first line) and predicted covariance (second line) between the two deficits, obtained from the unrestricted VAR and restricted VAR with the benchmark orderings. $*$, $**$, and $***$ indicate that the p-values associated with the χ^2 test that the components are null are smaller than 10%, 5%, and 1%. The test takes into account the uncertainty related to the estimates using the δ-method.

Table 3.VIII. Results: Decompositions of the Simulated Covariances

Parameter		ψ_r	ψ_τ	ψ_q	ψ_y	ψ_g	ψ_d	ψ_a
	0.200	4.429	4.320	2.407	1.926	3.715	1.167	0.601
$\rho =$	0.548	3.368	2.684	2.336	1.615	5.532	2.586	1.389
	0.800	2.705	1.772	2.618	1.300	4.868	3.571	1.150
	0.910	3.329	2.554	2.276	1.503	5.307	2.222	1.241
$\gamma =$	1.440	3.368	2.684	2.336	1.615	5.532	2.586	1.389
	3.330	3.406	2.810	2.395	1.725	5.751	2.941	1.533
	0.500	7.382	1.937	4.478	2.991	9.768	3.249	2.843
$\xi =$	1.222	3.368	2.684	2.336	1.615	5.532	2.586	1.389
	2.500	3.061	2.741	2.172	1.510	5.208	2.535	1.278
	0.500	3.622	3.650	1.806	1.388	4.800	2.084	1.233
$\zeta =$	1.434	3.368	2.684	2.336	1.615	5.532	2.586	1.389
	2.500	3.320	2.501	2.436	1.658	5.670	2.681	1.419

Note. $\psi' s$ are the components (multiplied by 10000) of the simulated covariances between the external and budget deficits, obtained by evaluating the restricted VAR with the benchmark ordering from various sets of parametrizations. These sets fix the preference parameters to the estimates for Sri Lanka (second line), and to smaller and larger values (first and third lines).

Table 3.IX. Robustness: Decompositions of the Empirical and Predicted Covariances

Country	ψ_y	ψ_g	ψ_d	ψ_r	ψ_τ	ψ_q	ψ_z or ψ_a
Costa Rica	−11.680**	9.104***	6.354	5.129	0.012	0.007	−0.235
	0.808	4.418***	0.552	1.106***	0.037***	0.380**	0.082
Honduras	3.031	1.819**	−0.146	0.032	0.320	1.658	0.462
	3.276***	4.159***	0.700***	0.573***	1.970***	1.688***	0.322***
India	0.047	0.083	0.477	0.155	−0.011	0.041	0.089
	0.062**	0.356***	1.265***	0.070	0.097**	0.084***	0.055***
Malaysia	6.636**	3.634	0.425	0.460	−0.970	0.392	0.137
	5.565*	3.131***	1.065***	0.808***	0.463**	0.277**	0.518***
Mauritius	9.247**	−5.732	−0.177	0.520	0.073	1.448	0.702
	3.453***	5.429***	1.900**	0.234	−0.084***	1.859***	1.289***
Morocco	2.494	5.275*	2.072	−0.008	0.493	−0.176	2.425
	1.495***	10.360***	5.492***	−0.362	−0.104	−0.075	0.472
Nigeria	25.900***	−27.030**	1.864	12.300	2.245	−0.454	−2.534
	4.113**	8.387***	0.712	0.802*	0.537	0.009	3.492***
Pakistan	0.930	10.190	−0.589***	1.084	0.211	0.156	0.153
	0.703***	3.880*	0.388	0.578	0.899***	0.388	0.205
Sierra Leone	5.753	10.490***	0.252	1.259	−0.045	0.307	0.620
	3.001*	8.460**	3.082**	−0.308	0.232***	0.833	1.376***
South Africa	2.289	1.785***	−1.175	0.088	0.003	−0.430	−0.142
	1.404***	0.697***	1.108***	0.040	0.006**	0.017	−0.042***
Sri Lanka	−0.859	2.122	3.483***	0.764	2.624	0.468	0.332
	2.631***	3.488***	2.918***	2.372***	4.936***	1.749***	1.389***
Tunisia	1.873	3.867	−11.280***	2.150	1.814	10.020	−0.123
	0.831***	1.761***	0.217	0.096***	0.122***	0.544***	−0.001

Note. $\psi^{'}s$ are the estimates of the components (multiplied by 10000) of the empirical covariance (first line) and predicted covariance (second line) between the two deficits, obtained from the unrestricted VAR and restricted VAR with the alternative orderings. $*$, $**$, and $***$ indicate that the p-values associated with the χ^2 test that the components are null are smaller than 10%, 5%, and 1%. The test takes into account the uncertainty related to the estimates using the δ-method.

Conclusion générale

Le premier essai de la thèse s'intéresse aux effets macroéconomiques des chocs de politique fiscale aux États-Unis. La méthodologie proposée dans ce travail consiste à identifier les chocs fiscaux en exploitant l'hétéroscédasticité conditionnelle des innovations structurelles. Cette méthodologie relâche les restrictions d'identification communément utilisées dans la littérature empirique. Ce faisant, cette approche n'impose aucune hypothèse sur l'indicateur de politique fiscale ainsi que sur le mécanisme de transmission des chocs. Conformément aux prédictions keynésiennes, les résultats obtenus montrent qu'une augmentation des dépenses gouvernementales est un outil de stimulation économique plus efficace qu'une baisse des taxes. De plus, les effets des chocs fiscaux et leurs contributions relatives dans les fluctuations de la production ont changé de manière significative durant la période post 1979. Aussi, la réponse positive et significative de la consommation privée à un choc positif des dépenses gouvernementales s'accorde avec la prédiction keynésienne. Enfin, la consommation privée est plus sensible aux chocs de dépenses gouvernementales qu'aux chocs de taxes, alors que l'inverse est vrai pour l'investissement privé.

Le deuxième essai de la thèse estime les effets des chocs fiscaux sur le compte courant et le taux de change réel pour un échantillon de quatre pays industrialisés. Les chocs sont identifiés à partir d'une méthodologie non contrainte qui se base sur les propriétés stochastiques des innovations structurelles. Les réponses du compte courant et du taux de change réel aux chocs fiscaux diffèrent d'un pays à l'autre. Les résultats soulignent très peu d'évidence empirique en faveur de l'hypothèse des déficits jumeaux. À l'exception du cas américain, les réponses du compte courant et du taux de change réel à une baisse non anticipée des taxes ne confirment pas les prédictions de certains modèles macroéconomiques de petites économies ouvertes. Enfin, nous montrons que les études antérieures sous évaluent la dépréciation réelle conditionnellement à un choc positif de dépenses gouvernementales.

Le troisième essai documente et explique le lien entre les déficits extérieur et budgétaire pour un échantillon de pays en développement. D'abord, nous nous basons sur les données historiques des déficits extérieur et budgétaire afin d'extraire le comouvement entre ces deux variables qui est spécifique à chaque pays. Empiriquement, la covariance entre les deux déficits est toujours positive et est statistiquement significative pour la majorité des pays de l'échantillon. Le comportement conjoint des déficits extérieur et budgétaire est ensuite expliqué à partir d'un modèle de petite économie ouverte à générations imbriquées avec des biens hétérogènes. Ce modèle fait intervenir les conditions internes et externes qui sont généralement considérées pour les pays en développement, en tant que déterminants des déficits extérieur et budgétaire. De plus, le modèle capte plusieurs caractéristiques des économies en développement telles que les imperfections des liens intergénérationnelles et des marchés financiers. Le modèle est estimé pour chaque pays tel que les moments du second ordre des déficits extérieur et budgétaire, et en particulier la covariance entre ces déficits, s'approchent de leurs contreparties empiriques. L'exercice de décomposition de la covariance prédite montre que les chocs associés aux conditions internes sont les principaux déterminants du comouvement positif entre les déficits extérieur et budgétaire pour la plupart des pays de l'échantillon. Ce résultat contraste avec les évidences empiriques des études à formes réduites expliquant soit le comportement individuel du déficit extérieur ou celui du déficit budgétaire. Comme pistes de recherches futures, nous proposons l'enrichissement du cadre théorique de l'analyse afin de vérifier la robustesse de nos résultats. Ces enrichissements consistent à modéliser les décisions de production et d'investissement, ainsi que les comportements des autorités monétaires et budgétaires.