# Determining the Transmission Channels of Fiscal Policy: Do Exchange Rate Regimes drive Fiscal Multipliers?

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**Abstract**

This paper examines the impact of government spending across differing exchange rate regimes. Using a panel of OECD countries from 1985s2 to 2011s2, we identify fiscal shocks as the one period-ahead error in the forecast of government spending and examine their impact on the macroeconomy across fixed and flexible exchange rate regimes. Systematic differences across regime type support the predictions of the Mundell-Fleming model, with fiscal multipliers larger under fixed exchange rates than in flexible exchange rate regimes.

Keywords: Fiscal policy, Exchange rate regimes, Fiscal multipliers, Direct projections
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# 1 Introduction

In the wake of the global financial crisis, the debate over the impact of fiscal policy has been revived. The rise in popularity of New Keynesian models expanded research into optimal monetary policy at the expense of designing optimal fiscal policy. So long as prices and wages were sticky, monetary policy was the preferred tool for economic stabilization, with fiscal policy unnecessary in the stabilization of macroeconomic shocks. With few exceptions, this vintage of macro models used for policy evaluation treated fiscal variables as exogenous.

As a consequence of this rise in popularity of monetary modelling, research into the design of optimal fiscal policy and its transmission mechanism have had limited impact on the practical analysis of economic policy. This result was evidenced through the lack of consensus among academic economists regarding the appropriate stance of fiscal policy in the immediate aftermath of the financial crisis. With policy interest rates falling towards the zero lower bound, monetary policy became ineffective in reversing the recessionary impact of the crisis. Initial policy responses, including the European Economic Recovery Plan, the American Recovery and Reinvestment Act, and the Japanese Policy Package to Address Economic Crisis, were geared towards national stimulus, through increased government spending, reduced tax rates and employment incentives.

Despite the seeming international coordination of these policies to stimulate the global economy, economic opinions regarding their effectiveness was mixed. Central to this debate is the size of the fiscal multiplier; the ratio of the change in output resulting from a change government spending to the change in government spending. Economic theory suggests that there are a number of factors that can influence the extent of the government spending multiplier, ranging from measures of financial and economic activity, to economic structures and policy regimes. However, there is a disagreement in the theoretical literature as to which of these factors determine the size of the multiplier.

The aim of this paper is to describe the dynamics of government spending shocks in determining the fiscal multiplier, by studying how fiscal transmission channels depend on the prevailing exchange rate regime. How monetary authorities respond to unanticipated changes in the level of government consumption can theoretically be a strong determining factor in the final impact of these shocks to the economy. The Mundell–Fleming model, a mainstay of macroeconomics, argues that the prevailing exchange rate regime is the most important determinant of the government multiplier. Under a flexible exchange rate regime, fiscal policy is ineffective at stimulating output, as without a monetary response, fiscal expansions fully crowd out net exports through exchange rate appreciation. Under a fixed exchange rate regime, fiscal policy can be an effective stabilization tool, if monetary expansions offset the exchange rate appreciation pressure from fiscal stimulus. In comparison, under the standard business cycle model assumptions of private consumption being governed by inter-temporal optimization, increased government spending crowds out private consumption, with the Baxter and King (1993) RBC model predicting fiscal multipliers below 1. These models do not provide for differences across exchange rate regime, due to the neutrality of money.

While research into the effectiveness of fiscal policy has advanced in recent years, there is still considerable variation in estimates of the fiscal multiplier across countries and time. Research into the role of exchange rate regimes in determining the size of the fiscal multiplier has developed in the last few years, with somewhat mixed results. Ilzetzki, Mendoza and Vegh (2011) use quarterly data for 44 developing and advanced countries in a panel SVAR framework to estimate differences in multiplier values across exchange rate regimes. They find a statistically significant long-run multiplier of 1.5 under predetermined exchange rate regimes, while the multiplier for flexible exchange rate regimes is found to be statistically indistinguishable from zero in the long-run. Born, Jüßen, and Müller (2012) also implement a panel SVAR methodology in their analysis, using semi-annual data for a panel of OECD countries. Controlling for forecasts of government spending, they estimate the fiscal multiplier to be 1.2 under pegged exchange rates and 0.75 under floating exchange rates.

Drawing on research by Ramey (2011), Auerbach and Gorodnichenko (2011), and the above work, we employ a direct-projections methodology that allows for a dynamic modelling of economic systems without the imposition of a linear global approximation to the underlying data generating process. Such an approach reduces the risk of model misspecification and allows for greater flexibility in determining causal order, relative to a vector-autoregressive approach. To derive the fiscal shocks that will be used to calculate the dynamic impact of government spending on output and a range of other macroeconomic variables, we employ real-time, one period-ahead forecasts of the growth rate of government spending. Computing the forecast error as the difference between the ex-post growth rate and the real-time forecast, this term should capture unanticipated changes in fiscal policy. We then employ these forecast errors in a dynamic panel data model to identify the long-run effects of government spending on the macroeconomy.

Our direct-projections estimates suggest an impact multiplier of 0.89 and a cumulative multiplier of 1.6 in fixed exchange rate economies versus an impact multiplier of 0.08 and a cumulative multiplier of -0.1 under flexible exchange rate regimes. The degree of monetary accommodation is also found to differ across regimes; real interest rates fall over the projected horizon in fixed exchange rate systems, while there is a positive average response of interest rates to the spending shock in flexible exchange rate economies. While these results are consistent with the theoretical responses proposed by the Mundell–Fleming model, the estimated response of net exports are at odds with the model's transmission mechanism.

The rest of this paper is structured as follows: Section 2 discusses the direct projections methodology and its comparability to a VAR structure. Section 3 details our estimation strategy. Section 4 describes the data. Section 5 reports the main results from our empirical work. Section 6 concludes.

## 2 Methodology

Beginning with Blanchard and Perotti (2002), the Structural Vector-Autoregression approach has become one of the most popular in estimating the dynamic response of the macroeconomy to fiscal policy shocks. While there are a number of reasons why SVARs are considered well-suited to modelling fiscal variables, two of the main factors supporting their use are exogeneity and contemporaneous independence. While there is a growing trend towards stricter fiscal rules designed to promote output stabilization, fiscal policy rules are considerably less rigid than their monetary counterparts. Thus, it is possible to treat shocks to fiscal variables as being exogenous with respect to output. Additionally, given the nature of fiscal decisions, the discretionary response of fiscal policy in relation to macroeconomic shocks is limited, both in absolute terms an relative to other policy tools. A further advantage of VAR-based models is that dynamic multipliers can be easily calculated and combined to form impulse response functions.

However, despite the ease of implementation and systematic modelling of multiple time series dynamics, VARs impose strict restrictions regarding the underlying data generating process for a variable of interest. Even if these restrictions hold true, it may not be the case that a VAR will deliver appropriate coefficient estimates. Zellner and Palm (1974) and Wallis (1977) show that, if the estimated model is a subset of the VAR system, individual variables will follow a VARMA system. Similarly, Cooley and Dwyer (1998) show that most RBC model dynamics follow VARMA structures, biasing VAR representations of these models. A consequence of these problems is that VAR-based impulse responses will require long lag lengths to appropriately model the dynamics of the system; an approach which may not be achievable given the relatively short time span of most macroeconomic series.

A further issue that arises when determining the appropriate statistical technique to implement is the appropriate identification of fiscal shocks. Given the effect of automatic stabilizers in the economy (particularly unemployment benefits and income taxes), the direction of causal impact between government spending and output plays an important role in determining what constitutes a fiscal shock.

## 2.1 Vector Autoregressions and the Wold Decomposition

A VAR is an $n$-equation, $n$-variable model in which each variable is explained by its own lagged values, plus current and past values of the remaining $n-1$ variables. A VAR can be considered the reduced form of a dynamic economic system involving a vector of variables, $z_t$. The structural form representation of the $VAR(p)$ process can be expressed as

$$\mathrm{A}z_t = \mathrm{B}_1 z_{t-1} + \mathrm{B}_2 z_{t-2} + \cdots + \mathrm{B}_p z_{t-p} + u_t \tag{1}$$

$$E(uu') = \Sigma_u = \begin{bmatrix} \sigma_{u_1}^2 & 0 & \cdots & 0 \\ 0 & \sigma_{u_1}^2 & \cdots & 0 \\ \vdots & \vdots & \ddots & \vdots \\ 0 & 0 & \cdots & \sigma_{u_1}^2 \end{bmatrix} \tag{2}$$

Converting this $VAR$ into its reduced form, these expressions become

$$z_t = \Phi_1 z_{t-1} + \Phi_2 z_{t-2} + \cdots + \Phi_p z_{t-p} + v_t \tag{3}$$

$$E(vv') = \Sigma_v = \begin{bmatrix} \sigma_{v_1}^2 & 0 & \cdots & 0 \\ 0 & \sigma_{v_1}^2 & \cdots & 0 \\ \vdots & \vdots & \ddots & \vdots \\ 0 & 0 & \cdots & \sigma_{v_1}^2 \end{bmatrix} \tag{4}$$

$$\text{where, } \Phi_i = \mathrm{A}^{-1}\mathrm{B}_i \text{ and } \Sigma_v = \mathrm{A}^{-1}\Sigma_u(\mathrm{A}^{-1})'. \tag{5}$$

From here, assuming the variance-covariance matrix of the reduced form residuals is symmetric and positive definite, the Choleski decomposition of the matrix will result in an upper triangular matrix, $\Gamma$, such that $\Sigma_v = \Gamma'\Gamma$. As

$$chol(\Sigma_v) = \Gamma = {\Sigma_u}^{1/2}(\mathrm{A}^{-1})', \tag{6}$$

the matrix $(\mathrm{A}^{-1})'$ can be calculated as

$$(\mathrm{A}^{-1})' = \Sigma_u{}^{1/2} chol(\Sigma_v) \tag{7}$$

where the element $\Sigma_u{}^{1/2}$ has been determined as

$$\Sigma_u{}^{1/2} = diag(chol(\Sigma_v)) \tag{8}$$

To estimate the impulse response functions, we use the definition from Hamilton (1994)

$$IRF(t,h,d_i) = \hat{E}\left(z_{t+h}\middle|v_t = d_i; z_{t-1}, \dots, z_{t-p}\right) - \hat{E}\left(z_{t+h}\middle|v_t = \underline{0}; z_{t-1}, \dots, z_{t-p}\right) \tag{9}$$

where $D$ is an $n \times n$ matrix, whose columns $d_i$ contain the unanticipated shocks and $\underline{0}$ is a $n \times 1$ vector of zeros, and $h = 0,1, \dots H$

Define

$$\mathrm{Z}_t = \begin{bmatrix} z_t \\ z_{t-1} \\ z_{t-2} \\ \vdots \\ z_{t-p+1} \end{bmatrix} \quad \mathrm{K} = \begin{bmatrix} \Phi_1 & \Phi_2 & \cdots & \Phi_{p-1} & \Phi_p \\ I & 0 & \cdots & 0 & 0 \\ 0 & I & \cdots & 0 & 0 \\ \vdots & \vdots & \ddots & \vdots & \vdots \\ 0 & 0 & \cdots & I & 0 \end{bmatrix} \quad \mathrm{V}_t = \begin{bmatrix} v_t \\ 0 \\ 0 \\ \vdots \\ 0 \end{bmatrix} \tag{10}$$

Through the combination of (3) and (10)

$$\mathrm{Z}_t = \mathrm{K}\mathrm{Z}_{t-1} + \mathrm{V}_t \tag{11}$$

$$E(VV') = \begin{bmatrix} \Sigma_v & 0 & \cdots & 0 \\ 0 & 0 & \cdots & 0 \\ \vdots & \vdots & \ddots & \vdots \\ 0 & 0 & \cdots & 0 \end{bmatrix} \tag{12}$$

Forecast values of $\mathrm{Z}_t$ can be calculated for each horizon period, h, such that

$$\mathrm{Z}_{t+h} = v_{t+h} + \mathrm{K}v_{t+h-1} + \cdots + \mathrm{K}^h v_t + \mathrm{K}^{h+1}\mathrm{Z}_{t-1} \tag{13}$$

With component forecasts, $z_{t+h}$, calculated as

$$z_{t+h} = \mathrm{v}_{t+h} + \mathrm{K}_1^1 \mathrm{v}_{t+h-1} + \cdots + \mathrm{K}_1^h \mathrm{v}_t + \mathrm{K}_1^{h+1} z_{t-1} + \mathrm{K}_2^{h+1} z_{t-2} + \cdots \mathrm{K}_p^{h+1} z_{t-p} \tag{14}$$

Assuming the eigenvalues of K all lie outside the unit circle (so that $\mathrm{Z}_t$ is covariance stationary), the recursive VAR specification can be re-expressed as an infinite vector moving-average representation

$$\mathrm{z}_t = v_t + \mathrm{K}_1^1 v_{t-1} + \mathrm{K}_1^2 v_{t-2} + \cdots + \mathrm{K}_1^h v_{t-h} + \cdots \tag{15}$$

while the impulse response function can be expressed as

$$IRF(t,h,d_i) = \mathrm{K}_1^h d_i \quad h = 0,1,2,\dots \tag{16}$$

A key assumption regarding the use of VARs for dynamic modelling lies in the use of the Wold theorem, which decomposes the vector stochastic process $\mathrm{z}_t$ into a linearly predictable component and a linearly unpredictable component. The process is then forecasted through the use of a weighted average of past forecast errors. In doing so, it matches the first and second moments of the data to the infinite vector moving-average representation. However, the Wold representation mat not be the true representation of the $\mathrm{z}_t$ process. If not, VAR estimates of the process will be biased.

While the Wold decomposition theorem can be applied to any covariance stationary process, there is a large class of models for which the generated data has a non-invertible moving average component. If any of the eigenvalues of K lie inside the unit circle, no VAR representation of the series exists, using current and past values of the endogenous variables, where the innovations are fundamental, representing the residual series resulting from the data generating process. However, there will always exist an invertible version of the series, with identical mean and autocovariance generating function (the z-transform of a sequence of autocovariances). Thus, VAR representations may recover the fundamental innovations in the invertible series, which may not correspond to the true economic shocks from the fundamental innovations in the non-invertible series.

Numerous examples of the importance of the invertibility of stochastic processes for VAR inferences are present in economic literature. Lippi and Reichlin (1993) modify the underlying DGP in Blanchard and Quah's (1989) structural model of output and unemployment, incorporating a non-invertible moving average series in the equation representing the dynamic response of productivity increases. Estimating the modified MA representation using the same VAR structure as Blanchard and Quah, they find output and unemployment impact-responses to supply shocks of opposite sign to the invertible MA representation, while the output response to demand shocks from the modified representation is less than one-quarter of the responses from the invertible representation.

Tenhofen and Wolff (2010) develop a simple DSGE model incorporating fiscal policy anticipation from the private sector. As the vector stochastic process cannot be accurately represented using only current and past values of the endogenous variables, moving average representation of the series contain a non-invertible component. Thus, VAR estimation recovers non-fundamental innovations from the invertible representation of the stochastic process. Simulating data from this model, the theoretical impulse response show a response of consumption in period $t-1$ to a predictable government spending shock in period $t$, falling in $t-1$ due to the negative wealth effect, then increasing over the model's horizon. In comparison, the SVAR representation estimates a statistically and economically insignificant response of consumption to the same shock, with almost no impact response in $t-1$.

An additional complication arises in the infinite vector moving-average representation of the VAR. Under this representation, the variance-covariance matrix of the reduced form residuals

$$E(vv') = \Omega_v \tag{17}$$

will most likely not be a diagonal matrix. As a result, there will exist a contemporaneous correlation between forecast errors.

## 2.2 Direct Projections

Given the issues presented above regarding the estimation of dynamic systems with Wold decomposition-based models, an alternative technique that is more robust to misspecification of the data generating process may yield more accurate insights into the macroeconomic impact of fiscal shocks. The approach taken here is to employ the single-equation direct projections approach developed by Jordà (2005) and Stock and Watson (2007). This method has recently been used by Auerbach and Gorodnichenko (2012) to examine differences in fiscal multipliers over the business cycle, with their results indicating that government spending multipliers are larger during recessionary economic periods.

The direct projections method uses a sequential regression structure to estimate the endogenous variable of interest for each forecast horizon period, as opposed to the recursive iteration methodology employed by a VAR. Direct forecasting methods have been show by Bhansali (1997) and Ing (2003) to outperform iterative methodologies for autoregressive models with artificially short lag length.

Using the notation from the previous section, the direct projection method projects the H step-ahead forecast of the exogenous vector of variables, $z_{t+h}$, and projects them onto the state space $(z_{t-1}, z_{t-2}, \dots, z_{t-p})'$, in the form

$$z_{t+h} = \Psi_1^{h+1} z_{t-1} + \Psi_2^{h+1} z_{t-2} + \cdots + \Psi_p^{h+1} z_{t-p} + \mathrm{u}_{t+h}^{h} \qquad \mathrm{h} = 0,1, \dots \mathrm{H} \tag{18}$$

where $\Psi_i^{h+1}$ is a matrix of coefficients for the $i^{th}$ lag of the vector of variables $z_t$ in horizon period $\mathrm{h}+1$. Comparing equations (15) and (18),

$$\Psi_i^{h+1} = \mathrm{K}_i^{h+1} \quad \forall \ i = 1 \dots p \tag{19}$$

$$\mathrm{u}_{t+h}^{h} = \mathrm{v}_{t+h} + \ \mathrm{K}_1^1 \mathrm{v}_{t+h-1} + \cdots + \mathrm{K}_1^h \mathrm{v}_t \tag{20}$$

it is apparent that, when the VAR accurately represents the true DGP, impulse responses estimated by each technique will be equivalent.

To further show the difference in how each technique calculates impulse responses, consider the AR(1) process

$$y_t = \alpha y_{t-1} + \varepsilon_t + error = \sum_{h=0}^{H} \alpha^h \varepsilon_{t-h} + error \tag{21}$$

where $\varepsilon_t$ is a structural shock and $error$ is a set of unidentified innovations. Under the approach employed in the VAR methodology, a single estimate of the process is calculated,

$$y_t = \hat{\alpha} y_{t-1} + \varepsilon_t + error \tag{22}$$

For each time period in the forecast horizon, a recursive calculation is made using the this estimate of $\alpha$, so that

$$h = 0 \colon y_t = \hat{\alpha} y_{t-1} + \varepsilon_t + error \tag{23}$$

$$h = 1 \colon y_{t+1} = \hat{\alpha} y_t + \varepsilon_{t+1} + u_{t+1} = \hat{\alpha}^2 y_{t-1} + \hat{\alpha}\varepsilon_t + \varepsilon_{t+1} + error$$

$$h = 2 \colon y_{t+2} = \hat{\alpha} y_{t+1} + \varepsilon_{t+2} + u_{t+2} = \hat{\alpha}^3 y_{t-1} + \hat{\alpha}^2\varepsilon_t + \hat{\alpha}\varepsilon_{t+1} + \varepsilon_{t+2} + error$$

$$\vdots$$

$$h = H \colon y_{t+H} = \hat{\alpha} y_{t+H-1} + \varepsilon_{t+H} + u_{t+H} = \hat{\alpha}^{H+1} y_{t-1} + \hat{\alpha}^H \varepsilon_t + \sum_{s=0}^{H-1} \alpha^s \varepsilon_{t+H-s} + error$$

From this recursive calculation, the impulse response function is calculated as

$$I\hat{R}F = \{1, \hat{\alpha}, \hat{\alpha}^2, \dots, \hat{\alpha}^H\} \tag{24}$$

In contrast, the direct projections approach estimates an equation for each time period $h = 0, \dots, H$

$$\begin{gathered} h = 0 \colon y_t = \varphi_0 y_{t-1} + \hat{\beta}_0 \varepsilon_t + error \\ h = 1 \colon y_{t+1} = \varphi_1 y_{t-1} + \hat{\beta}_1 \varepsilon_t + error \\ \vdots \\ h = H \colon y_{t+H} = \varphi_H y_{t-1} + \hat{\beta}_h \varepsilon_t + error \end{gathered} \tag{25}$$

From each equation, the estimated coefficient attached to $\varepsilon_t$ is used to generate the impulse response function

$$I\hat{R}F = \{\hat{\beta}_0, \hat{\beta}_1, \hat{\beta}_2, \dots, \hat{\beta}_h\} \tag{26}$$

Under the null hypothesis for both projection methods, the calculated impulse response functions are estimates of the true impulse responses

$$IRF = \{1, \alpha, \alpha^2, \dots, \alpha^H\} \tag{27}$$

Under the employed DGP, the direct projections method and the VAR return the same impulse response functions. However, the real benefits to using the direct projections method result when the true DGP does not possess a VAR representation and thus the VAR cannot recover the true impact of the economic shock $\varepsilon_t$.

## 2.3 Simulation and Model Performance

While the theoretical properties of the direct projection approach show distinct advantages to its use over implementing a standard VAR methodology, it is the empirical performance of the method that determines its relative merits. Jordà (2005) runs a Monte-Carlo simulation to compare both methodologies under correct specification of the DGP and misspecification. Using monthly data from 1960-2001, Jordà estimates a VAR with lag-length 12 to model the movements in the money policy.

Fitting both a VAR and a direct projection model with two lags of the endogenous variables, impulse responses are estimated for a shock to the federal funds rate and compared against the true responses. While the VAR(2) responses are found to have the same general shape as the true response, the direct projections' responses provide a greater degree of accuracy in representing detailed movements in the responses of the endogenous variables over the entire horizon period. In particular, the VAR predicts a positive response in the price level to the shock for 23 out of the 24 horizon periods, with a statistically different response to the true model (which estimates a negative price response) for the first 17 periods. In comparison, responses generated by the direct projections model are within two standard error bands for the entire horizon period, with smaller prediction errors than the VAR responses across all time periods.

## 2.4 Identification of Fiscal Shocks

A critical element in estimating the fiscal multiplier lies in the choice of technique for identifying fiscal shocks. As automatic stabilizers act as a transmission channel for output shocks to impact on government spending, identifying truly exogenous changes in government spending is a non-trivial task.

The approach taken in this paper is based on the Ramey (2011) use of professional forecasts to control for anticipated changes in government spending. The use of forecasts in determining fiscal shocks has a number of desirable properties that provide clear benefits to alternative approaches. Auerbach and Gorodnichenko (2012) calculate correlations between the unpredictable component of government spending forecasts and errors from a government spending equation in an SVAR. A significant correlation of 0.36 is estimated between both series, indicating that a large component of government spending shocks estimated by VARs is anticipated, potentially introducing a considerable source of bias into impulse response functions generated by VAR shocks. Similarly, Ramey (2011) shows that both one and four-quarter-ahead professional forecasts Granger-cause shocks generated by VARs, leading to potential anticipation bias in impulse response functions generated by VAR shocks.

It should be noted, however, that there are certain drawbacks to using professional forecast data to generate fiscal shocks. Primarily, there is limited real-time forecast data available at an international level. As a result, this places a constraint on the sample data that can be used to estimate the impact of fiscal shocks on the macroeconomy. In addition, there will be a disassociation between the model used by the forecaster and the model employing the forecast errors as an exogenous shock. While forecast values may anticipate VAR shocks, this does not imply certainty of their representation of the true residual values from the underlying DGP.

# 3 Empirical Strategy

## 3.1 Empirical Framework

Following Auerbach and Gorodnichenko (2012), we estimate a panel direct projection model to identify the impact of unanticipated shocks to government spending on the macroeconomy. While the primary purpose of this research is to compare the size of the fiscal multiplier across exchange rate regimes, the transmission channels through with the estimated effects (if any) occur is also of considerable interest. The empirical model includes six variables; government spending, $g_{i,t}$ , forecast errors of government expenditure, $fe_{i,t}$ , output, $y_{i,t}$ , interest rates, $r_{i,t}$ , exchange rates, $ex_{i,t}$ , and net exports, $nx_{i,t}$ .

Our values of government spending, output and the forecast error are measured as growth rates. To obtain impulse responses that measure the level response of output to government spending, rather than the percentage change, we scale these variables by the ratio of their lagged value to the lagged value of output. Net exports are measured as a ratio to output. Interest rates are measured in percentage points, while exchange rates are measured in logs. All series are measured in real terms.

We specify government spending, forecast errors of government spending and output in the same units to prevent biasing our estimates of the fiscal multiplier. Traditional VAR analysis of the fiscal multiplier estimates the IRF for the response of output to fiscal shocks, then scales the responses by the sample average ratio of output to government spending. Owyang, Ramey and Zubairy (2012) find that using a sample average, rather than adjusting at each point in time, can lead to positive bias in estimates of the fiscal multiplier, due to variance in the ratio of output to government spending over time.

We estimate the system equation at a time, regressing the variable of interest onto current and lagged values of the variables in the system. To determine the impact of government spending on output at horizon $h$, we estimate the regression

$$y_{i,t+h} = \alpha_h + \Gamma_h fe_{i,t} + \Gamma_h^{Fix} fe_{i,t} + \Sigma_{\tau=1}^{\mathrm{T}} \Theta_{h,\tau} y_{i,t-\tau} + \Sigma_{\tau=1}^{\mathrm{T}} \Theta_{h,\tau}^{Fix} y_{i,t-\tau} + \Sigma_{\tau=1}^{\mathrm{T}} \Pi_{h,\tau} g_{i,t-\tau} + \Sigma_{\tau=1}^{\mathrm{T}} \Pi_{h,\tau}^{Fix} g_{i,t-\tau} + \Sigma_{\tau=1}^{\mathrm{T}} \Upsilon_{h,\tau} r_{i,t-\tau} + \Sigma_{\tau=1}^{\mathrm{T}} \Upsilon_{h,\tau}^{Fix} r_{i,t-\tau} + \Sigma_{\tau=1}^{\mathrm{T}} \Phi_{h,\tau} ex_{i,t-\tau} + \Sigma_{\tau=1}^{\mathrm{T}} \Phi_{h,\tau}^{Fix} ex_{i,t-\tau} + \Sigma_{\tau=1}^{\mathrm{T}} \Psi_{h,\tau} nx_{i,t-\tau} + \Sigma_{\tau=1}^{\mathrm{T}} \Psi_{h,\tau}^{Fix} nx_{i,t-\tau} + u_{i,t,h} \tag{28}$$

where $i$ and $t$ index countries and time periods.

Our equations can be viewed as containing two components. The first component, containing the $\Gamma_h^j$ coefficients, dictates the dynamics of the system. By estimating a separate regression for each horizon period, we generate the impulse response functions

$$\mathrm{IRF}^{Flex} = \{\Gamma_h\}_{h=0}^{H} \qquad \mathrm{IRF}^{Fix} = \{\Gamma_h + \Gamma_h^{Fix}\}_{h=0}^{H} \tag{29}$$

for flexible and fixed exchange rate regime countries respectively.

The second component, containing the set of coefficients on the lagged variables $\{\Theta_{h,\tau}^j, \Pi_{h,\tau}^j, \Upsilon_{h,\tau}^j, \Psi_{h,\tau}^j\}$, controls for the history of the system. By including these terms in the regression, any element of the forecast error that could have been predicted from lagged values of the variables in the system is eliminated from the dynamic component of the estimation. This increases the likelihood that the $fe_{i,t}$ term represents unanticipated shocks to government spending, as any predictable component in the forecast error series that would have been removed through VAR analysis will be removed here.

There are a number of advantages in estimating the above system using a direct projections approach. As discussed above, we convert government spending and our measure of fiscal shocks into output units, allowing for direct estimation of the fiscal multiplier (rather than the conversion of the elasticity of output with respect to government spending to a multiplier via sample averages). Such a transformation is considerably more difficult under a VAR approach. As we are only interested in the impact of fiscal shocks on the system of variables, we do not need to produce the set of responses to a shock for each variable in the system, greatly reducing the number of parameters to be estimated. Finally, as the direct projections approach estimates the coefficient on the fiscal shock for each horizon period, it does not constrain the shape of the IRF in the same manner that a VAR would. Thus any misspecification of the DGP should result in less of a bias in the estimates from the direct projection method.

## 3.2 Definition of Fiscal Multipliers

Across the fiscal policy literature, there are several methods used to measure the fiscal multiplier. As stated earlier, the standard definition of the multiplier is the ratio of the change in output resulting from a change in government spending to the change in government spending. However, this definition does not specify the time horizon over which the multiplier is calculated.

To maintain consistency with as broad a range of research as possible, we focus on two specific definitions of the fiscal multiplier, as per Ilzetzki, Mendoza and Vegh (2011) and Corsetti, Meier and Müller (2011). The *Impact Multiplier*, defined as

$$\text{Impact Multipler} = \frac{\Delta y_0}{\Delta g_0} \tag{30}$$

measures the ratio of the output change resulting from a change in government spending in the initial horizon period, $h = 0$. The second version of the multiplier we examine measures the cumulative change in output resulting from the combined change in government spending resulting from the initial fiscal innovation. This measure, the *Cumulative Multiplier*, is defined as

$$\text{Cumulative Multipler} = \frac{\sum_{h=0}^{H} \Delta y_0}{\sum_{h=0}^{H} \Delta g_0} \tag{31}$$

As there are $H + 1$ possible cumulative multipliers that can be considered, we will focus on the cumulative multiplier that includes all $H$ horizon periods in our analysis.

# 4 Data

The macroeconomic data used in this paper comes primarily from the OECD's Statistics and Projections database. The database provides consistent coverage of a wide range of macroeconomic data, with real-time semi-annual data and forecasts available for a large panel of OECD economies. As there are few institutions that provide real-time and forecast macroeconomic data, the OECD data is unique in its cross-country availability.

The OECD forecasting process uses a combination of expert judgement from policy makers and governments with analytical economic models to produce forecasts of a number of macroeconomic series. The forecasts employ recent changes to a range of relevant variables (including commodity prices, exchange rates, fiscal trends and output) by incorporating new data into simulations, using the NIGEM DSGE model and short-term indicator models, taking the previous projections as a starting point.

In their forecasting of fiscal variables, the OECD uses the primary cyclically-adjusted budget balance as the baseline measure for discretionary fiscal policy changes. In addition, measures of government debt, including gross government debt and general government gross financial liabilities, are incorporated into the forecast assessment, as they act as indicators of fiscal sustainability and the degree of manoeuvrability in fiscal policy. Forecast data is available on a semi-annual basis from 1985s2 for a set of seven older member countries, and for an additional 19 newer members from 1996s1. Our sample runs until 2011s2. We measure our fiscal shock as the forecast error for the growth rate of government spending at time $t - 1$ for period $t$.

In addition to the OECD data, we use data from two other sources in our regressions. Our exchange rate measure comes from the narrow real exchange rate index of the Bank for International Settlements, calculated as geometric weighted averages of bilateral exchange rates, adjusted by relative consumer prices. Finally, our classification of exchange rate regimes is based on the Ilzetzki, Reinhart, and Rogoff (2008), with the exception of Canada, which we consider to be a flexible regime across the sample. While they argue that the Canadian dollar can be seen as following a crawling peg against the US dollar of $\pm 2\%$, the exchange rate series over the sample period shows a trough to peak appreciation of 37.8% and a peak to trough depreciation of 37.5%. Given the scale of these changes, it is unlikely that monetary expansion (contraction) was consistently used to prevent exchange rate depreciations (appreciations) beyond those observed in the data, following periods of expansionary (contractionary) fiscal policy. Thus, from a theoretical perspective, we do not consider it appropriate to include Canada a fixed exchange rate regime.

Table 1 presents summary statistics for our data, across fixed and flexible exchange rate subsamples. Means and standard deviations are broadly similar, indicating that there are no obvious structural differences across the subsamples, other than exchange rate regime.

## 5 Results

The primary focus of our research is estimating the economic response to government spending that has no causal relationship to the state of the economy. As such, it is critical that our estimates of fiscal policy shocks no are unit roots or common stochastic drifts with the other endogenous variables in our model. Thus, prior to our main econometric analysis, we test for unit roots across our sample. Results are shown in Table 2.

Results from the Fisher-type unit-root tests suggest the presence of non-stationarity in our real exchange rate and net export series. We control for this by detrending using the Hodrick-Prescott filter, with a smoothing parameter of $\lambda = 400$. Testing the detrended versions, both variables are found to be stationary at the 1% level. All other series are I(0) in levels.

While our series are free from unit roots, it is likely that the error terms in our regressions are correlated across countries, as well as potentially being correlated across time. In addition, our regression equation contains lags of the dependent variable as explanatory variables. Given these problems have been shown to produce downward bias in standard errors and upward bias in t-statistics, we use the panel corrected standard error model of Beck and Katz (1995), controlling for pairwise covariance of the error terms across panels.

## 5.1 Baseline Results

Figure 2 presents the impulse response of government spending to the unanticipated fiscal shock. As per Auerbach and Gorodnichenko (2011) and Corsetti, Meier and Müller (2011) Impulse responses are normalized, so that the initial increase in government spending is one percent of output, for both exchange rate regimes. For each regime, this requires a shock greater than one percent of output, as the response of government spending to the shock is 0.88 under flexible regimes and 0.85 under fixed regimes.

Impulse responses for both regimes show dynamics similar to those observed by Ramey (2011) using forecast errors derived from the Survey of Professional Forecasters. Spending shocks are found to be transitory in nature; initial spikes in government spending are followed by a minor reduction in spending in the following period, after which spending patterns appear to return to normal. Testing the restriction that the responses are identical within horizon periods, the only statistical differences are found in $h = 1$, where the reduction in spending following the initial shock is greater under fixed exchange rates, and $h = 2$. Consistent with the idea that these shocks are transitory, tests that the government spending responses beyond the first horizon period are jointly zero cannot be rejected in either exchange rate regime.

Figure 3 displays the response of GDP to a government spending shock equal to one percent of output. The impulse responses are consistent with the Mundell-Fleming model, showing a larger and more significant response across the horizon period under fixed exchange rates. The impact multiplier in fixed exchange rate regimes is strongly positive (0.89) and significant, while the estimate of the cumulative multiplier is 1.6 and significant. Against these results, both the impact and cumulative multipliers are found to be small and insignificant under flexible exchange rates. While the impact multiplier is mildly positive (0.08), suggesting considerable crowding-out due to exchange rate flexibility, the cumulative multiplier is marginally negative (-0.1). However, at no point in the forecast horizon is the response of output to the government spending shock found to be significant under flexible exchange rates.

Turning to the impact of the government spending shock on short-term real interest rates, Figure 4 again shows results consistent with the Mundell-Fleming model. As real interest rate responses reflect the joint response of both nominal interest rates and inflation, they represent a comprehensive indicator of the stance of monetary policy following shocks to government spending. Overall, real interest rates are found to decline in fixed exchange rate regimes, with an average reduction of 0.61 percentage points over the horizon period. In contrast, real interest rates rise initially in flexible regimes, before a minor reduction and return to trend by the end of the horizon period. The average response over the horizon period is 0.006, indicating a greater degree of monetary accommodation under fixed exchange rate regimes, as per standard monetary theory assumptions. Ilzetzki, Mendoza and Vegh (2011) and Born, Jüßen, and Müller (2012) also observe increased interest rates across flexible exchange rate regimes and interest rate decreases under fixed exchange rates.

The response of real exchange rates is displayed in Figure 5. Following a 1% of output shock to government spending, our model estimates a large initial appreciation in the real exchange rate under the flexible regime of 1.25 percentage points, declining over the horizon period and eventually becoming negative after two years. In contrast, there is considerably less movement in the real exchange rate under the fixed rate regime. The initial response to the fiscal shock results in a minor appreciation of 0.34 percentage points. While the Mundell-Fleming model suggests no movement in exchange rates, this can be explained by the fact that a number of de facto fixed exchange rate economies use a crawling peg system with a $\pm 2\%$ band. Such a system allows for some degree of exchange rate movement, consistent with the result we find. This also explains the small degree of crowding-out we observe in the initial response of output under a fixed exchange rate. While the direction of movement in the real exchange rate is consistent with our earlier results, the results are not found to be significant over the horizon period.

While the previous results are all found to be consistent with the Mundell-Fleming model's predicted responses to a fiscal shock, we find an inconsistency in the response of net exports, shown in Figure 6. Despite the theoretical prediction and empirical evidence of an appreciation in the real exchange rate under a flexible regime, the initial response of net exports is to increase by 0.61 output units, steadily declining over the next two years before returning to trend. The response of net exports under a fixed regime is consistent with our earlier findings, falling initially in response to the real exchange rate appreciation, before becoming positive as real exchange rates depreciate. However, these results are not statistically significant.

Despite the inconsistency of the response of net exports to economic theory, these results have been observed in other research. Ilzetzki, Mendoza and Vegh (2011) observe depreciation in real exchange rates under fixed regimes following a fiscal shock, while at the same time net exports decline, although neither result is significant. Similarly, Born, Jüßen, and Müller (2012) observe short-run reductions in net exports across exchange rate regimes, despite opposing short-run movements in exchange rates. Impact responses in net exports are not significant in either regime, although the long run response under flexible rates is significant. These differences clearly suggest that further analysis into the current account as a transmission channel is warranted.

## 5.2 Sensitivity Analysis.

To test the robustness of our results, we make two alternate changes to our empirical specifications. A primary concern is that fiscal policy has changed since the financial crisis, altering the transmission of fiscal shocks and both the speed and size of their impact on the macroeconomy. We control for this by interacting our measure of fiscal shocks with a dummy variable, equal to one for all time periods post 2007s2. Results are shown in Figure 6.

The response of government spending to a fiscal shock remains virtually unchanged, suggesting that the nature of the spending shocks captured by our forecast error measure has remained unchanged. The cumulative impact of a fiscal shock of 1% is to raise government spending by 1% (0.9% in our baseline results) under flexible exchange rates and by 1.06% (1% baseline) under fixed exchange rates.

A slight change is observed in the response of output, with movements in both impact and cumulative multipliers. For fixed exchange rate regimes, the impact multiplier rises to 1.14 (0.89 baseline), while the cumulative multiplier falls to 1.49 (1.6 baseline). Similarly, the impact multiplier in flexible exchange rate regimes rises to 0.42 (0.09 baseline), while the cumulative multiplier falls to -0.25 (-0.1 baseline). While these results indicate that impact multipliers are weaker during financial crises but cumulative multipliers are stronger, the results from our baseline analysis, that fiscal multipliers are larger under fixed exchange rates than under flexible exchange rates, still hold.

Regarding the effects on the transmission channels of fiscal shocks, the directional effects are all found to be similar to those obtained under the baseline scenario. The estimated monetary policy stance is virtually unchanged from the baseline results, although the significance of the results under fixed exchange rates is somewhat reduced. However, monetary policy is still found to be more accommodating under a fixed exchange rate regime, with a difference in the average policy response of 0.62 percentage points between both regimes.

As our second robustness check, we remove Austria, the Czech Republic and Poland from our sample and re-estimate the baseline model. These countries all have gaps in the forecast of government spending growth rates. By removing these countries, we increase the number of time periods common to all countries from 5 to 18, allowing for a casewise estimation of the covariance matrix of inter-panel error terms when controlling for serial correlation. Impulse response functions for this specification are presented in Figure 7.

Results from this specification increase the magnitude of the gap between fiscal multipliers under fixed and flexible exchange rate regimes. Impact multipliers rise to 0.15 under flexible rate regimes and 1.25 under fixed rate regimes. Similarly, the cumulative multiplier rises to 2.67 under fixed exchange rates and falls to -0.6 under flexible exchange rates. This can be explained by the increased divergence in the degree of accommodation of monetary policy over the horizon period. Under the fixed regime, the reduction in real interest rates averages 0.77 percentage points in each period, while there is an average increase in real interest rates under the flexible regime of 0.37 percentage points.

# 6 Conclusion

Economic theory is divided on the role of exchange rate regimes in the transmission of fiscal policy. We test the theoretical implications of the Mundell-Fleming model regarding the effect of exchange rate regimes in determining the magnitude of the fiscal multiplier. Using panel direct projection methods and real-time forecast errors in the growth rate of government spending as our definition of fiscal shocks, we find strong evidence to suggest that fiscal multipliers are considerable larger under fixed exchange rate regimes, with almost no output impact resulting from increased government spending under flexible exchange rate regimes.

Our evidence indicates that interest rates are a key transmission channel for these fiscal shocks. Monetary policy is found to be more accommodating under fixed exchange rates, with short-run real interest rates declining sooner, and by a greater amount than under flexible rate regimes. Under flexible exchange rates, short-term interest rates are found to increase following a fiscal shock, accounting for the limited impact of the shock on output. These results provide further evidence supporting the theoretical arguments proposed in the Mundell-Fleming model.

With austere fiscal policy prevailing in Eurozone economies, our results suggest that cuts to government spending could be counterproductive, potentially reducing output and aggravating the recessionary impacts that have affected most of Europe in the wake of the global financial crisis. However, given the fact that interest rates remain close to the zero lower bound, monetary policy tools may not be an available to accommodate expansionary fiscal policy.

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# Tables and Figures

**Figure 1: Fiscal Shocks in Percentage Output Units, Selected Countries**

**Figure 2: Response of Government Spending to Fiscal Shock**

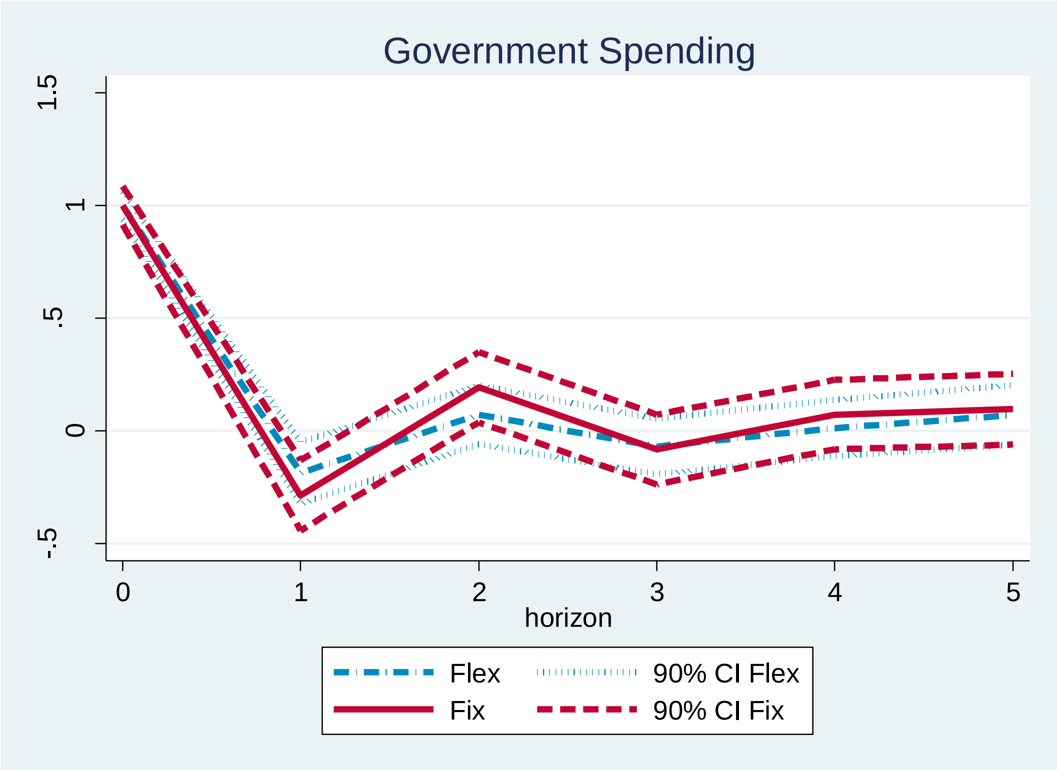



**Figure 2A: Cumulative Response of Government Spending to Fiscal Shock**

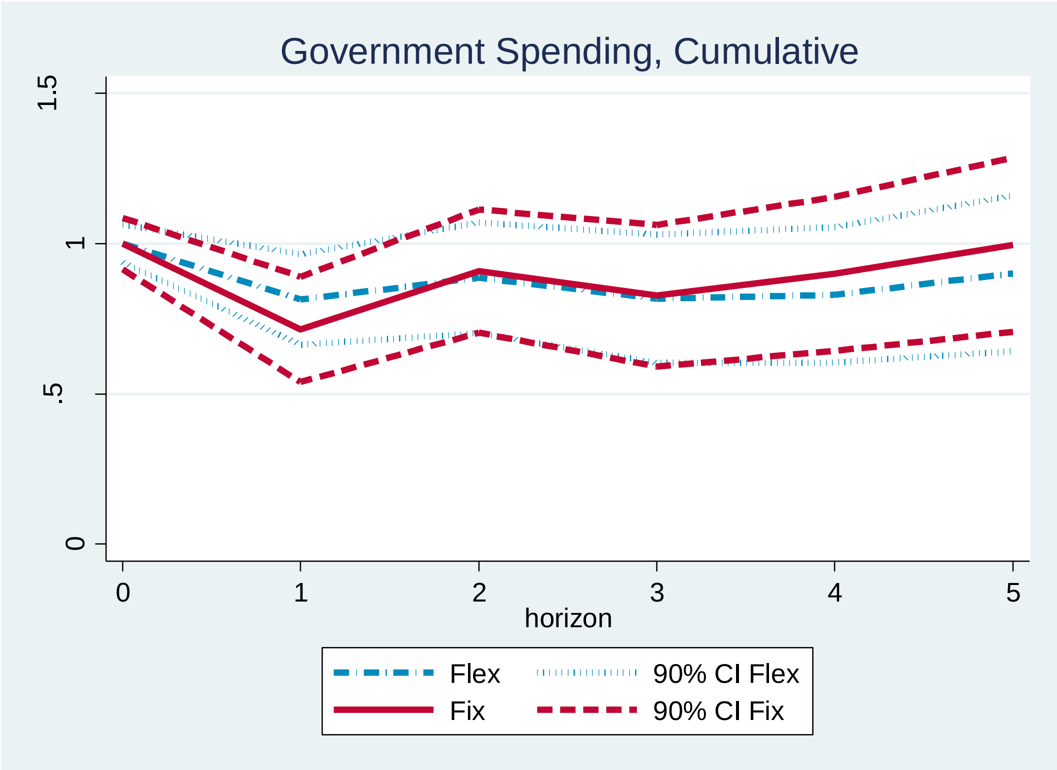


**Figure 3: Response of Output to Fiscal Shock**

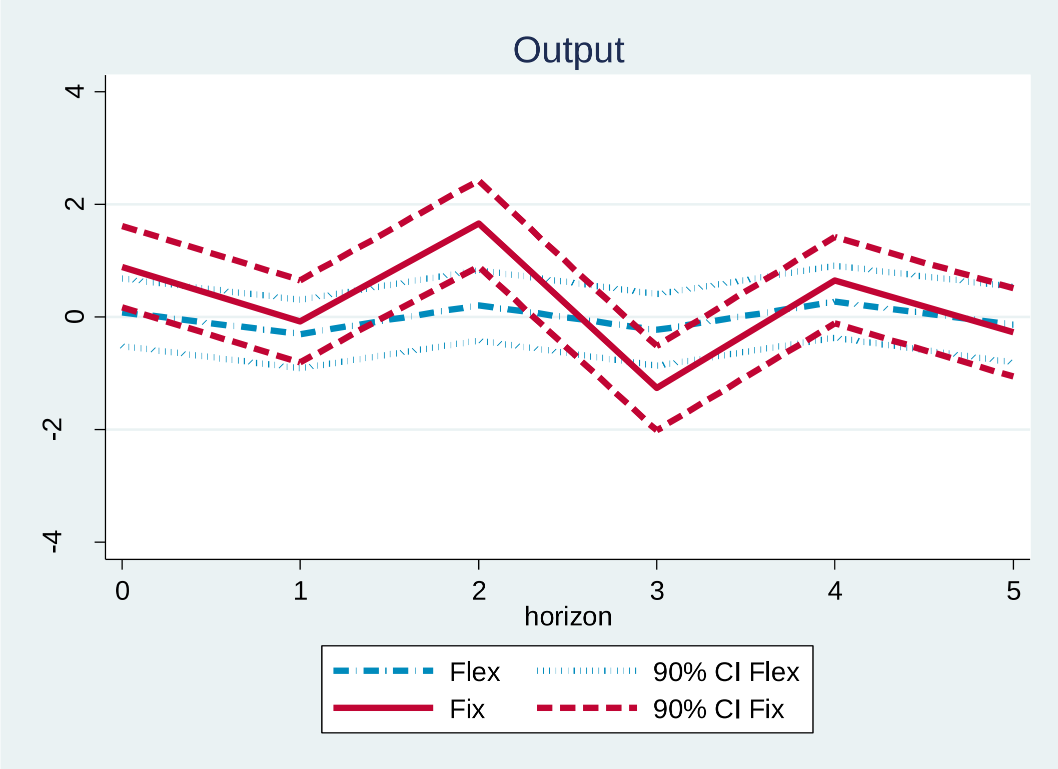
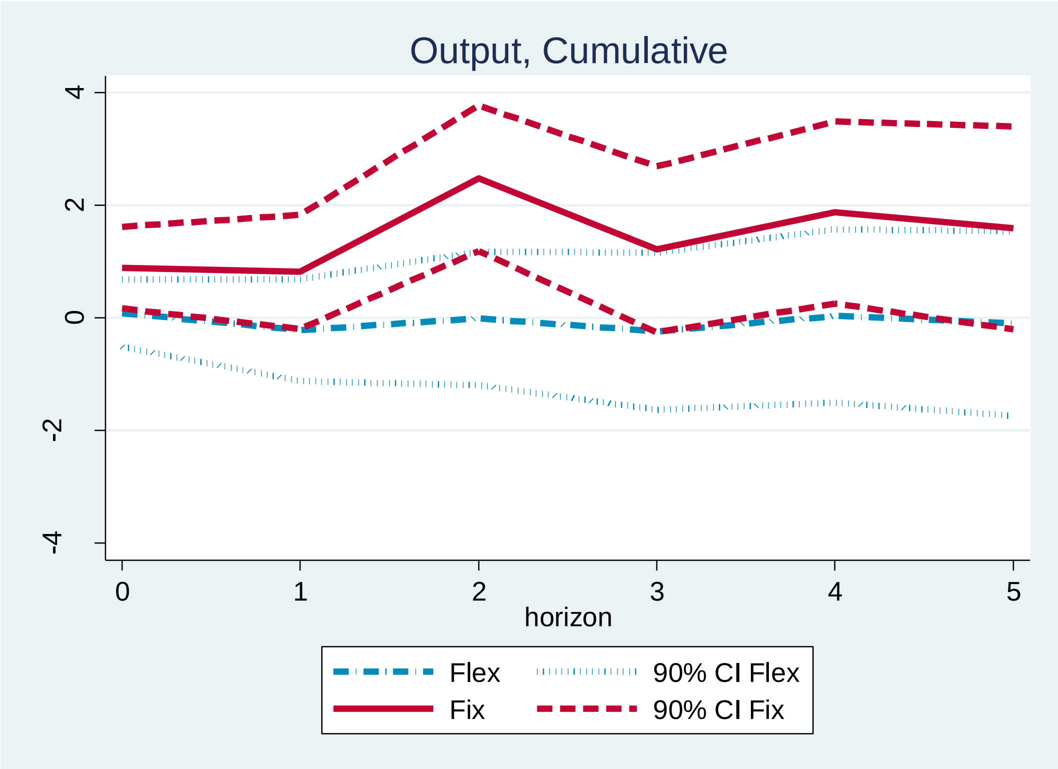
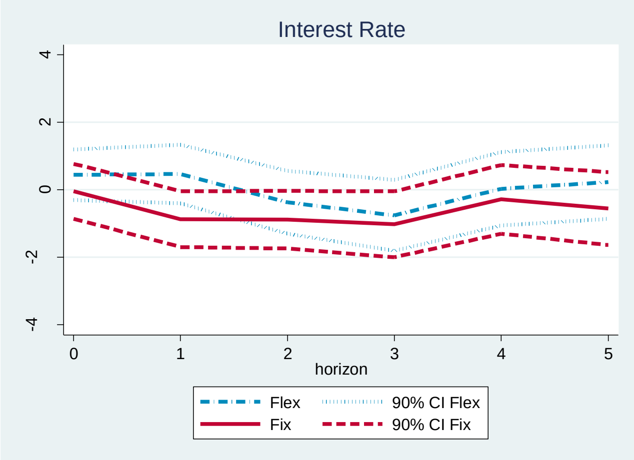


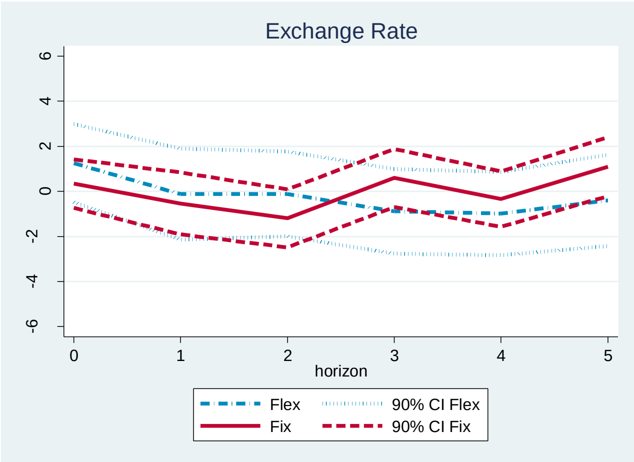

**Figure 3A: Cumulative Response of Output to Fiscal Shock**



## Figure 4: Response of Interest Rates to Fiscal Shock




## Figure 5: Response of Exchange Rates to Fiscal Shock




## Figure 6: Response of Net Exports to Fiscal Shock

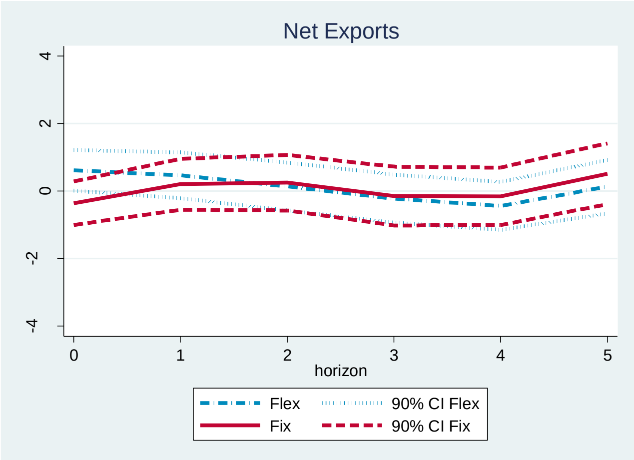


**Figure 7: Alternative Specification, Controlling for Financial Crisis**

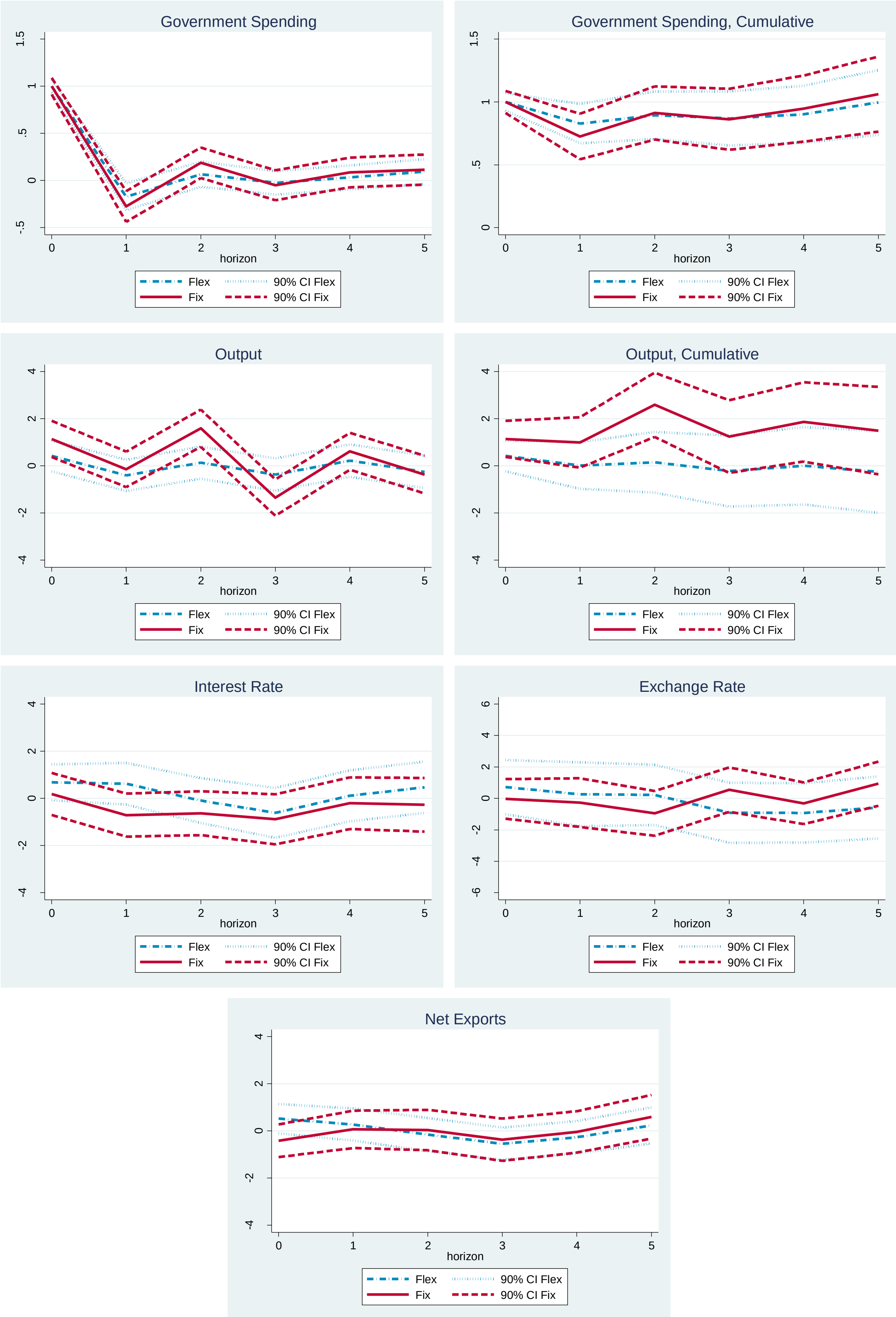


**Figure 8: Alternative Specification, Reduced Sample**

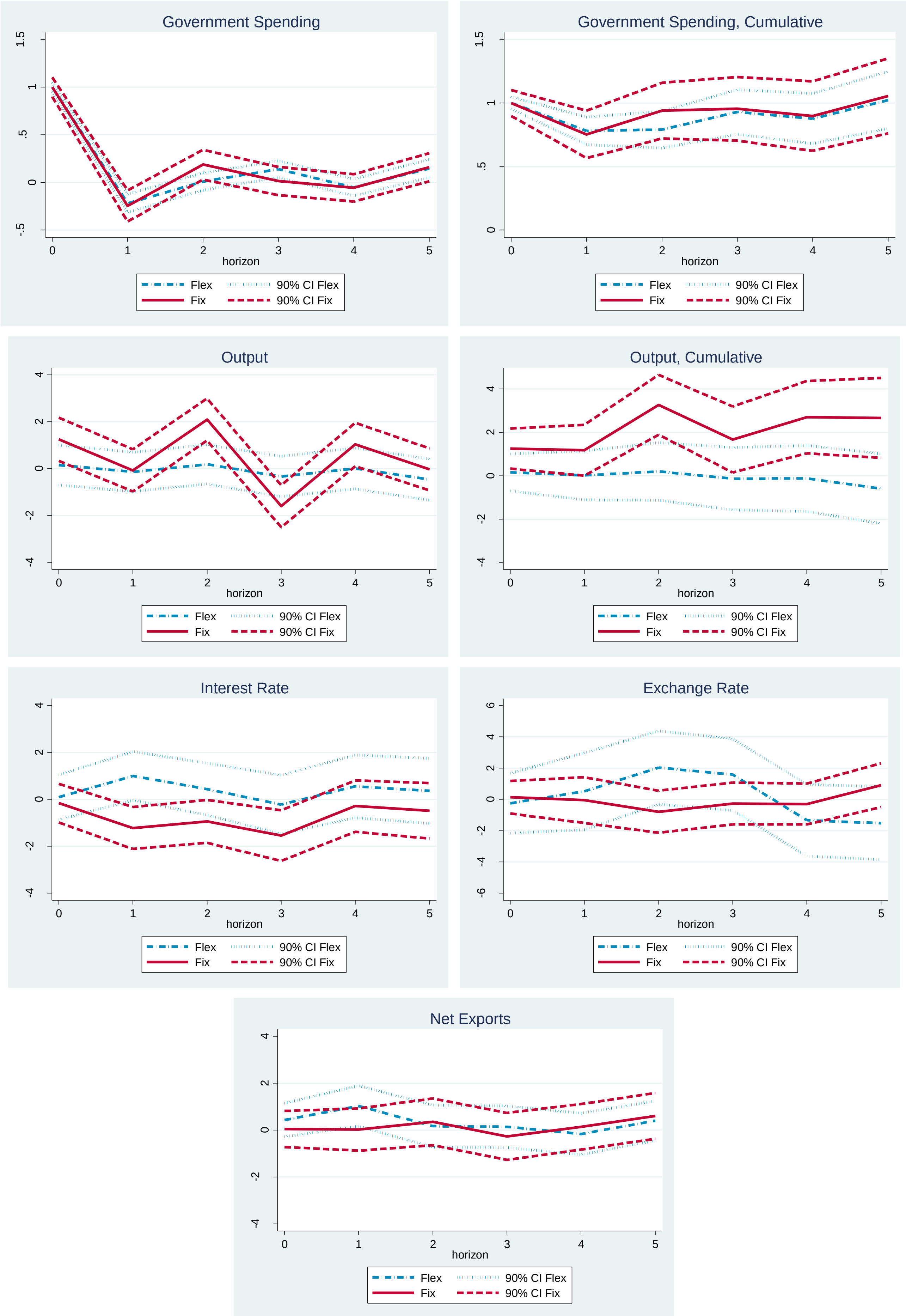


**Table 1: Summary Statistics**

| | Flexible Exchange Rate Regimes | | | | | | Fiexed Exchange Rate Regimes | | | | |
|---|---|---|---|---|---|---|---|---|---|---|---|
| | Obs | Mean | Std. Dev. | Min | Max | | Obs | Mean | Std. Dev. | Min | Max |
| *fe* | 495 | 0.05 | 0.22 | -0.90 | 1.01 | | 406 | 0.07 | 0.21 | -0.90 | 1.35 |
| *g* | 501 | 0.18 | 0.24 | -1.04 | 1.26 | | 425 | 0.19 | 0.27 | -1.68 | 1.76 |
| *y* | 501 | 1.30 | 1.52 | -7.43 | 7.59 | | 425 | 1.10 | 1.57 | -7.86 | 6.84 |
| *r* | 514 | 4.02 | 3.62 | -4.82 | 24.17 | | 437 | 2.89 | 2.57 | -3.63 | 14.66 |
| *ex* | 514 | -0.01 | 0.15 | -0.44 | 0.46 | | 438 | -0.04 | 0.12 | -0.70 | 0.34 |
| *nx* | 514 | 0.55 | 9.54 | -24.04 | 38.96 | | 438 | 0.63 | 13.53 | -40.63 | 57.09 |

*Note: Samples split by exchange rates regime classification. Source: OECD's Statistics and Projections database and BIS*

**Table 2: Stationarity Tests based on Augmented Dickey-Fuller Tests**

| | P | | Z | | L* | | Pm | |
|---|---|---|---|---|---|---|---|---|
| | Statistic | p-value | Statistic | p-value | Statistic | p-value | Statistic | p-value |
| *fe* | 100.25 | 0.00 | -3.97 | 0.00 | -4.02 | 0.00 | 4.73 | 0.00 |
| *g* | 88.40 | 0.00 | -3.35 | 0.00 | -3.33 | 0.00 | 3.57 | 0.00 |
| *y* | 88.25 | 0.00 | -3.81 | 0.00 | -3.67 | 0.00 | 3.55 | 0.00 |
| *r* | 80.85 | 0.01 | -2.41 | 0.01 | -2.45 | 0.01 | 2.83 | 0.00 |
| *ex* | 43.68 | 0.79 | 0.23 | 0.59 | 0.21 | 0.58 | -0.82 | 0.79 |
| *nx* | 38.68 | 0.91 | 1.65 | 0.95 | 1.74 | 0.96 | -1.31 | 0.90 |

*Note: Statistics and p-values based on ADF tests*

**Table 3: Impulse Responses to Spending Shock of 1% of Output**

| Variable | Horizon | Flexible | Fixed |
|---|---|---|---|
| Government Spending | 0 | 1 | 1 |
| | | (0.04) | (0.05) |
| | 1 | -0.19 | -0.29 |
| | | (0.08) | (0.09) |
| | 2 | 0.07 | 0.19 |
| | | (0.08) | (0.09) |
| | 3 | -0.07 | -0.08 |
| | | (0.07) | (0.09) |
| | 4 | 0.01 | 0.07 |
| | | (0.08) | (0.09) |
| | 5 | 0.07 | 0.1 |
| | | (0.08) | (0.1) |
| Output | 0 | 0.08 | 0.89 |
| | | (0.36) | (0.44) |
| | 1 | -0.3 | -0.08 |
| | | (0.37) | (0.44) |
| | 2 | 0.21 | 1.66 |
| | | (0.38) | (0.46) |
| | 3 | -0.23 | -1.26 |
| | | (0.39) | (0.46) |
| | 4 | 0.27 | 0.65 |
| | | (0.39) | (0.47) |
| | 5 | -0.13 | -0.27 |
| | | (0.41) | (0.48) |
| Inerest Rates | 0 | 0.44 | -0.05 |
| | | (0.46) | (0.49) |
| | 1 | 0.47 | -0.87 |
| | | (0.53) | (0.5) |
| | 2 | -0.37 | -0.89 |
| | | (0.57) | (0.52) |
| | 3 | -0.76 | -1.02 |
| | | (0.64) | (0.59) |
| | 4 | 0.03 | -0.29 |
| | | (0.66) | (0.62) |
| | 5 | 0.23 | -0.56 |
| | | (0.66) | (0.66) |
| Exchange Rates | 0 | 1.25 | 0.34 |
| | | (1.03) | (0.64) |
| | 1 | -0.12 | -0.53 |
| | | (1.2) | (0.82) |
| | 2 | -0.11 | -1.2 |
| | | (1.12) | (0.77) |
| | 3 | -0.89 | 0.59 |
| | | (1.12) | (0.77) |
| | 4 | -0.99 | -0.34 |
| | | (1.1) | (0.73) |
| | 5 | -0.4 | 1.09 |
| | | (1.21) | (0.78) |
| Net Exports | 0 | 0.61 | -0.36 |
| | | (0.36) | (0.38) |
| | 1 | 0.47 | 0.2 |
| | | (0.4) | (0.45) |
| | 2 | 0.14 | 0.25 |
| | | (0.42) | (0.49) |
| | 3 | -0.22 | -0.15 |
| | | (0.42) | (0.52) |
| | 4 | -0.44 | -0.16 |
| | | (0.42) | (0.51) |
| | 5 | 0.13 | 0.51 |
| | | (0.47) | (0.54) |

*Note: Horizon measured in semi-annual units, standard errors in parentheses*

**Table 4: Impulse Responses to Spending Shock, Alternative Specification, Controlling for Financial Crisis**

| Variable | Horizon | Flexible | Fixed |
|---|---|---|---|
| Government Spending | 0 | 1 | 1 |
| | | (0.04) | (0.05) |
| | 1 | -0.17 | -0.28 |
| | | (0.09) | (0.1) |
| | 2 | 0.07 | 0.19 |
| | | (0.08) | (0.1) |
| | 3 | -0.03 | -0.05 |
| | | (0.08) | (0.1) |
| | 4 | 0.03 | 0.09 |
| | | (0.08) | (0.1) |
| | 5 | 0.09 | 0.12 |
| | | (0.08) | (0.1) |
| Output | 0 | 0.42 | 1.14 |
| | | (0.4) | (0.47) |
| | 1 | -0.4 | -0.15 |
| | | (0.4) | (0.46) |
| | 2 | 0.14 | 1.59 |
| | | (0.41) | (0.48) |
| | 3 | -0.37 | -1.34 |
| | | (0.42) | (0.47) |
| | 4 | 0.22 | 0.62 |
| | | (0.42) | (0.48) |
| | 5 | -0.25 | -0.37 |
| | | (0.42) | (0.49) |
| Inerest Rates | 0 | 0.68 | 0.19 |
| | | (0.46) | (0.54) |
| | 1 | 0.62 | -0.71 |
| | | (0.54) | (0.55) |
| | 2 | -0.09 | -0.63 |
| | | (0.58) | (0.57) |
| | 3 | -0.61 | -0.89 |
| | | (0.64) | (0.65) |
| | 4 | 0.11 | -0.21 |
| | | (0.66) | (0.67) |
| | 5 | 0.47 | -0.27 |
| | | (0.66) | (0.69) |
| Exchange Rates | 0 | 0.71 | -0.04 |
| | | (1.03) | (0.75) |
| | 1 | 0.26 | -0.27 |
| | | (1.21) | (0.92) |
| | 2 | 0.22 | -0.96 |
| | | (1.14) | (0.85) |
| | 3 | -0.91 | 0.56 |
| | | (1.14) | (0.85) |
| | 4 | -0.93 | -0.31 |
| | | (1.12) | (0.79) |
| | 5 | -0.58 | 0.93 |
| | | (1.18) | (0.84) |
| Net Exports | 0 | 0.52 | -0.42 |
| | | (0.37) | (0.41) |
| | 1 | 0.27 | 0.07 |
| | | (0.41) | (0.47) |
| | 2 | -0.16 | 0.04 |
| | | (0.42) | (0.51) |
| | 3 | -0.54 | -0.37 |
| | | (0.41) | (0.53) |
| | 4 | -0.27 | -0.04 |
| | | (0.41) | (0.52) |
| | 5 | 0.23 | 0.6 |
| | | (0.46) | (0.55) |

*Note: Horizon measured in semi-annual units, standard errors in parentheses*

**Table 5: Impulse Responses to Spending Shock, Alternative Specification, Reduced Sample**

| Variable | Horizon | Flexible | Fixed |
|---|---|---|---|
| Government Spending | 0 | 1 | 1 |
| | | (0.03) | (0.06) |
| | 1 | -0.22 | -0.25 |
| | | (0.06) | (0.1) |
| | 2 | 0.01 | 0.19 |
| | | (0.06) | (0.09) |
| | 3 | 0.14 | 0.01 |
| | | (0.05) | (0.09) |
| | 4 | -0.05 | -0.06 |
| | | (0.05) | (0.09) |
| | 5 | 0.15 | 0.16 |
| | | (0.06) | (0.09) |
| Output | 0 | 0.15 | 1.25 |
| | | (0.52) | (0.56) |
| | 1 | -0.14 | -0.07 |
| | | (0.51) | (0.55) |
| | 2 | 0.18 | 2.09 |
| | | (0.51) | (0.55) |
| | 3 | -0.33 | -1.6 |
| | | (0.52) | (0.54) |
| | 4 | 0.01 | 1.03 |
| | | (0.53) | (0.56) |
| | 5 | -0.47 | -0.03 |
| | | (0.53) | (0.54) |
| Inerest Rates | 0 | 0.09 | -0.16 |
| | | (0.58) | (0.5) |
| | 1 | 1 | -1.22 |
| | | (0.63) | (0.54) |
| | 2 | 0.44 | -0.94 |
| | | (0.67) | (0.56) |
| | 3 | -0.22 | -1.54 |
| | | (0.76) | (0.66) |
| | 4 | 0.56 | -0.29 |
| | | (0.81) | (0.67) |
| | 5 | 0.36 | -0.49 |
| | | (0.84) | (0.72) |
| Exchange Rates | 0 | -0.25 | 0.13 |
| | | (1.14) | (0.62) |
| | 1 | 0.52 | -0.05 |
| | | (1.47) | (0.87) |
| | 2 | 2.03 | -0.79 |
| | | (1.4) | (0.8) |
| | 3 | 1.58 | -0.27 |
| | | (1.37) | (0.8) |
| | 4 | -1.33 | -0.3 |
| | | (1.37) | (0.78) |
| | 5 | -1.53 | 0.91 |
| | | (1.39) | (0.84) |
| Net Exports | 0 | 0.43 | 0.05 |
| | | (0.43) | (0.46) |
| | 1 | 1.03 | 0.02 |
| | | (0.52) | (0.54) |
| | 2 | 0.17 | 0.35 |
| | | (0.54) | (0.59) |
| | 3 | 0.14 | -0.27 |
| | | (0.53) | (0.6) |
| | 4 | -0.16 | 0.14 |
| | | (0.53) | (0.58) |
| | 5 | 0.4 | 0.61 |
| | | (0.5) | (0.58) |

*Note: Horizon measured in semi-annual units, standard errors in parentheses*