Commodity Prices and Fiscal (Pro)Cyclicality*

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Abstract

Consensus holds that Emerging Markets and Developing Economies (EMDEs) engage in procyclical fiscal behavior. We emphasize that considering conditional responses to macroeconomic shocks is crucial when evaluating fiscal cyclicality, as neglecting this can result in significant biases. This study investigates the effects of exogenous commodity price shocks on fiscal variables in EMDEs by exploiting major narrative episodes and the heterogeneous exposure of countries to these shocks. Our results reveal that, following an expansionary shift in the terms of trade, fiscal authorities raise government spending and moderately increase taxes. The overall fiscal stance mitigates the effect of commodity price booms while leading to an improvement in the primary balance. These findings contrast with conventional wisdom but align with the optimal policy response to export price shocks predicted by a multi-good small open economy model with incomplete financial markets. We also highlight the role of institutional quality in shaping fiscal policy responses.

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1 Introduction

The cyclicality of fiscal policy in Emerging Markets and Developing Economies (EMDEs) has long been a subject of intense debate in macroeconomics. Conventional wisdom, supported by extensive empirical evidence, suggests that fiscal policy in these economies is predominantly procyclical: Governments increase spending and cut taxes during expansions, exacerbating macroeconomic volatility (Kaminsky, Reinhart, and Végh, 2004).[1] This pattern stands in sharp contrast to standard policy recommendations that advocate for a countercyclical approach, reducing expenditures and increasing taxes during economic expansions to create fiscal space to stabilize the economy during downturns. In fact, fiscal procyclicality is widely regarded as a key stylized fact that any theory of fiscal policy in EMDEs must address. Various explanations have been proposed, including institutional weaknesses, challenges in maintaining sustainable fiscal management, and the impact of financial frictions under high sovereign risk.

In this paper, we challenge the prevailing consensus on the procyclicality of fiscal policy in EMDEs. We demonstrate that the standard empirical approach–relying on unconditional correlations between fiscal instruments and economic activity–can fundamentally misrepresent fiscal cyclicality. *Unconditional* correlations capture fiscal policy responses to various economic shocks but are also biased by reverse causality in the presence of fiscal shocks. Standard approaches to correct for this bias can also be problematic, as the resulting estimates do not necessarily represent a weighted average of the underlying conditional correlations. Instead, we argue that *conditional* correlations more accurately capture whether fiscal policy mitigates or amplifies the economy's response to shocks, making them a more appropriate basis for assessing fiscal cyclicality. To apply this approach, we focus on commodity price shocks, which are a key driver of macroeconomic fluctuations in EMDEs, and therefore provide a natural setting to analyze fiscal policy responses to external disturbances.

Measuring *conditional fiscal cyclicality* requires a credible identification framework to disentangle the transmission of a specific shock to the economy and the corresponding fiscal response. In our setting, this hinges on distinguishing exogenous movements in commodity prices to avoid confounding commodity shocks with changes in commodity prices driven by shifts in global economic activity or financial conditions. Indeed, other global shocks affecting commodity prices impact EMDEs' aggregate activity and fiscal policy through channels that may differ from those that characterize the transmission of commodity shocks. To identify the dynamic transmission of commodity shocks, we exploit the heterogeneous exposure of countries to changes in global commodity prices arising from geopolitical events, weather shocks, and natural disasters.

The fiscal response to commodity price shocks contrasts sharply with what unconditional correlations suggest. Unconditional correlations indicate a procyclical stance, with government spending rising alongside export prices and tax rates declining, resulting in a

[1] The stark contrast in the cyclicality of fiscal policy between advanced and developing countries has been extensively documented by, among others, Gavin and Perotti (1997); Ilzetzki and Végh (2008); Kaminsky (2010); Talvi and Vegh (2005); Végh and Vuletin (2015).

weak link between the primary balance and commodity price fluctuations. This aligns with the broader literature, which portrays fiscal policy in EMDEs as amplifying external shocks rather than mitigating them. However, when focusing on commodity price booms driven by commodity-specific shocks, a different picture emerges: While spending remains procyclical, tax policies are markedly countercyclical. As a result, despite higher spending, the primary balance improves, enabling EMDEs to strengthen their fiscal position during booms.

Through a counterfactual analysis, we show that although increased spending weakens the primary balance and amplifies economic fluctuations, the rise in tax rates more than compensates for this effect. Coupled with a broader tax base, this leads to an improved fiscal balance. More importantly, the overall fiscal response dampens the impact of commodity price shocks on domestic GDP, challenging the notion that fiscal policy in EMDEs is inherently destabilizing. These findings fundamentally challenge the prevailing consensus on procyclical fiscal policy in EMDEs. They also carry significant implications: Assessing policy effectiveness without accounting for the conditional nature of fiscal responses may lead to misguided recommendations. Furthermore, since theories of fiscal policy in EMDEs often assume procyclicality as a stylized fact, our results call this assumption into question.

The seminal work of Barro (1979) emphasizes tax smoothing and countercyclical debt as fundamental components of prudent fiscal policy–a perspective that has since gained widespread acceptance (see, e.g., Yared, 2019). However, these insights are typically derived from models focusing on domestic shocks (often productivity shocks) and do not adequately account for the role of the export sector and its susceptibility to globally driven commodity price fluctuations, which are crucial factors for many EMDEs. Our evidence shows that EMDEs' fiscal response to commodity price shocks, involves an increase in taxes with an improvement in the primary balance, despite the increase in spending. How does this align with the recommended optimal fiscal policy response? To address this question, we develop a small open economy (SOE) multi-good model encompassing importable, exportable, and non-tradable (MXN) goods set against the backdrop of incomplete financial markets. We consider a framework where government spending yields benefits but requires financing through a consumption tax, which induces a wedge between the (sectoral) marginal product of labor and the marginal rate of substitution between consumption and leisure.

In this setting, the optimal response to shifts in commodity prices and the country's terms of trade depends on balancing the benefits against the costs of reallocating resources over time. Following an exogenous increase in export prices, the optimal fiscal strategy involves increasing government spending at a rate lower than that of output, in line with prudent fiscal management. Since the government levies taxes on consumption, higher consumption expands the tax base and improves fiscal revenues. However, the revenue boost from the expansion in aggregate demand alone is insufficient to meet future resource allocation needs. Therefore, the optimal policy requires further increasing fiscal revenues by raising the tax rate despite the distortion it introduces. This decision is shaped by the interaction between tax policy and financial market frictions, which influence interest rates on private and public debt and, consequently, the intertemporal cost of consumption. Overall, this strategy reallocates the windfall from the commodity price boom towards the future. Therefore, optimal policy prescribes (conditional) procyclical government spending and countercyclical tax policy. Our analysis indicates that the optimal policy response to an export price shock aligns with our empirical findings.

We also use the model setting to demonstrate that the optimal policy can generate markedly different patterns in conditional and unconditional correlations, even within a controlled environment featuring a well-specified optimal policy. Consequently, unconditional correlations offer unreliable guidance on the appropriateness of the fiscal stance, even within our theoretical framework. Moreover, we show that the outcomes of the Ramsey optimal policy can be closely replicated by simple, implementable fiscal rules on spending and revenues, where the fiscal authority responds countercyclically and smooths debt over time.

Finally, we uncover significant heterogeneity linked to institutional quality. EMDEs with strong institutions pursue countercyclical fiscal policy during commodity booms, while those with weak institutions adopt a procyclical stance. The results for countries featuring high institutional quality align with our baseline findings. By contrast, in countries with lower institutional quality, both spending and spending as a proportion of GDP increase, and taxes are not raised, leading to a more muted improvement in revenues and the primary balance. Although political economy factors may help explain divergent fiscal responses, we show that these outcomes are also consistent with optimal policy in an environment characterized by financial market frictions and production inefficiencies. In countries with weaker institutions, economic rents and borrowing constraints increase the costs of shifting resources over time, reducing the benefits of windfalls. This framework helps account for the distinct fiscal responses observed in economies with lower institutional quality.

Related Literature.— This paper contributes to the extensive empirical literature on the cyclical behavior of fiscal policy. The existing evidence consistently shows that fiscal procyclicality is a defining feature of EMDEs (see, Frankel, Végh, and Vuletin, 2013; Gavin and Perotti, 1997; Ilzetzki and Végh, 2008; Kaminsky, Reinhart, and Végh, 2004; Végh and Vuletin, 2015, among others).[2] Our key contribution lies in emphasizing the critical distinction between the conditional and unconditional cyclicality of fiscal policy, both empirically and theoretically. Our evidence challenges the conventional assumption of procyclical fiscal policy in EMDEs, underscoring the need to consider heterogeneity in fiscal responses to different macroeconomic shocks. We also demonstrate that fiscal responses to commodity price shocks vary significantly across countries, depending on institutional strength–complementing and extending the findings of Arezki and Brückner (2012) and Céspedes and Velasco (2014).

Commodity price shocks are a major source of external vulnerability for EMDEs and play a central role in shaping their business cycles (see, e.g., Fernández, González, and Rodríguez, 2018; Fernández, Schmitt-Grohé, and Uribe, 2017). Consequently, understanding fiscal responses to commodity booms has received significant attention (see, e.g., Arroyo Marioli and Végh, 2023; Céspedes and Velasco, 2014; Kaminsky, 2010). A widely recommended pol-

[2] Frankel, Végh, and Vuletin (2013) note some progress in mitigating this tendency over the past two decades.

icy approach is to adopt countercyclical fiscal strategies, building buffers during booms to cushion downturns. Assessing fiscal responses to commodity price shocks is therefore essential when formulating policy recommendations for EMDEs (see, e.g., IMF, 2016; World Bank, 2024). This is particularly important for developing economies, where exposure to macroeconomic volatility and external shocks underscores the need for effective stabilization policies. Our analysis offers new evidence on fiscal policy responses to commodity price booms. More importantly, our findings highlight the need to evaluate fiscal responses within the context of specific shocks to develop effective policy recommendations.

Several factors have been proposed to explain fiscal policy responses in developing countries, including weak political institutions (see, e.g., Alesina et al., 2008; Ilzetzki, 2011; Lane and Tornell, 1999), challenges in committing to sustainable fiscal management over time (see, e.g., Chari and Kehoe, 1993; Halac and Yared, 2024), incomplete markets, borrowing constraints, and sovereign default (see, e.g., Azzimonti and Mitra, 2023; Bianchi et al., 2023; Cuadra et al., 2010). While our analysis abstracts from many relevant features of EMDEs, it underscores the importance of recognizing the heterogeneity in responses when considering different types of economic shocks, especially when examining optimal fiscal policy.

In line with Fernández et al. (2021) and Riascos and Végh (2003) we emphasize the role of financial market incompleteness–a salient feature of developing countries' economies. However, while their focus is on productivity shocks–as is much of the previous work in these settings–we investigate these issues within a multisectoral environment where traded prices are a source of external vulnerability for the domestic economy. In particular, our theoretical framework builds on Mendoza (1995) and Schmitt-Grohé and Uribe (2018), using a simplified MXN framework, and incorporating independent shocks to export prices as in Di Pace, Juvenal, and Petrella (2024). Within this framework, we introduce a financial channel through which commodity shocks impact government financing costs by altering sovereign risk (as in Drechsel, McLeay, and Tenreyro, 2019; Drechsel and Tenreyro, 2018; Hamann, Mendez-Vizcaino, Mendoza, and Restrepo-Echavarria, 2023). While Drechsel et al. (2019) examine this channel's impact on EMDEs' monetary policy, we analyze its influence on EMDEs' fiscal space.

Outline.— The paper is organized as follows. Section 2 presents the data and preliminary evidence. Section 3 outlines the framework for analyzing the cyclicality of fiscal instruments. The identification strategy is described in Section 4, followed by the empirical analysis and baseline results, including extensions and robustness checks, in Section 5. Section 6 uses a theoretical model to examine the optimal fiscal policy response. Section 7 concludes.

2 Data

The estimation period runs from 1990 to 2019. The yearly dataset covers 54 emerging and developing countries. Within this category, 31 belong to the upper middle income group, 14 to the lower middle income group, and 9 to the low income group. The sample of countries covers all the regions in the world. The dataset includes information on fiscal expenditures, fiscal revenues, primary balance, government interest expenditures, value added tax (VAT) rates, output, EMBI spreads, a measure of institutional quality, and export prices. The selection of countries is dictated by data availability, considering that EMBI spreads are only available from the 1990s.[3]

Fiscal expenditures, fiscal revenues, and primary balance data are obtained from the IMF World Economic Outlook (WEO). Starting in 2010, the WEO switched to using the Government Finance Statistics Manual (GFSM) 2001 framework, replacing the previous GFSM 1986 standard. While for numerous countries, data under the new methodology are available retroactively, the coverage can be expanded by integrating it with the preceding version. Therefore, we use the updated framework as a baseline and splice the data backwards with the preceding version. VAT rates are mainly sourced from Végh and Vuletin (2015) and extended using information from the Inter-American Center of Tax Administrations. We focus on the VAT rate as a key fiscal instrument for several reasons. First, VAT represents one of the largest revenue source in EMDEs, accounting for approximately 30% of total revenue in 2019 across our sample countries.[4] This contrasts with other forms of taxation, such as personal income taxes, which constitute only 12% of total revenue and are often less effective in economies with large informal sectors.[5] Second, VAT rate data offer the most comprehensive coverage in our dataset. Finally, as demonstrated by Végh and Vuletin (2015), tax rate changes tend to exhibit strong comovement, suggesting that the VAT rate serves as a reliable proxy for broader tax policy shifts.

Country-specific nominal GDP is sourced from the World Bank's World Development Indicators (WDI) database. Emerging market sovereign spreads are measured as spreads over Treasuries of the J.P. Morgan EMBI global diversified index obtained from Datastream, Bloomberg, and J.P Morgan. Institutional quality is measured using the political risk rating from the International Country Risk Guide (ICRG). This metric ranges from 0 (highest risk) to 100 (least risk) such that a higher value implies a country has a higher quality of institutions.

Export price indices denominated in U.S. dollars are obtained by extending the sample in Di Pace et al. (2024), constructed following the IMF Export and Import Price Manual (IMF, 2009). Specifically, export prices are a weighted average of commodity and manufacturing prices where the country-specific export shares give the weights. Commodity prices are from the World Bank's Commodity Price Data, manufacturing prices from the Federal Reserve Bank of St. Louis FRED, and export shares from the MIT Observatory of Economic Complexity.

[3] Data sources and coverage specifics can be found in Appendix A.

[4] This figure, sourced from the OECD's Global Revenue Statistics dataset, reflects the average across 40 countries in our sample.

[5] This is also sourced from the OECD's Global Revenue Statistics dataset, and covers 33 countries in our sample in 2019.

Table 1: Unconditional Correlations

	Corr w/ GDP		Corr w/ Px	
Spending	**0.16**	[0.05]	**0.23**	[0.04]
VAT Rate	**-0.10**	[0.03]	**-0.08**	[0.04]
Primary Balance	0.03	[0.04]	0.06	[0.04]
Revenue	**0.17**	[0.05]	**0.30**	[0.04]
Spending/GDP	**0.10**	[0.04]	**0.16**	[0.04]
Revenue/GDP	**0.13**	[0.04]	**0.24**	[0.03]
EMBI Spread	**-0.10**	[0.05]	-0.05	[0.05]

Notes: For each variable listed, we report the average (unconditional) correlation and its associated standard error (in square brackets). The left column measures correlation with respect to detrended GDP, whereas the right column measures correlation with respect to detrended export prices. Values in bold denote that the average correlation is significant at a 10% level.

2.1 Unconditional Evidence

Building on the seminal work of Kaminsky et al. (2004), a large body of research has investigated the pro or countercyclicality of fiscal policy by focusing on the unconditional correlation between key fiscal instruments–namely government spending and tax rates–and GDP. A central insight from this literature is the need to distinguish between fiscal instruments, which lie under direct policymaker control, and fiscal outcomes, which are strongly influenced by endogenous macroeconomic dynamics and other non-policy factors (Kaminsky et al., 2004; Végh and Vuletin, 2015). This distinction clarifies the extent to which observed correlations reflect intentional policy actions rather than indirect responses to economic fluctuations. Among these outcomes, the primary balance, defined as total revenues minus non-interest expenditures, is often highlighted as a summary indicator of the "fiscal stance."[6]

Investigating the behavior of fiscal policy over commodity price cycles, Céspedes and Velasco (2014) and Kaminsky (2010) examine how fiscal instruments correlate with a commodity-based measure of export prices. Following this approach, Table 1 presents the average unconditional correlations between fiscal instruments (in the upper portion of the table) and GDP or export prices, and then, in the lower portion, between fiscal outcomes and these same two variables.[7]

Turning to the patterns in Table 1, government spending and the VAT rate both exhibit procyclicality on average, consistent with findings in earlier literature (see, e.g., Arroyo Marioli and Végh, 2023; Céspedes and Velasco, 2014; Frankel et al., 2013; Végh and Vuletin, 2015). Meanwhile, revenues show a clear positive correlation with GDP, largely reflecting the mechanical rise in taxable income during expansions. By contrast, the primary balance displays a near-zero correlation, suggesting that the tendency for both spending and revenues to move with GDP yields minimal net effect on the balance itself–a pattern consistent with the ten-

[6]For a detailed discussion of the primary balance as a measure of fiscal stance, see, e.g., IMF (2024).

[7]While commodity and export prices are typically treated as exogenous to the domestic economy, these unconditional correlations nonetheless reflect the influence of all global shocks, not just those originating in commodity markets (Juvenal and Petrella, 2024).

dency of some EMDEs to follow a balanced budget rule (as explained in Ilzetzki and Végh, 2008). Similar results emerge when considering the correlations with export prices. The unconditional correlations show some heterogeneity across countries and institutional quality (these results are shown in Appendix B).

At first glance, these correlations may suggest a deviation from fiscal best practices that prescribe countercyclical behavior. However, as we show in the next section, unconditional correlations can mask whether fiscal policy is truly pro or countercyclical once different shocks are disentangled.

3 Fiscal Cyclicality: Unconditional vs. Conditional Evidence

Kaminsky et al. (2004) define countercyclical fiscal policy as actions aimed at stabilizing the business cycle by mitigating the transmission of shocks, whereas procyclical policy amplifies business cycle fluctuations. Crucially, their framework focuses on the behavior of policy instruments–in our setting, government spending and tax rates–rather than outcomes like revenues or the primary balance, which naturally respond to economic conditions even absent changes in discretionary policy. In practice, fiscal policy is classified as procyclical if spending (tax rates) increases (decreases) during expansions and decreases (increases) during downturns; it is countercyclical if the opposite pattern holds. A positive (negative) comovement between GDP and spending (tax rates) thus signals that fiscal policy amplifies rather than dampens the business cycle (i.e., "when it rains, it pours").

In what follows, we argue that *unconditional* correlations between fiscal instruments and aggregate demand can be misleading when identifying whether fiscal policy amplifies or stabilizes the cycle. By contrast, examining *conditional* comovement isolates how policy instruments respond to particular shocks, offering a more accurate account of whether and how fiscal interventions mitigate or magnify the transmission of disturbances to aggregate demand.

3.1 Illustrative example

Consider the following simplified framework capturing the dynamics of aggregate demand, $\mathtt{y}_t$, and a fiscal policy instrument, $\mathtt{f}_t$:

$$\begin{aligned} \mathtt{y}_t &= a^y_f \mathtt{f}_t + a^y_{px} \mathtt{p}^x_t + e^y_t, \\ \mathtt{f}_t &= b^f_y \mathtt{y}_t + b^f_{px} \mathtt{p}^x_t + e^f_t, \end{aligned} \tag{1}$$

where $\mathtt{p}^x_t = e^x_t$, represents export price shocks. Focusing on government spending as the fiscal instrument, we assume $a^y_f > 0$ and $a^y_{px} > 0$, indicating that both government spending and export prices positively affect aggregate demand. The term e^y_t captures exogenous "output" shocks.[8] The fiscal authority sets policy according to a simple rule that responds to

[8]In the case of taxes, $a^y_f < 0$ would typically apply. The analysis in this section carries through seamlessly under this alternative scenario, with policy response coefficients adjusting signs accordingly.

changes in aggregate demand and export prices. Deviations from this rule, represented by e_t^f, correspond to fiscal policy shocks. To ensure aggregate demand responds positively or remains unaffected by shocks e_t^y and e_t^x, we impose meaningful restrictions on policy reaction coefficients.[9] All shocks are *iid* and uncorrelated between each other.

By combining the system's equations, we can characterize the response of aggregate demand to shocks in the economy:

$$\mathtt{y}_t = \frac{1}{1 - a_f^y b_y^f}[(a_f^y b_{px}^f + a_{px}^y)e_t^x + a_f^y e_t^f + e_t^y]. \tag{2}$$

We can therefore formally determine whether fiscal responses amplify or mitigate the impact of shocks on the economy. Focusing on the simplest case where the sign of the policy coefficients is the same, i.e., spending increases or decreases with output after a positive shift in e_t^y or e_t^x. In this setting, fiscal policy *amplifies* the impact of shocks (is procyclical) if both $b_y^f > 0$ and $b_{px}^f > 0$ (i.e. $\text{Var}(\mathtt{y}_t|b_y^f > 0 \;\&\; b_{px}^f > 0) > \text{Var}(\mathtt{y}_t|b_y^f = b_{px}^f = 0)$). Conversely, fiscal policy *mitigates* the effects of shocks (i.e., is countercyclical) if both $b_y^f < 0$ and $b_{px}^f < 0$, implying that $\text{Var}(\mathtt{y}_t|b_y^f < 0 \;\&\; b_{px}^f < 0) < \text{Var}(\mathtt{y}_t|b_y^f = b_{px}^f = 0)$.

Looking at the unconditional covariance between the fiscal instrument and output, we can express it as:

$$\text{Cov}(\mathtt{y}_t, \mathtt{f}_t) = \underbrace{\frac{[b_y^f(a_f^y b_{px}^f + a_{px}^y)^2 + b_{px}^f(a_f^y b_{px}^f + a_{px}^y)]}{(1 - a_f^y b_y^f)^2}\sigma_x^2}_{=\text{Cov}(\mathtt{y}_t, \mathtt{f}_t|e_t^x)} + \underbrace{\frac{b_y^f}{(1 - a_f^y b_y^f)^2}\sigma_y^2}_{=\text{Cov}(\mathtt{y}_t, \mathtt{f}_t|e_t^y)} + \underbrace{\frac{a_f^y}{(1 - a_f^y b_y^f)^2}\sigma_f^2}_{=\text{Cov}(\mathtt{y}_t, \mathtt{f}_t|e_t^f)}. \tag{3}$$

$\text{Cov}(\mathtt{y}_t, \mathtt{f}_t) > 0$ would typically be interpreted as an indication that fiscal policy is procyclical. However, such positive co-movement can arise even when both $b_y^f < 0$ and $b_{px}^f < 0$, when fiscal response mitigates the impact of both e_t^y and e_t^x. This occurs, for instance, when the policy responses b_y^f and b_{px}^f are not strong enough and/or a_f^y and σ_f^2 are large.[10] This case may be particularly relevant for developing countries, which are prone to large discretionary fiscal policies, for example, during election periods (i.e., σ_f^2 is large), and the endogenous countercyclical reaction of fiscal policy may be modest (i.e., b_y^f and b_{px}^f are small).

The co-movement of spending and output depends on the conditional covariance with respect to each shock. Fiscal shocks add a positive component to the unconditional covariance, $(\text{Cov}(\mathtt{y}_t, \mathtt{f}_t|e_t^f) > 0$), which is independent of the policy response coefficients. As a result, unconditional covariances or correlations are unreliable measures for determining whether fiscal policy mitigates or amplifies the effects of shocks on the economy.

More generally, it is possible to expand this example to account for multiple shocks and different channels of transmission (consider, for instance, the case of a U.S. monetary policy shock or a global financial shock). These shocks can affect output (and potentially enter the fiscal rule) both through their impact on the export prices (i.e., by influencing global commodity prices) and through other channels, such as altering financing costs or sovereign risk.

[9] It is straightforward to show that this implies $b_y^f \leq 1/a_f^y$ and $b_{px}^f \geq -a_{px}^y/a_f^y$.

[10] Formally, this occurs when $[b_y^f(a_f^y b_{px}^f + a_{px}^y)^2 + b_{px}^f(a_f^y b_{px}^f + a_{px}^y)]\sigma_x^2 + b_y^f\sigma_y^2 > -a_f^y\sigma_f^2$.

In such a scenario, it is possible for fiscal policy to mitigate the effects of export price shocks while simultaneously amplifying the transmission of other shocks. Consequently, unconditional correlations are a poor proxy for assessing fiscal cyclicality. In contrast, conditional correlations, which focus on the role of fiscal policy in the transmission of specific shocks, remain consistently informative. They provide a clear understanding of whether fiscal interventions mitigate or amplify the impact of a particular shock on aggregate demand.

3.2 Measurement

Let us now turn to the issue of measurement of (pro)cyclicality. Let us consider a simplified setting with a single policy instrument, $\mathtt{f}_t$, and a business cycle indicator, $\mathtt{y}_t$, both in deviation from steady state. Fiscal cyclicality is typically measured by the sign of the correlation between the two variables or, equivalently, the sign of the regression coefficient in

$$\mathtt{f}_t = \gamma \mathtt{y}_t + \eta_t. \tag{4}$$

To simplify the derivations, assume $\mathtt{f}_t$ and $\mathtt{y}_t$ are iid and driven by three shocks:

$$\begin{aligned} \mathtt{f}_t &= \alpha_1 \upsilon_{1,t} + \alpha_2 \upsilon_{2,t} + \alpha_3 \varepsilon_{p,t}, && (5) \\ \mathtt{y}_t &= \beta_1 \upsilon_{1,t} + \beta_2 \upsilon_{2,t} + \beta_3 \varepsilon_{p,t}, && (6) \end{aligned}$$

where $\varepsilon_{p,t} \sim iid\mathcal{N}\left(0, \sigma_p^2\right)$ denotes the policy shock, and $\upsilon_{j,t} \sim iid\mathcal{N}\left(0, \sigma_j^2\right)$ for $j = \{1, 2\}$ are two non-policy shocks.[11] In this setting, it is clear that the OLS coefficient in equation (4) yields

$$\widehat{\gamma}_{ols} = \frac{\alpha_1 \beta_1 \sigma_1^2 + \alpha_2 \beta_2 \sigma_2^2 + \alpha_3 \beta_3 \sigma_p^2}{\beta_1^2 \sigma_1^2 + \beta_2^2 \sigma_2^2 + \beta_3^2 \sigma_p^2}. \tag{7}$$

Therefore, the sign of this coefficient depends on the response of the two variables to each of the shocks $\{\alpha_j, \beta_j\}_{j=1}^3$, as well as the relative importance of each shock.

Similar to the argument presented in equation (3), the OLS coefficient reflects the sum of two components: The conditional covariances associated with non-policy shocks, which are directly influenced by the fiscal authority's behavior, and the conditional covariance with policy shocks, which remains independent of policy response. The latter introduces a bias, as it can distort the interpretation of the covariance's sign when used as a proxy for assessing the procyclicality or countercyclicality of fiscal policy.

The literature has long recognized the need to account for the potential endogeneity in these regressions (Rigobon, 2004) and, in particular, to the endogeneity associated with the policy shock which mechanically biases the results towards finding procyclicality in fiscal policy (see, e.g., Ilzetzki and Végh, 2008).[12] To address this, it is common to use an instru-

[11] These simplifying assumptions do not limit the generality of the argument, which only requires the presence of more than one non-policy shock affecting the variables. The example presented in Section 3.1 follows this structure.

[12] Consider the case of government spending: if it increases demand, then $\alpha_3\beta_3 > 0$, introducing a positive bias to $\widehat{\gamma}_{ols}$. Conversely, for taxes or revenues, if higher taxes reduce demand, $\alpha_3\beta_3 < 0$, resulting in a negative bias to $\widehat{\gamma}_{ols}$.

ment, $\mathtt{z}_t$, orthogonal to the policy shock:

$$\mathtt{z}_t = \kappa_1 \upsilon_{1,t} + \kappa_2 \upsilon_{2,t} + u_t, \tag{8}$$

where $u_t \perp \varepsilon_{p,t}$. The typical choices for these types of instruments are foreign variables, such as the growth rate of the major trading partners (Panizza and Jaimovich, 2007), the real returns on U.S. Treasury bills, or changes in the price of exports (Ilzetzki and Végh, 2008).

In this context, the instrumental variable coefficient can be written as:

$$\widehat{\gamma}_{iv} = \frac{Cov\,(\mathtt{f},\mathtt{z})}{Cov\,(\mathtt{y},\mathtt{z})} = \frac{\kappa_1\alpha_1\sigma_1^2 + \kappa_2\alpha_2\sigma_2^2}{\kappa_1\beta_1\sigma_1^2 + \kappa_2\beta_2\sigma_2^2} = \frac{\alpha_1}{\beta_1}\varkappa + \frac{\alpha_2}{\beta_2}\,(1-\varkappa)\,, \tag{9}$$

where $\varkappa = \frac{\kappa_1\beta_1\sigma_1^2}{\kappa_1\beta_1\sigma_1^2+\kappa_2\beta_2\sigma_2^2}$. This shows that the regression coefficient is an affine combination of the conditional policy responses $\frac{df_t}{dy_t}|\upsilon_{j,t} = \frac{\alpha_j}{\beta_j}$. The term $\varkappa$ or $(1-\varkappa)$ can be negative unless $\mathtt{sign}\,(\kappa_1\beta_1) = \mathtt{sign}\,(\kappa_2\beta_2)$, in which case $\varkappa \in (0,1)$.

Hence, the sign of $\widehat{\gamma}_{iv}$ does not necessarily align with the (conditional) policy responses.[13] For instance, taking export prices as an instrument, even when totally independent from domestic conditions, as long as they are driven by multiple shocks, and those shocks have a distinct transmission to the domestic economy (i.e., $\frac{\alpha_1}{\beta_1} \neq \frac{\alpha_2}{\beta_2}$), the IV estimate does not provide any guidance of policy pro or counter-cyclicality of fiscal policy.[14]

Reverse causality is the sole concern when assessing the cyclicality of the policy instrument only if the instrument responds to the shocks only because the policymaker directly reacts to observable movements in the business cycle indicator. In this case, the policy variable can be represented by a simple rule: $\mathtt{f}_t = \chi\mathtt{y}_t + \varepsilon_{p,t}$.

Given the DGP for $\mathtt{y}_t$, equation (6), it follows that the DGP for the policy variable can be represented as a particular restriction of equation (5), with $\alpha_1 = \chi\beta_1$ and $\alpha_2 = \chi\beta_2$, hence, $\chi = \frac{\alpha_1}{\beta_1} = \frac{\alpha_2}{\beta_2}$. Therefore, the IV estimator in equation (9) delivers $\widehat{\gamma}_{iv} = \chi = \frac{\alpha_1}{\beta_1} = \frac{\alpha_2}{\beta_2}$. Beyond this restrictive case, for example, as long as the policy reaction function incorporates a response to an indicator other than aggregate demand, standard strategies to address the endogeneity of the business cycle's response to policy do not offer effective guidance on the cyclicality of the policy response.[15]

These derivations highlight that standard approaches to addressing endogeneity may not provide clear guidance on the cyclicality of fiscal policy responses. In Section 3.1, we argued that conditional cyclicality–defined as the sign of the correlation between the fiscal in-

[13] This is also true when either the fiscal instrument or the business cycle indicator are affected by only one shock, e.g., $\alpha_2 = 0$ or $\beta_2 = 0$, in which case the sign of the IV estimator is still a function of the response coefficients as well as the sensitivity of the shock to the two instruments. In fact, for $\beta_2 = 0$: $\widehat{\gamma}_{iv} = \frac{\alpha_1}{\beta_1} + \frac{\kappa_2\alpha_2\sigma_2^2}{\kappa_1\beta_1\sigma_1^2}$ and for $\alpha_2 = 0$: $\widehat{\gamma}_{iv} = \frac{\alpha_1}{\beta_1}\varkappa$ (where $\varkappa$ can be positive or negative).

[14] It is important to note that the same reasoning applies to studies that infer the procyclicality or countercyclicality of fiscal policy by examining the correlations between fiscal instruments and export prices directly. This approach is essentially analogous to a reduced-form regression in an instrumental variable (IV) framework, where export prices serve as instruments for aggregate demand.

[15] To see that, consider the policy rule $\mathtt{f}_t = \chi\mathtt{y}_t + \tilde{\chi}\tilde{\mathtt{y}}_t + \varepsilon_{p,t}$, where $\tilde{\mathtt{y}}_t = \delta_1\upsilon_{1,t} + \delta_2\upsilon_{2,t} + \delta_3\varepsilon_{p,t}$. The IV estimator of the regression coefficient in equation (4) would yield $\widehat{\gamma}_{iv} = \chi + \tilde{\chi}[\frac{\delta_1}{\beta_1}\varkappa + \frac{\delta_2}{\beta_2}(1-\varkappa)]$ (where $\varkappa$ can be positive or negative).

strument and the business cycle indicator, conditional on a single shock–provides a reliable framework for assessing whether fiscal policy interventions mitigate or amplify the transmission of the shock to aggregate demand. Identifying these conditional correlations requires an identification strategy in which the instrument isolates the effects of a single shock (that is, in (8), either $\kappa_1 = 0$ or $\kappa_2 = 0$). In that case, the IV estimate delivers $\widehat{\gamma}_{iv} = \frac{\alpha_j}{\beta_j} = \frac{df_t}{dy_t}|\upsilon_{j,t}$, where υ_j is the only shock for which $Cov\left(\mathrm{z}_t, \upsilon_{j,t}\right) \neq 0$.

To sum up, we demonstrated that identifying the cyclicality of fiscal policy requires examining *conditional* evidence–specifically, focusing on the transmission of a particular shock. In the next Section, we outline our empirical strategy to identify fiscal policy responses to export price shifts driven by commodity-specific shocks.

4 Identifying Commodity Price Shocks

We use export prices as the main channel through which fluctuations in commodity prices are transmitted to EMDEs. The high share of primary commodities in total exports (0.68 on average in our sample) and the greater volatility of commodity prices compared to other goods make them the dominant driver of export price variation. However, most commodity price fluctuations stem from changes in global demand and broader global factors, which can affect domestic economies through multiple channels beyond export prices. To isolate the causal impact of export price changes, we follow Juvenal and Petrella (2024) and construct an instrument based on price variations driven by idiosyncratic commodity market events–such as natural disasters, weather shocks, and significant local geopolitical events–that induce large price swings but are unrelated to global economic activity. This quasi-natural experiment allows us to identify the impact of export price shifts on countries not directly affected by these events.

Given the price-taker behavior of the countries in our study within each of the (global) commodity markets, an exogenous shift in a global commodity prices–driven by changes in global supply or demand *for a specific commodity*–leads to a corresponding shift in foreign demand for domestically produced commodities. Higher commodity prices, whether due to supply contractions or increased global demand within the specific commodity market, generally incentivize greater domestic production and lead to higher export revenues, ultimately boosting aggregate demand in the domestic economy. Crucially, the key assumption is that these events are commodity-specific and affect the domestic economy only through their impact on global commodity prices. Table 2 details the 24 events. For instance, the environmental policies of the Clinton Administration in 1993 led to a substantial reduction in U.S. domestic timber production, resulting in a roughly 40% increase in global prices. This can be exploited as an exogenous event to identify a shift in export prices for countries whose exports relied heavily on timber that year, including Cameroon and Malaysia. Similarly, a positive shock in the price of cotton in 2003, resulting from global shortages associated with severe weather damage to cotton crops in China, provided us with an event for an exogenous

Table 2: List of Events

Year	Commodity	Sign	Source of Shock
1993	Timber	+	Clinton's environmentally friendly policies
1993	Tobacco	-	Worldwide increase in competition for exports
1994	Aluminum	+	Reduction in stocks of major producing countries
1994	Coffee	+	Frost in Brazil
1994	Cotton	+	Decline in production due to bad weather in key producing countries
1997	Cereals/Food	-	Favorable production forecast
1998	Crude oil	-	Expectations of higher supply
1999	Cocoa	-	Supply surplus in major producing countries
2000	Natural gas	+	California gas crisis
2000	Nickel	+	Technical problems in key producing countries
2002	Cocoa	+	Attempted coup in Cote d'Ivoire
2003	Cotton	+	Severe weather damage in China
2005	Natural gas	+	Effects of hurricanes Katrina and Rita
2006	Sugar	+	Severe draughts in Thailand
2007	Lead	-	Rising stocks and resumed production from the Magellan mine in Australia
2008	Rice	+	Trade restrictions of major suppliers
2008	Soybean	+	Expectations of a reduction in supply
2010	Cereals/Food	+	Adverse weather conditions in key producing countries
2010	Cotton	+	Negative weather shocks in the U.S. and Pakistan
2010	Rubber	+	Severe draughts in Thailand and India
2015	Energy	-	Booming in U.S. shale oil production
2017	Cocoa	-	Favorable weather conditions in major producing countries
2019	Energy (excluding crude oil)	-	The U.S. became a net energy exporter
2019	Iron ore	+	Collapse of a mining dam in Brazil

Notes: This Table lists each of the episodes identified as generating large exogenous variations in commodity prices and provides a brief description of the source of the shock.

shift in the price of cotton for major cotton exporters in our sample, such as Burkina Faso.[16] Appendix C provides a detailed narrative and evidence supporting our choice of events.

To create the commodity price instrument, we generate a surprise metric for each event. This metric is calculated as the difference between the real price of the commodity and its expected price.Specifically, the surprise is defined as: $e_{\mathfrak{c},t} = p_{\mathfrak{c},t} - E_{t-1}[p_{\mathfrak{c},t}]$, where $p_{\mathfrak{c},t}$ is the (log real) price of commodity $\mathfrak{c}$ at time t (deflated using the U.S. CPI), and E is the expectation operator. The expectation of the price before the event is retrieved from a regression on the commodity's own price history as well as the overall (log) level of real commodity price indices (including lags) for the group of commodities to which the commodity does not belong. The latter set of variables is included to control for global economic conditions that affect all commodity price indices. Specifically, we estimate: $p_{\mathfrak{c},t} = \sum_{j=1}^{2} a_j p_{\mathfrak{c},t-j} - \sum_{\forall \mathfrak{g} \neq \mathfrak{g}_\mathfrak{c}} \sum_{j=1}^{2} b_{\mathfrak{g},j} p_{t-j}^{\mathfrak{g}} + e_{\mathfrak{c},t}$, where $\mathfrak{g}_\mathfrak{c}$ represents the commodity group $\mathfrak{g}$ to which commodity $\mathfrak{c}$ belongs.[17] The surprise component, $e_{\mathfrak{c},t}$, contains two pieces of information: the "sign" and the "magnitude." The "sign" must align with the direction of the shock reported in Table 2. The "magnitude" acts as a scaling factor, ensuring that events affecting the same commodity are adjusted according to the size of the unexpected price shift.[18] For

[16] To avoid selecting events that might represent both an export price shock and a capital or productivity shock, we exclude events that arise from weather conditions or political events within a specific country. For example, an attempted coup in Côte d'Ivoire in 2002, a leading cocoa-producing country, generated an increase of 66 percent in cocoa prices. This shock served us for an event for cocoa exporting countries except Côte d'Ivoire.

[17] We consider the three main commodity indices: agricultural, energy, and metals. When we evaluate, for instance, the surprise in one of the agricultural commodity prices, we include the lagged value of the energy and metal commodity price indices as a proxy of the global component.

[18] Alternative methods for calculating the expected component of the price shift will impact the magnitude

each event, j, we define $\mathfrak{q}_{j,t} = e_{\mathfrak{c},t}$ for t corresponding to the year of the event, and $\mathfrak{q}_{j,t} = 0$ for all other periods. By doing so, we assume that a predominant part of the unexpected variation in the commodity price at the time of the event can be attributed to the exogenous event.[19]

The instrument is then constructed as $z_{i,t} = \sum_j \mathbf{1}(\mathbb{w}_{i,\mathfrak{c}} > \underline{\mathbb{w}})\mathbb{w}_{i,\mathfrak{c}}\mathfrak{q}_{j,t}$, where $\mathbb{w}_{i,\mathfrak{c}}$ denotes the export weight of commodity $\mathfrak{c}$ (associated with event j) for country i, calculated as the median export weight in the 80s, and $\mathbf{1}(x)$ denotes an indicator function that takes value 1 when condition x is satisfied. The surprise component, $\mathfrak{q}_{j,t}$, reveals that the exogenous fluctuations in the export price for a country with similar exposure to relevant commodities for two distinct events are approximately proportional to the surprise in the commodity price changes that occurred during the respective events mediated by the country's export share (see upper left panel of Figure 1). Most importantly, within a panel setting, we can take advantage of the cross-sectional variation in the sensitivity of different countries to the same commodity for each of the events, i.e., $\mathbb{w}_{i,\mathfrak{c}} \neq \mathbb{w}_{j,\mathfrak{c}}$ for each $i \neq j$. Therefore, for each event, differences in the responses between countries with higher and lower exposure to commodity exports are exploited to identify the causal effect of a commodity price shock (see upper right panel of Figure 1).Lastly, we choose a lower bound $\underline{\mathbb{w}} = 2\%$, so that the term $\mathbf{1}(\mathbb{w}_{i,\mathfrak{c}} > \underline{\mathbb{w}})$ limits the amount of noise in the instruments for countries with limited exposure to the commodity price at the time of specific events.[20] In our sample, the median country experiences up to four events, while some, such as Cameroon and Côte d'Ivoire, are exposed to as many as ten major events over the period. Notably, only two countries–Armenia and Morocco–are not exposed to any of the events considered.

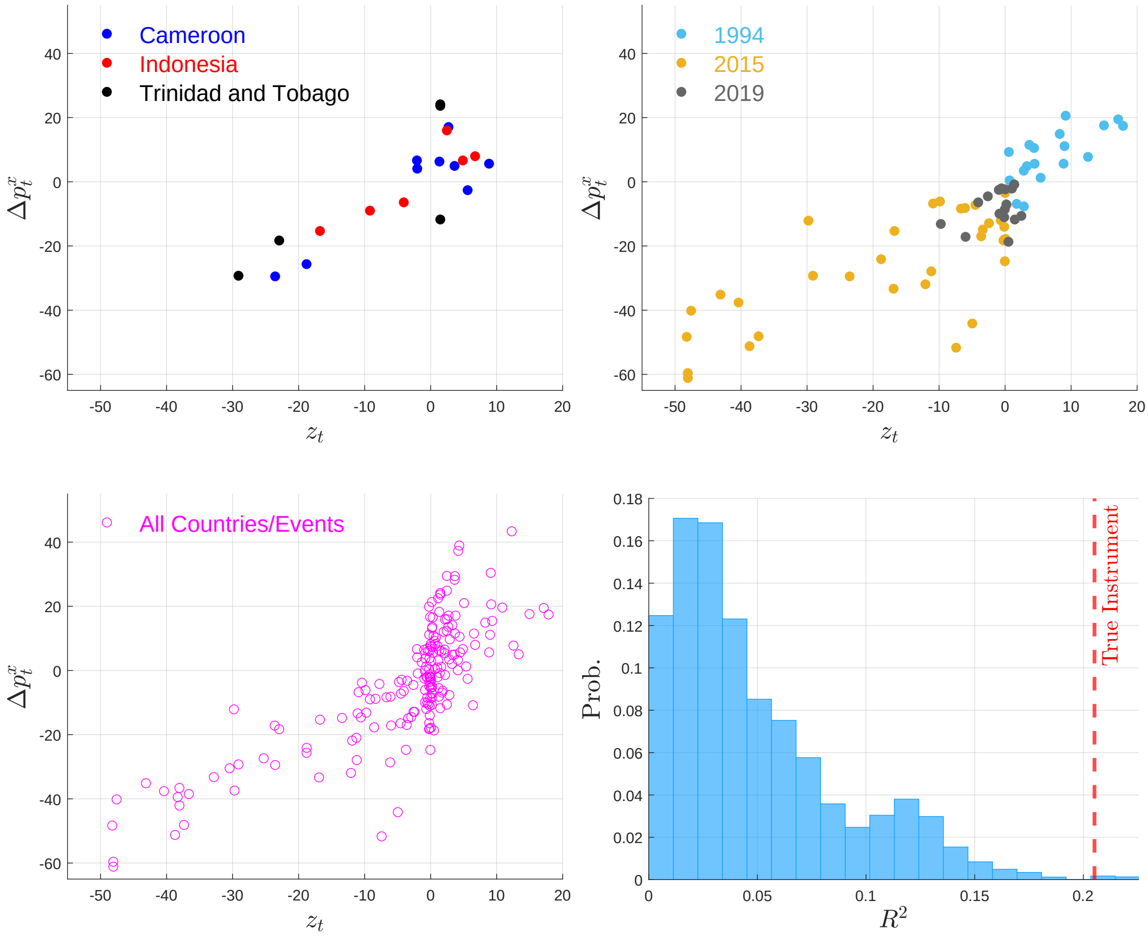
The instrument is a specific application of the shift-share research design (Bartik, 1991; Blanchard and Katz, 1992). Export shares, defined as the median over the ten years preceding our estimation sample, are predetermined with respect to the events. In fact, these shares, which reflect structural factors such as natural resource endowments or climate conditions, serve as a robust proxy for countries' exposure to commodity price shocks due to their persistent nature. Countries with export compositions heavily skewed towards a particular commodity experience significant changes in export prices due to the shock. The selection of events, corresponding to substantial changes in export prices, is equally crucial. Although each commodity represents a small portion of total exports, the volatility in export price changes for affected countries is approximately 80% higher during event periods compared to non-event periods. This heightened volatility can be traced to the significant fluctuations in the specific commodity impacted by the event. Therefore the selection of the events is a crucial step for a credible identification of the transmission of commodity price

component while leaving the sign unchanged. While the "magnitude" of the news provides useful information, other methods for computing it are certainly feasible. In recognition of this, Section 5.3 demonstrates that our results remain robust even when the instrument is constructed using only the sign component, with only a marginal loss in efficiency.

[19]This procedure is in line with the approach proposed by Hamilton (2003), who identifies oil supply shocks as reductions in oil prices from their previous peaks and shows these to be closely related to a fall in oil supply for the countries specifically affected by the event over the same period. The use of the surprise avoids the inclusion of price fluctuations into $\mathfrak{q}_{j,t}$, which would have been anticipated "ex-ante" based on the information available.

[20]The results that we report are robust to an alternative choice of $\underline{\mathbb{w}}$ at 1% or 0.5%.

Figure 1: Instrument Relevance



Notes: The first three panels reports a scatter of the changes in the price of exports (Δp_t^x) and the instrument (z_t), for a selection of countries (upper left panel), events (upper right panel) and all events/years (bottom left panel), considering only the periods where $z_t \neq 0$. The right left panel reports the R^2 of the regression $\Delta p_{i,t}^x = a_i + bz_{i,t} + e_{i,t}$, for all the observations in the sample, for the true instrument and the histogram of values consistent with a placebo experiments where the dates of the events are resampled randomly.

shocks.

To illustrate the latter point, we create 10,000 "placebo" instruments by replicating the procedure described above with randomly assigned events (maintaining the original subset of commodity events). Despite events appearing in only 13% of the sample, the R^2 associated with our instrument explains roughly 20% of the variability in export price changes over the entire sample, after controlling for country fixed effects. This is larger than the 99.7% quantile of all possible R^2 values obtained from the placebo exercise (see Figure 1, bottom right panel), highlighting the relevance of commodity price variability during the identified dates for instrument validity. Apart from the relevance condition, the instrument must be conditionally independent of other determinants of the outcome variable for unbiased instrumental variable estimates. This depends on the plausibility of export share exogeneity (Goldsmith-Pinkham et al., 2020) or shift exogeneity (Borusyak et al., 2022), i.e., the assumption that identified shocks are not systematically related to other determinants of commodity prices.

In our framework, even when using pre-sample weights, we cannot rule out the possibility that countries with similar shares may exhibit other features that predict changes in the outcome through channels unrelated to the endogenous variable, thus violating the exclusion restriction. Therefore, exogeneity relies on the assumption that the shifts are effectively random. Specifically, major idiosyncratic, commodity-specific events that generate exogenous variation in $\mathfrak{q}_{j,t}$ should impact a country only through their effect on export prices. In this setting, shift-based identification follows from the simple observation that a share-weighted average of a random shift is itself as-good-as-random (Borusyak et al., 2024). To verify this, we test and do not reject the orthogonality of each country-level instrument to several key global factors: The BAA spread, the VIX, world real GDP growth, and the U.S. real interest rate. This confirms that the variation in shifts is uncorrelated with other global factors influencing commodity prices and, more broadly, the business cycle of EMDEs.

Therefore, unanticipated variations in commodity prices during major, commodity-specific events–modulated by the significance of each commodity in the total export basket–result in exogenous fluctuations in export prices for the countries under investigation (see the bottom left panel of Figure 1). This creates a robust instrument, associated with an F-statistic of over 500 in the first stage regression, which can be used to calculate the local average treatment effect (LATE, see Imbens and Angrist, 1994) of a boom in export prices driven by exogenous commodity price shocks.

5 Empirical Analysis

We use the framework proposed by Cloyne et al. (2023), which expands upon the conventional LP method (Jordà, 2005) to incorporate interaction effects and estimate the following regression:

$$y_{i,t+h} - y_{i,t-1} = \mu_i^h + \Delta p_{i,t}^x \beta^h \ + \ (\mathbf{x}_{i,t} - \bar{\mathbf{x}}_i)\boldsymbol{\gamma}_0^h \ + \ \Delta p_{i,t}^x (\mathbf{x}_{i,t} - \bar{\mathbf{x}}_i)\boldsymbol{\theta}_x^h \ + \ \omega_{i,t+h}, \tag{10}$$

for $h = 0, 1, ..., H$. Here, the dependent variable represents the cumulative change in country i outcome variable y from year $t-1$ to $t+h$, Δp^x is the change in (the log of detrended) export prices, and β^h captures the average dynamic causal effect associated with the intervention variable. We control for country fixed effect, μ_i^h, and a set of additional covariates $\mathbf{x}_{i,t}$ (in deviation from their mean, $\bar{\mathbf{x}}_i$). Moreover, we follow Cloyne et al. (2023) and include interaction terms $\Delta p_{i,t}^x(\mathbf{x}_{i,t} - \bar{\mathbf{x}}_i)$. This modification generalizes the traditional LP which typically assumes $\boldsymbol{\theta}_x^h = 0$ and allows for (potential) heterogeneity in the causal effect arising from the interplay between intervention and control variables. Therefore, in our setting, $\mathbf{x}_{i,t}$ are both control variables as well characteristics of the treated subpopulation that may influence the way in which treatment affects outcomes.

The outcome variables used in our analysis are the log of GDP (detrended), log of export prices (detrended), the log of the EMBI spread, and a battery of fiscal indicators. In turn, $\mathbf{x}_{i,t}$ includes two lags of the (change in the) dependent variable, as well as two lags of real GDP growth, export price growth, the BAA spread, and the primary balance, both as a control and interacted.[21] The change in export prices is instrumented as discussed in Section 4, therefore the reported impulse response functions (IRFs) can be interpreted as the LATE (see, e.g., Jordà et al., 2020).

5.1 Baseline Results

Figure 2 displays the average effect of a one standard deviation increase in export prices driven by commodity-specific shocks.[22] Panel (a) shows the impact on macro variables. In response to a commodity price shock, there is a quite persistent increase in export prices. The increase in export prices leads to a steady increase in domestic GDP, which is what would be expected from a positive terms-of-trade shock. Moreover, the increase in export prices leads to a mild decrease in borrowing costs, as shown by the decline of the EMBI spread.

The presentation of the fiscal results are categorized into three domains: fiscal *instruments* (Panel b), the primary balance as a key fiscal *outcome* (Panel c), and other fiscal *outcomes* (Panel d). As emphasized before, fiscal *instruments*, government spending and the VAT rate, provide guidance to define the cyclicality of the fiscal response, the primary balance is a core metric within fiscal outcomes since it is a summary of the fiscal stance. Other fiscal *outcomes* include government revenues, and fiscal variables with respect to GDP.[23] It is pertinent to acknowledge the non-discretionary nature of these variables.

Panel (b) of Figure 2 displays the response of fiscal instruments to an increase in export prices induced by commodity-specific shocks. In response to a commodity shock government spending shows a positive and significant response, reaching a peak three years post-shock. The VAT rate increases significantly, although with a lag, following a commodity boom. This implies that while spending is procyclical, tax rates are countercyclical.

The primary balance (Panel c) increases on impact, peaks after one year and then reverts slowly. Government revenues surge immediately but start tapering off after the first year (Panel d). The revenue patterns in commodity-exporting countries are as anticipated, given the direct association between government revenues and commodity sector performance (Céspedes and Velasco, 2014). This happens because there is a "base" effect, whereby the positive commodity shock leads to a higher GDP and since revenues depend positively on GDP, they increase. In addition, commodity-linked revenues, either from taxes on production, royalties, or profits, rise during commodity booms.

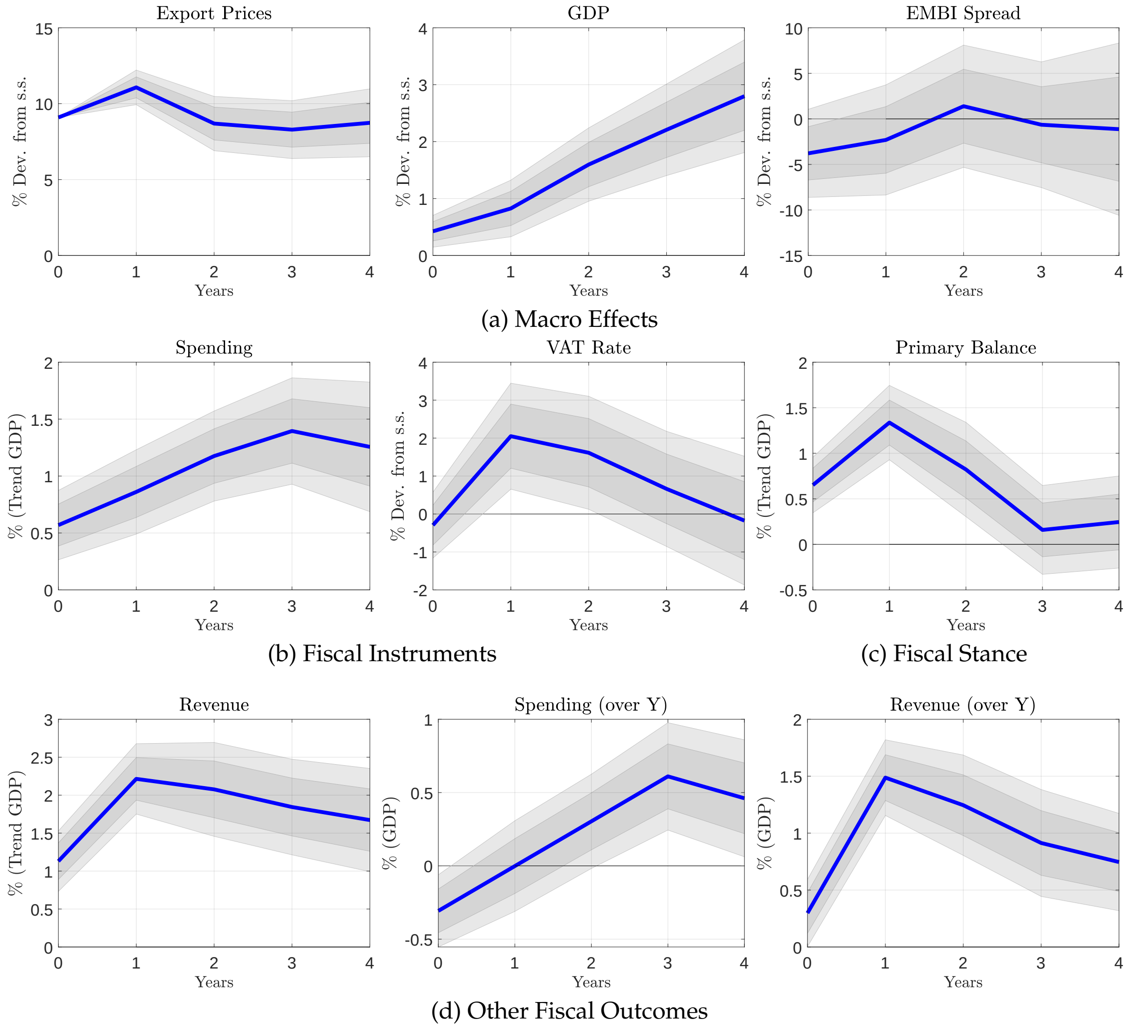
The response of the VAT rate is particularly noteworthy. First, it shows an increase, which contrasts with the negative correlation in Table 1. Second, it differs from the procyclical re-

[21] The interaction terms include only one lag of the same variables. We include only a single lag in the change of VAT rate as a control to minimize the loss of observations, since this variable has the most incomplete data in our dataset.

[22] If we extend the horizon, all impulse response functions exhibit mean reversion. However, the bands become considerably larger after four periods. This can be attributed to the fact that many countries have relatively short samples.

[23] We measure spending and revenues with respect to the trend of GDP in local currency units, computed using an HP filter (with λ=100). To take care of outliers, we eliminate the data for which at least one of the ratios of revenue, expenditure, and the primary balance over trend GDP is larger that 100 (in absolute value for the primary balance).

Figure 2: Dynamic Causal Effect to a Commodity Price Shock



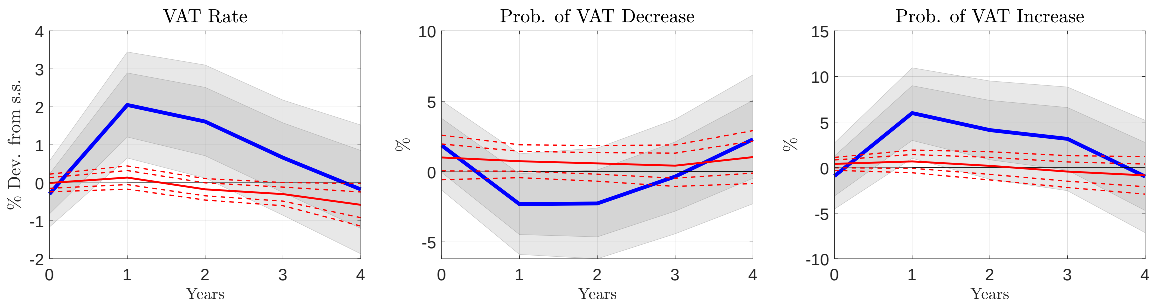
Notes: The Impulse Responses show the LATE (in blue) of one standard deviation increase in Px driven by commodity price shocks. Gray areas denote 68% and 90% confidence intervals.

sponse documented by Végh and Vuletin (2015). Figure 3 provides additional insights into the VAT rate response. The top left chart compares the LATE baseline response with the ATE response estimated using OLS without instrumenting export prices.[24] The ATE response shows no statistically significant effect, consistent with the weak unconditional correlation between export prices and VAT rates (Table 1), highlighting the importance of a robust identification strategy to capture the causal impact of commodity price changes on fiscal policy.

VAT rates are generally stable, with infrequent changes. We assess whether a boom in export prices, triggered by a commodity price shock, affects the likelihood of VAT rate changes. To that aim, we run the LP regression in equation (10) where the dependent variable takes the value of zero if there is no VAT rate change and a value of 1 if there is a change between $t-1$ and $t+h$ with $h = 0, ..., 4$ (so that the LP takes the form of a linear probability model

[24]In the next Section, we repeat the same exercise for all other variables, confirming the existence of substantial differences between the IV and OLS estimate.

Figure 3: VAT Response to a Commodity Price Shock



Notes: The Impulse Responses show the LATE (in blue) of one standard deviation increase in Px driven by commodity price shocks, for the VAT (left panel), the probability of observing a VAT decrease and increase (middle and right panel). Gray areas denote 68% and 90% confidence intervals. All figures also report, in red, the LP coefficients estimated without instrumenting export prices (i.e., OLS coefficients). Dashed lines denote 68% and 90% confidence intervals.

with IV). For the exercise, we distinguish between VAT decreases and increases (middle and right panels of Figure 3, respectively).[25] The results show that a commodity price shock does not significantly raise the probability of a VAT rate decrease, in fact it marginally decreases the probability of a VAT cut. Moreover, it is associated with a significant increase in the likelihood of a VAT rate hike in the year following the shock. For context, the unconditional probability of a VAT increase within a two-year period in our sample is approximately 9.5%.

Following a commodity price shock that leads to a one standard deviation increase in export prices, this probability increases to roughly 17%. This represents a substantial increase, comparable to the unconditional probability of a VAT rate increase within any four-year period in our sample. Examining VAT rate changes in the year following each event, we find that adjustments align with the instrument's sign 67% of the time, compared to only 54% when using the contemporaneous export price change. This pattern underscores the importance of exogenous export price variations in identifying VAT responses.

Ideally, one would assess whether VAT changes during event years are at least partially endogenous rather than exogenous by analyzing the narrative behind legislative tax reforms (see, e.g., Cloyne, 2013; Romer and Romer, 2010). This classification is challenging even for advanced economies, as policymakers rarely cite specific shocks they aim to address, and the task is even more difficult in developing countries. In our sample, the most significant VAT reduction during an identified exogenous event occurred in Ghana in 2018, when the VAT rate was cut by 2.5 percentage points, nearly 20% from its original level. A speech by Finance Minister Ken Ofori-Atta to the Ghanaian Parliament suggests the cut was countercyclical. The minister discussed the decline in cocoa prices and its impact on Ghana's economy, stating that the government sought to *"use tax policy as a tool to stimulate investment and to shape economic behavior."*[26]

[25] This analysis also allows us to verify whether the baseline results, indicating a significant VAT increase following a commodity price shock, are not driven by a few large tax changes coinciding with our identified commodity events.

[26] The minister also noted that the cocoa price decline in 2017 was "due to strong West African

Table 3: Conditional Correlations

	Corr w/ GDP			Corr w/ Px		
	Base	Low QI	High QI	Base	Low QI	High QI
Spending	0.96	0.65	0.99	0.94	0.97	0.84
VAT Rate	0.47	-0.70	0.78	0.68	-0.90	0.81
Primary Balance	0.54	-0.98	0.86	0.88	-0.74	0.97
Revenue	0.88	-0.83	0.97	0.98	-0.40	0.97
Spending (over Y)	0.85	-0.17	0.69	0.49	-0.03	0.35
Revenue (over Y)	0.81	-0.83	0.96	0.93	-0.34	0.88
EMBI Spread	-0.30	0.86	-0.79	-0.63	0.29	-0.63

Notes: For each variable listed we report the conditional correlations calculated from the IRFs reported in Figures 2 and 5. The left columns measures the correlation with respect to detrended GDP, whereas the right column measures the correlation with respect to detrended export prices. Low QI and High QI denote low and high institutional quality, respectively.

Conversely, one of the largest VAT increases in our sample took place in El Salvador in 1995. According to an IMF Staff Country Report,[27] the VAT rate rose from 10% to 13% amid a consumption boom over this period and is no. This increase was not motivated by fiscal consolidation efforts. Similarly, Jamaica's VAT hike in 1995 is not listed among fiscal measures driven by consolidation in David and Leigh (2018).[28]

The findings indicate a procyclical nature of government spending conditional on the commodity shock: It expands with positive commodity shocks and contracts with negative ones. By contrast, revenues and taxes exhibit a countercyclical behavior conditional on the commodity shock, acting as stabilizers during business cycles–they expand following positive commodity shocks and contract after negative shocks. Taken together, the primary balance serves as a key indicator of the fiscal stance and, more broadly, reflects progress in repaying public debt.

We also examine the dynamics of spending, and revenue relative to GDP. The responses are shown in Panel (d) of Figure 2. Given that a positive commodity shock leads to an increase in spending, revenues, the primary balance, and GDP, the effect of fiscal ratios in terms of GDP can, in principle, take on any value. The results indicate that government spending relative to GDP initially declines but rises after one year. Conversely, revenues as a proportion of GDP increase on impact, peaking after one year and reverting thereafter. The same happens with the primary balance-to-GDP ratio.

Table 3 is the counterpart of Table 1 but the correlations are calculated conditional on the export price shock, from the IRFs reported in Figures 2 and 5.[29] Zooming into the baseline

output and increased production in Latin America," aligning with our narrative identification. Full speech available at `https://www.modernghana.com/news/816731/ghana-2018-budget-full-text-presented-by-ken-ofori-atta.html`.

[27]Available at `https://www.elibrary.imf.org/view/journals/002/1998/032/article-A002-en.xml?rskey=aP5jSa&result=34`.

[28]However, the 1995 budget was identified as motivated by fiscal consolidation for Guatemala and Costa Rica. Excluding these countries from our sample does not affect our results.

[29]The correlations are computed directly from the impulse response coefficients following Den Haan (2000). Specifically, we compute conditional covariances (variances), as the cumulative sum of the cross-product (squares) of the impulse response coefficients. Therefore, the conditional correlation between two generic vari-

results, three important observations stand out from this table. First, government spending remains positively correlated with GDP. Second, in contrast with the results in Table 1, the VAT rate shows a countercyclical pattern. Finally, the primary balance increases with GDP and with export prices. Therefore, our findings challenge the conventional wisdom of balanced budget rules. While a balanced budget behavior is often observed unconditionally, as illustrated in Table 1, our results conditional on the export price shock reveal a deviation from a balanced budget behavior.

5.2 Do Fiscal Interventions Amplify the Transmission of Shocks?

Our analysis reveals that government spending is procyclical, while tax rates exhibit a countercyclical pattern. However, given the opposing cyclical behavior of spending and taxes, it is unclear whether fiscal policy amplifies or mitigates these shocks. To assess this, we conduct a counterfactual exercise in which we hold government spending, tax rates, or both constant in response to the commodity shock. Specifically, we estimate the following equation:

$$y_{i,t+h}-y_{i,t-1}=\widetilde{\mu}_i^h+\Delta p_{i,t}^x\widetilde{\beta}^h+(\mathbf{x}_{i,t}-\bar{\mathbf{x}}_i)\widetilde{\gamma}_0^h+\Delta p_{i,t}^x(\mathbf{x}_{i,t}-\bar{\mathbf{x}}_i)\widetilde{\boldsymbol{\theta}}_x^h+(\mathtt{f}_{i,t+h}-\bar{\mathtt{f}}_i)\widetilde{\boldsymbol{\theta}}_f^h+\widetilde{\omega}_{i,t+h}, \quad (11)$$

where $\mathtt{f}_{i,t+h}$ represents the cumulated changes in the fiscal instrument between period t and period $t+h$. The estimated coefficients $\widetilde{\beta}^h$ in equation (11) capture the effect of commodity shocks on the outcome variables at horizon h holding government spending, taxes, or both constant over the same time period. Therefore, this provides an estimate of the direct effects of commodity shocks. Any differences between the estimated $\widetilde{\beta}^h$ in equation (11) and β^h in (10) reflect the indirect effects of fiscal policy in shaping the transmission of commodity shocks.

Figure 4 compares the effects of commodity shocks on GDP and the primary balance under the baseline scenario (solid blue) and three counterfactual exercises (dashed black), estimated using equation (11). Panel (a) presents the counterfactual in which the VAT response is held constant. In the absence of an increase in VAT, GDP rises more sharply, while the primary balance improves less relative to the baseline. This suggests that the VAT response plays a stabilizing role by dampening the output expansion. Panel (b) examines the case where government spending is held fixed. In this case, GDP remains largely unchanged, but the primary balance improves more significantly, indicating that increased spending in the baseline partially offsets fiscal consolidation. Panel (c) jointly holds both spending and the VAT rate constant.

The shift in taxes more than offsets the impact of commodity price shocks on aggregate demand, despite the increase in spending. Additionally, the combined effect of higher tax rates and an expanded tax base plays a dominant role in shaping the overall response of the primary balance. Taken together, these findings indicate that the fiscal policy response to commodity shocks is countercyclical, smoothing GDP fluctuations while strengthening the primary balance.

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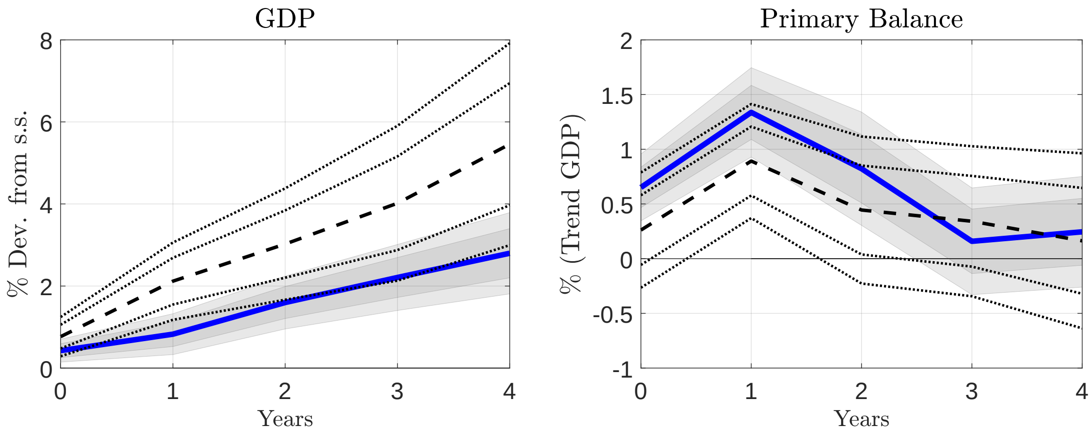
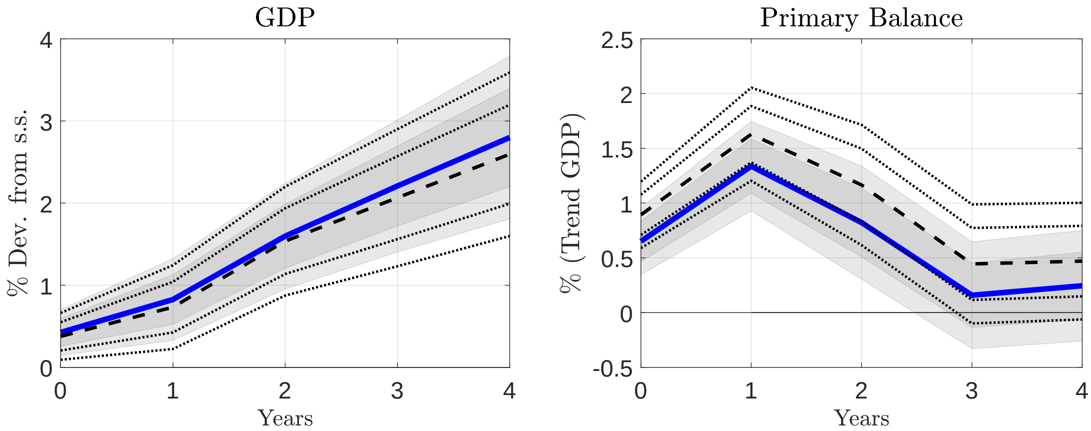
ables $\mathtt{z}$ and $\mathtt{w}$ is calculated as $\sum_h(\beta_{\mathtt{z}}^h,\beta_{\mathtt{w}}^h)/\sqrt{\sum_h(\beta_{\mathtt{z}}^h)^2\sum_h(\beta_{\mathtt{w}}^h)^2}$, where $\beta_{\mathtt{z}}^h$ ($\beta_{\mathtt{w}}^h$) denotes the impulse response associated with variable $\mathtt{z}$ ($\mathtt{w}$) at horizon h.

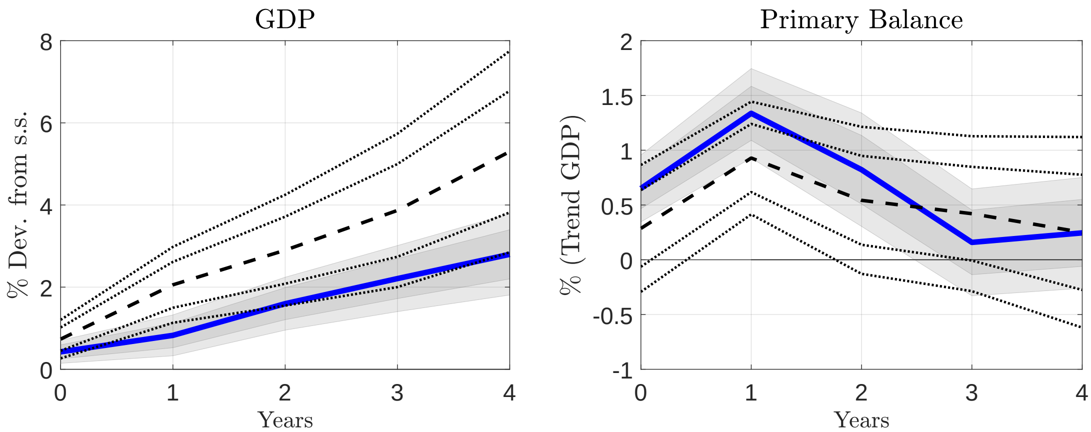
Figure 4: Counterfactual Analysis of Fiscal Responses



(a) Holding Fixed VAT Response



(b) Holding Fixed Spending Response



(c) Holding Fixed Spending & VAT Response

Notes: The Impulse Responses show the LATE (in blue) of one standard deviation increase in Px driven by commodity price shocks. Gray areas denote 68% and 90% confidence intervals. In black we report the counterfactuals, with confidence intervals.

5.3 Sensitivity Analysis and Additional Results

In this Section, we summarize the main takeaways of some additional results and sensitivity analysis. The results are presented in Appendix D.

LATE vs. OLS. The empirical methodology conventionally used in previous studies is based on unconditional correlations and would be akin to determining the impact of export price changes using ordinary least squares (OLS), as opposed to the IV methodology we implement. The OLS approach yields a mean response, amalgamating the effects of both exogenous shifts in export prices and those originating in endogenous responses to global shocks. The results derived from OLS (as shown in Appendix Figure D.1) differ significantly from the effects of increased export prices driven by commodity-specific shocks, which we identify using IV. The response of several variables is notably different. For instance, with OLS, the response of the EMBI spread is muted and the impact on the VAT rate is flat and insignificant. Moreover, many of the fiscal variables (e.g., primary balance, spending, and revenues) exhibit a smaller response, consistent with the weaker unconditional correlations observed in Table 1.

Alternative Instruments. To assess the robustness of our findings to the importance of the measurement of the "news" ($\mathfrak{q}_{j,t}$) in the construction of the instrument, we use an alternative instrument that relies on the direction of the shock for each event and the heterogeneous country exposure, determined by the commodity's weight in the export share. The alternative instrument, $\tilde{z}$, is constructed as follows: $\tilde{z}_{i,t} = \sum_j \mathbf{1}(\mathbb{w}_{i,\mathfrak{c}} > \underline{\mathbb{w}})\mathbb{w}_{i,\mathfrak{c}}\text{Sign}_{j,t}$, where the signs are the ones reported in Table 2. The results obtained using this alternative instrument (shown in Appendix Figure D.2) are consistent with our baseline findings, underscoring the importance of cross-sectional heterogeneity in exposure to each commodity for identification. This demonstrates that our conclusions are robust to the specific measurement of the "news" component in the instrument construction and only depend on the sign of the event being correctly identified. This exercise also highlights that concerns regarding the identification of the news component and its potential "contamination" by other global shocks are unwarranted. Such concerns would only be valid under two conditions: first, if an alternative shock were systematically related to the pattern of signs dictated by the identified event; and second, if there existed an economic mechanism through which the effect of this alternative shock was proportional to the exposure to the specific commodity in question.

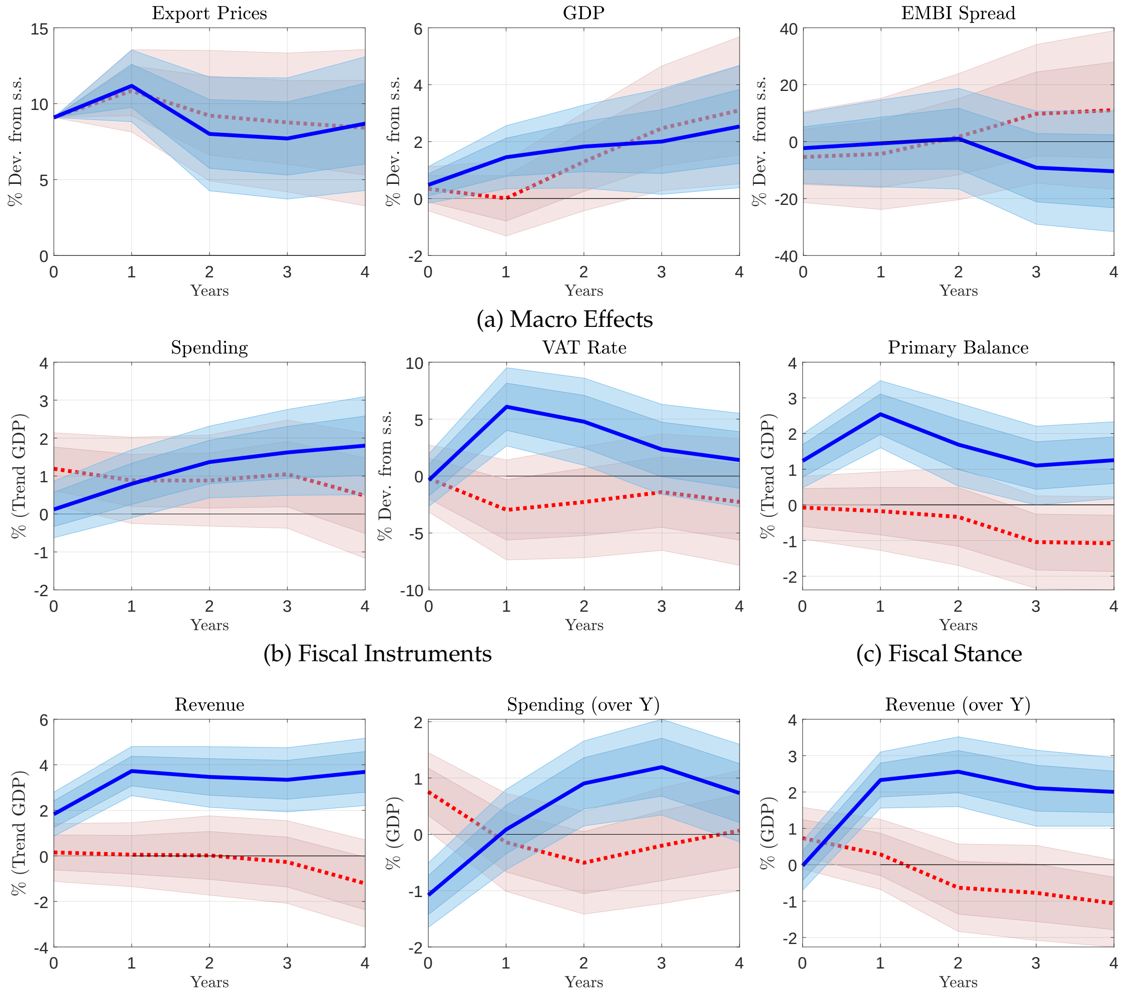
We also construct an instrument that modifies the constant weights used in the baseline to time-varying export shares from the year prior to the event: $\tilde{z}_{i,t} = \sum_j \mathbf{1}(\mathbb{w}_{i,\mathfrak{c},t-1} > \underline{\mathbb{w}})\mathbb{w}_{i,\mathfrak{c},t-1}\mathfrak{q}_{j,t}$. Following Borusyak et al. (2022), the exogeneity of this instrument stems from the exogeneity of $\mathfrak{q}_{j,t}$, which is ensured through the selection of events. This instrument also captures the evolving reliance of developing countries on commodity exports. The results, shown in Appendix Figure D.3, remain consistent with our baseline findings.

Placebo Exercise. To address the concern about the robustness of our results and the importance of the choice of events, we conduct a placebo test. In this exercise, we randomly reshuffle the instrument, maintaining the original subset of commodity events but assigning them to random dates. This ensures that the new instrument has the same number of events and composition in terms of commodities considered. We estimate impulse responses using 10,000 placebo instruments and plot the median point estimates along with the 68% and 90% point-wise quantiles (Appendix Figure D.4). As expected, the placebo instrument is insignificant, and the confidence bands are extremely wide, yielding no statistically significant effects. While the mean of the point estimates of the coefficients is similar to the OLS results, the variance is substantially larger. This placebo exercise confirms that we cannot identify the causal effect when commodity events are resampled randomly, highlighting the importance of the specific choice of events in our analysis.[30]

Omitting Countries/Events. One concern in our analysis is related to the possibility that a country could be playing a large role in driving the results. We therefore assess the sensitivity of our findings by excluding from the sample one country at a time. The results are presented in Appendix Figure D.6. Although the commodity events that we selected are idiosyncratic and unrelated to the business cycle, some of them overlap with recession periods. We therefore check the robustness of the effects of an increase in export prices driven by commodity-specific shocks when we exclude from our sample events which can be contaminated by major crises at a global scale.[31] The impulse responses are shown in Appendix Figure D.7. Our results remain robust in both cases, with only the GDP response appearing slightly overstated when events from major recessions are included in the sample.

Additional Shocks. The contrast between unconditional correlations and those conditional on commodity shocks suggests that other shocks influence fiscal variables differently. To further illustrate this, we follow Juvenal and Petrella (2024) and examine the fiscal response to a decline in the BAA spread driven by U.S. monetary policy shocks and shifts in global risk appetite. The results, presented in Appendix Figures D.8 and D.9, indicate that while both shocks lead to a temporary rise in commodity prices and a persistent boost to EMDEs business cycles, their fiscal effects differ significantly. An improvement in global financial conditions following U.S. monetary easing results in a persistent increase in government spending but only a modest–and statistically insignificant–increase in taxes.[32] By contrast, shifts in global risk appetite lead to a sharp decline in government spending, while tax changes remain relatively muted. As a result, taxes remain acyclical in both scenarios. However, fiscal spending is strongly countercyclical after a global financial shock and highly procyclical when global financial conditions are improved by easier U.S. monetary policy.

Figure 5: Dynamic Causal Effects of a Commodity Price Shock: the Role of Institutional Quality



(a) Macro Effects

(b) Fiscal Instruments

(c) Fiscal Stance

(d) Other Fiscal Outcomes

Notes: The Impulse Responses show the LATE (in blue) of one standard deviation increase in Px driven by commodity price shocks. Higher quality of institutions countries are in blue and lower quality of institutions countries in red. Areas denote 68% and 90% confidence intervals.

5.4 The Role of Institutional Quality

Figure 5 illustrates how the quality of institutions is an important determinant of the heterogeneity in the impact of commodity shocks on macro variables and the fiscal effects. Institutional quality is gauged using a political risk index from the International Country Risk Guide (ICRG), ranging from 0 to 100, where a higher value indicates higher quality. Given the

[30] One potential concern is that the broad range of LATE estimates identified in the placebo exercise could result from the weak relevance of the instrument generated by randomly selecting events. To address this, Figure D.5 presents the same chart, but only retains placebo instruments with a high F-statistic in the first-stage regression (F-stat > 100). Even in this case, the causal effect remains unidentified, highlighting the critical importance of constructing a valid instrument to capture the transmission of commodity shocks.

[31] These events include the Asian crisis in 1997 and the global financial crisis and associated recovery in 2008-2010.

[32] The similarity to our baseline findings underscores the central role of commodity prices in the global transmission of U.S. monetary policy to EMDEs (as highlighted by, e.g., Degasperi et al., 2020; Juvenal and Petrella, 2024).

index's stability over time, we take the median ratings over the analyzed period and include this as an interaction term for the treatment, which we instrument using the interacted instrument. This approach enables us to identify the LATE for countries with high institutional quality (depicted in blue) and compare it to those with lower institutional quality (shown in red).[33]

Countries with both high and low institutional quality exhibit comparable initial responses in export prices to a commodity shock (Figure 5, Panel a). GDP is initially muted in low-institutional quality countries while it increases in high-quality ones. Over time, however, the response of the EMBI spread diverges: it declines in countries with stronger institutions but follows an upward trajectory in those with weaker institutions. This pattern is consistent with Arezki and Brückner (2012), who find that rising commodity prices lower default probabilities in strong institutional settings but increase default risk in environments with weaker institutional quality.

The response of government spending is initially similar across both groups but diverges over time, with higher institutional quality countries showing a greater increase (Panel b). These findings suggest a stronger procyclicality in government spending within countries of higher institutional quality. In these countries, taxes display a clear countercyclical trend, acting as economic stabilizers. In contrast, in countries with lower institutional quality, taxes display a procyclical behavior (although confidence level includes the zero).

In countries with higher institutional quality, there is a larger increase in the primary balance and revenue (Panels c and d). In fact, despite a stronger increase in spending, the combination of higher taxes and an improved tax base leads to a positive and sustained primary surplus. In contrast, countries with lower institutional quality initially display a stable primary balance, turning negative two years after the shock.

Figure 5 (Panel d) also shows the effects of other fiscal outcomes in terms of GDP, split by institutional quality. When analyzing the impact of fiscal ratios, there is a stark contrast between countries with high and low quality of institutions. While the results for the former mimic those of the baseline LATE, whereby spending relative to GDP goes down; revenue over GDP increases, implying that the primary balance over GDP goes up; for the latter both spending and revenue over GDP go up, yielding a muted response of the primary balance relative to GDP.

Focusing on the conditional correlations derived from the IRFs when we split the sample of countries by institutional quality, Table 3 highlights some contrasts between the two group of countries. Both groups exhibit procyclical spending. However, in countries with low institutional quality, the VAT rate is negatively correlated with GDP and export prices, whereas in countries with strong institutions there is clear evidence of countercyclical tax policy. Similarly, the primary balance is negatively correlated with GDP in low-quality institutions countries but positively correlated in countries with high institutional quality.[34]

[33]Specifically, we extend Eq.(10) by adding the term $\Delta p^x_{i,t} QI_i \beta^h_{QI}$, allowing us to estimate how the impact of export price shocks varies with institutional quality, QI_i, as $\beta^h + \beta^h_{QI} \times IQ_i$. Figure 5 shows the results for $QI_i = 30$ and $QI_i = 80$, (approximately) the lowest and highest values in our sample.

[34]These results contrast with the unconditional correlations for both groups (see Table B.1), which suggest

In countries with strong institutions, fiscal policy tends to be more countercyclical, with higher tax revenue and improved tax collection during commodity booms. This allows for increased spending without jeopardizing fiscal sustainability, leading to a decline in the spending-to-GDP ratio as GDP growth outpaces the growth in spending. By contrast, countries with weaker institutions struggle to generate sufficient revenue, making their fiscal responses less effective in smoothing economic fluctuations. Strengthening institutional quality can therefore enhance fiscal resilience and improve the management of commodity-driven economic cycles.

6 Understanding the Fiscal Responses

Our empirical analysis reveals that fiscal policy in EMDEs is characterized by procyclical spending but countercyclical tax adjustments. This countercyclical tax response, combined with an improved tax base, leads to a stronger fiscal position, as reflected in the primary balance. In this section, we evaluate whether these empirical findings align with the predictions of an optimal (Ramsey) policy in a model that accounts for financial market imperfections. These frictions, a defining characteristic of developing economies, have been shown to significantly influence the procyclicality of fiscal policy in small open economies (see Riascos and Végh, 2003; Fernández et al., 2021). Our analysis is based on a multi-good model featuring importable, exportable, and non-tradable goods, along the lines of a traditional MXN model (Mendoza, 1995; Schmitt-Grohé and Uribe, 2018).

The production of exportable and non-tradable goods occurs domestically, whereas importable goods are exclusively produced abroad. Final output is derived from labor and imported intermediates. The utility function incorporates separable preferences, where agents derive utility from both private and government consumption, each modeled as CES bundles that include imports, exports, and non-tradable goods.[35] Government spending is financed through either foreign currency-denominated debt issued on international markets or distortionary consumption taxes, which introduce a wedge between the marginal rate of substitution and the sectoral marginal product of labor. There are incomplete asset markets, and households and the government can only borrow internationally in foreign units (as in Di Pace et al., 2024). The model incorporates a credit spread channel, where higher export prices lower borrowing costs (as in Drechsel and Tenreyro, 2018). We assume the law of one price holds for both exportable and importable goods, and EMDEs are modeled as small open economies, acting as price-takers in global markets.[36] The full details and derivations of the model are provided in Appendix E.

In what follows, we examine the optimal fiscal policy response to export price shocks.

VAT acyclicality in high-quality institutional settings and the primary balance only weakly correlated with GDP and Px for both groups.

[35] We omit physical capital from the model. To capture the hump-shaped responses seen in the data, we incorporate external habit preferences in both private consumption and government spending.

[36] We focus on the impact of an exogenous shock to export prices in foreign currency, while keeping import prices in foreign currency fixed. The law of one price in import markets implies that domestic import prices move in tandem with the real exchange rate.

First, we demonstrate that, consistent with our empirical evidence, this response involves a combination of increased spending and taxes. We subsequently describe the key channels that rationalize these findings. Second, we present robustness exercises to validate our findings. Third, we show that the outcomes of the Ramsey optimal policy can be replicated by simple, implementable fiscal rules governing spending and revenues. Fourth, we emphasize the notable differences between the model-implied conditional and unconditional correlations of fiscal variables with GDP and export prices. Lastly, we extend the baseline model to account for the empirical results on institutional quality.

6.1 Baseline Impulse Responses

Figure 6 displays the responses of key macro aggregates. The responses of fiscal variables and ratios are shown in Figure 7. In this Section, we focus on the discussion of the transmission channels based on the baseline model with incomplete markets. The Figures also include the impulse responses under two restricted settings: one without a credit spread channel and another one assuming financial autarky. Contrasting these results against the baseline model will allow us to highlight the role of the major channels driving the optimal fiscal response which will be discussed in Section 6.2.

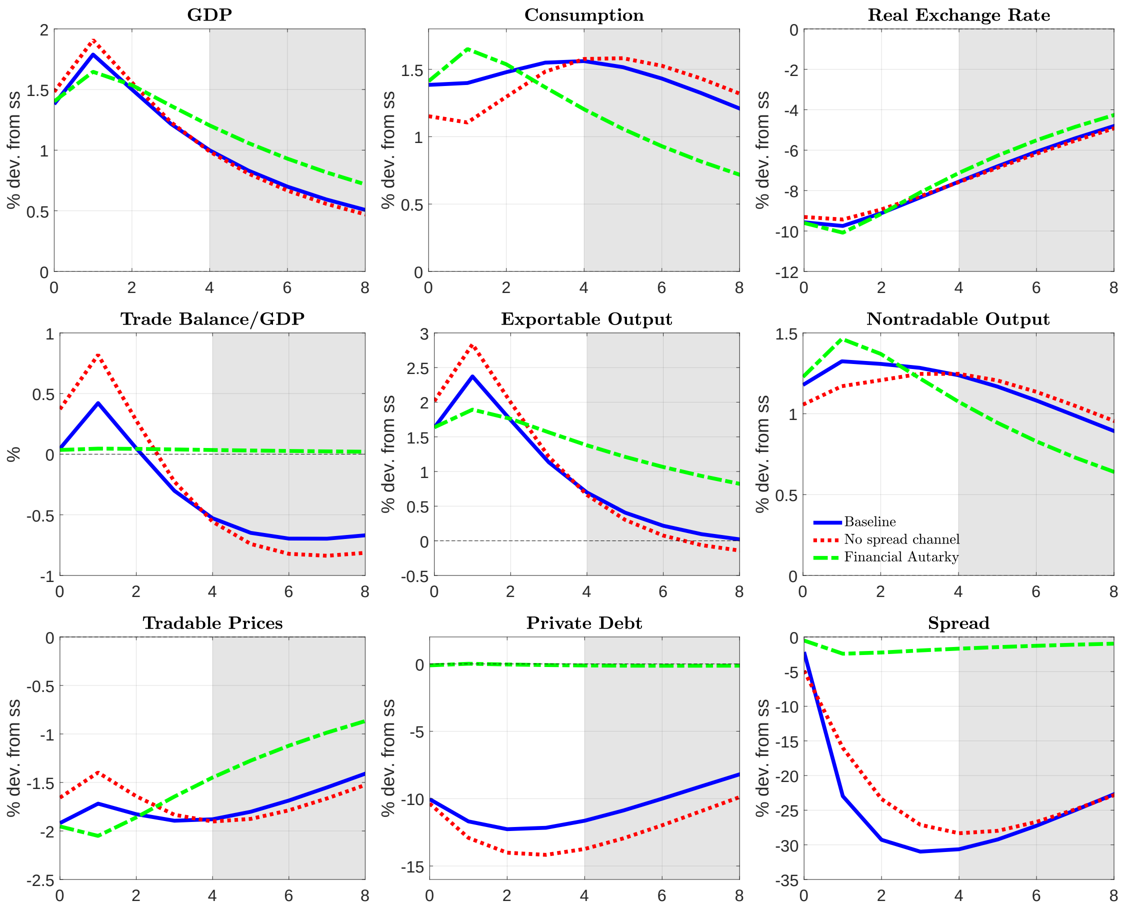
Following an export price shock, the exportable sector experiences an initial surge in activity due to higher prices, which then wanes after the first year. An increase in export prices triggers a substitution effect, reducing domestic demand for exportable goods as households shift towards importable and non-tradable goods. Simultaneously, a wealth effect increases demand across all goods, including non-tradables. The appreciation of the real exchange rate, resulting from the shock, lowers the cost of imported goods and reduces the value of existing foreign-denominated debt, further boosting consumption. Initially, the rise in export prices improves the trade balance due to increased export revenues, but as the positive impact on exports wanes, this improvement fades, and the trade balance eventually turns negative. In the non-tradable sector, the wealth effect associated with the positive export price shock leads to increased demand for both private and public goods. This surge in demand drives an expansion in non-tradable output and prices.

Given the symmetric treatment of private and public consumption in the utility function, the optimal fiscal policy response is associated with synchronized movements in government spending and private consumption.[37] A desire to smooth (public and private) consumption over time leads to a sustained increase in public spending, which requires an improvement in the primary balance and a transitory increase in public savings (Figure 7).[38] The rise in consumption expands the tax base. However, the resulting revenue increase from this "base effect" is insufficient to simultaneously finance higher spending and increase public savings.

[37]The presence of external habits in the utility function leads to more persistent responses without substantially affecting the qualitative results. The Ramsey planner's internalization of habit dependence introduces a small asymmetry in the response of consumption and government spending.

[38]The hump-shape response in consumption and government spending can be attributed to agents' dislike for jumps in expenditure, underscored by two elements: (i) the role of habit formation in consumption and government spending and (ii) access to financial instruments that facilitate the smoothing of export price shocks.

Figure 6: Macroeconomic Responses to an Export Price Shock

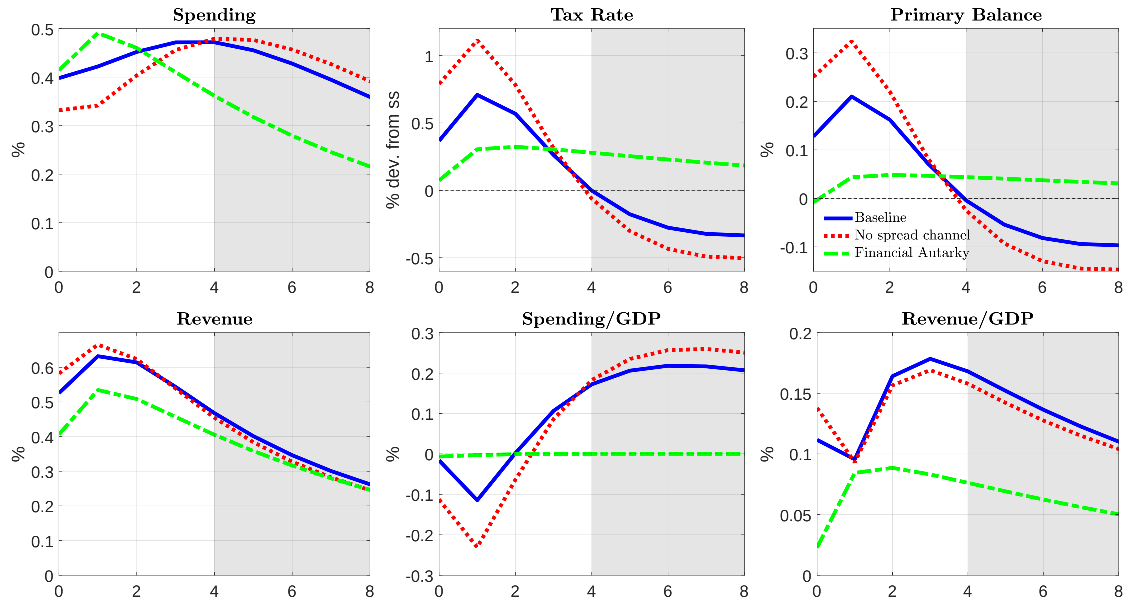


Notes: This Figure shows the impulse responses of the main macroeconomic aggregates to (one standard deviation) export price shock in the baseline model (in solid blue), the baseline model excluding the credit spread channel (in dotted red), and a model with financial autarky (in dashed-dotted green). The main macroeconomic aggregates plotted are observationally equivalent counterparts (expressed in constant prices). In the financial autarky case, instead of the spread we plot the interest rate. Responses are expressed in percentage deviations from steady state values, except for the trade balance-to-GDP ratio, which is shown in percentage. Gray areas denote the periods outside of the empirical horizon. Horizontal axes denote years.

Consequently, the Ramsey planner optimally implements a temporary tax increase to improve the primary balance. Figure 7 suggests a positive conditional comovement between output and fiscal variables (government spending, revenue, the primary balance, and the tax rate) after an export price shock. As a direct consequence of the "spread channel" and the overall reduction in debt, government borrowing costs fall. This partially counteracts the desire to save to smooth spending over time.

Figure 7 shows the initial transitory decline and subsequent rise in the government spending-to-GDP ratio, highlighting the preference for smoothing spending over time. The fiscal revenue-to-GDP ratio increases due to expanded consumption and a temporary rise in the consumption tax. This behavior of revenue and spending relative to GDP results in a persistent increase in the primary balance-to-GDP ratio, which gradually reverses in the long run. These dynamics suggest a procyclical government spending response and a countercyclical

Figure 7: Fiscal Responses to an Export Price Shock



Notes: This Figure shows the impulse responses of fiscal variables to (one standard deviation) export price shock in the baseline model (in solid blue), the model under financial autarky (in dashed-dotted green), and the baseline model with lower financial frictions (in dotted red). The main macroeconomic aggregates plotted are observationally equivalent counterparts (expressed in constant prices). Responses are expressed in percentage points, except for the tax rate, which is shown in percentage deviations from steady state values. Gray areas denote the periods outside of the empirical horizon. Horizontal axes denote years after the shock.

tax rate. Overall, the fiscal stance is associated with a primary surplus and, therefore, higher government savings following a commodity boom.

6.2 Disentangling the Channels

The fiscal policy response is driven by the interplay of three distinct mechanisms: the consumption preference channel (Fernández et al., 2021), the consumption smoothing channel (Riascos and Végh, 2003), and the credit spread channel (Drechsel and Tenreyro, 2018). The consumption preference channel captures households' valuation of consumption volatility, including both private and public consumption. The consumption smoothing channel reflects the ability and willingness of households and the Ramsey planner to smooth temporary shocks over time. In our framework, this channel assumes greater quantitative significance due to the foreign currency denomination of debt. The real exchange rate appreciation associated with improved terms of trade leads to a substantial reduction of existing debt, amplifying the effect. The credit spread channel represents the immediate impact of export price shocks on external borrowing costs.

To understand the role played by these mechanisms, we consider two alternative financial market arrangements: (i) financial autarky and (ii) incomplete markets without the credit spread channel. Financial autarky eliminates the smoothing and credit spread channels, allowing us to focus on the consumption preference channel. In this setting, the inability to borrow internationally constrains consumption and government spending smoothing.[39] Conversely, removing the credit spread channel emphasizes the consumption smoothing mechanism by increasing incentives to smooth shocks over time. The persistence of export price shocks plays a crucial role in determining the relative importance of these channels, with more persistent shocks limiting the scope for intertemporal smoothing.

Figures 6 and 7 depict in dashed green lines the responses to an export price shock under financial autarky. In this setting, the household's ability to buffer export price shocks is reduced, leading to more pronounced fluctuations in consumption and conditionally procyclical government spending. Since consumption increases, the tax base reacts more strongly on impact and the government can finance increased spending with a smaller tax increase. The tax rate therefore becomes less responsive to export price shocks and less countercyclical (consistent with the findings of Fernández et al., 2021). This is reflected in a smaller response of revenues and revenues as a share of GDP. Taken together, under financial autarky, we observe a more limited ability to run a primary surplus, as a consequence of the lack of access to international financial markets.

Alternatively, shutting down the credit spread channel helps us to highlight the significance of the consumption smoothing channel. In this case, illustrated in Figures 6 and 7 with dotted red lines, the fall in the interest rate is less pronounced so that both households and the benevolent planner have a further incentive to save on impact to better smooth out shocks (i.e., stronger savings). This arrangement leads to a more muted response of government spending and consumption. Compared to the baseline, consumption and government spending become less front-loaded, and the interest rate paid on debt becomes less responsive. As the tax base is smaller, the planner rises the tax rate by more to allow for higher savings. The tax rate's heightened responsiveness reflects the reduced cost of temporal resource allocation, prompting the planner to favor a stronger hike in the consumption tax rate.[40] Therefore, the tax policy is more countercyclical.

6.3 Robustness

This Section examines the sensitivity of conditional fiscal cyclicality to alternative calibrations of our baseline model. An increase in export prices is expansionary, and leads to a rise in the tax base and, consequently, an improvement in government revenues, irrespective of changes in the tax rate. The desire to smooth the fiscal bonanza implies an increase in government spending (for reasonable levels of persistence of export prices) and an improved primary balance. This results in a procyclical government spending response, while the overall fiscal stance (captured by the primary balance) correlates positively with the cycle, irrespective of parameter variations.[41] However, alternative model calibrations can potentially alter

[39]The full problem under financial autarky is detailed in Appendix F.3. To close the model in this setting, we introduce a quadratic portfolio adjustment cost, in line with Schmitt-Grohé and Uribe (2003).

[40]The planner must balance the distortionary effect of the tax rate on intra-period allocations (the wedge that it generates between the marginal product of labor and the marginal rate of substitution) against the intertemporal effects it has over time (by affecting the effective intertemporal price of consumption or the effective interest rate).

[41]The results remain qualitatively robust under two key modifications to the model: (i) the use of alternative distortionary taxes, such as payroll or income taxes, and (ii) the introduction of non-separable preferences

the relative responses of private and public consumption, as well as government revenues. Consequently, they may influence the cyclicality of the tax rate and fiscal ratios. Below, we present a summary of our key findings. Additional details are provided in Appendix G.

Consumption Preference vis-à-vis Consumption Smoothing Channel. The interaction between the consumption preference and consumption smoothing channels is crucial for understanding the fiscal response to commodity price shocks (see Fernández et al., 2021). We explore this interaction by varying the elasticities of intertemporal substitution of government spending (σ_g) relative to the one of aggregate consumption (σ_c), as well as the degree of financial market incompleteness, which in our setting is captured by the sensitivity of domestic spreads to (private and government) debt ($\psi = \psi_g$). When $\sigma_g > \sigma_c$, households' preferences imply a stronger willingness to smooth public over private consumption. As a result, private consumption increases proportionately more than public consumption following an increase in export prices. The relatively higher tax base allows the fiscal authority to reduce the tax rate while still financing higher public consumption, thus engaging in procyclical tax policy. In this setting, the increase in public spending (and revenue) is limited relative to the expansion in economic activity, leading to a countercyclical spending-to-GDP ratio (and revenue-to-GDP). These effects are amplified by greater asset market incompleteness (i.e., for large values of $\psi = \psi_g$), as spending decisions become more tightly linked to the government's ability to raise revenue through taxes.

Shock Persistence and Intertemporal Smoothing. The intertemporal consumption smoothing channel plays a key role in determining optimal Ramsey policy in our framework. This channel is fundamentally affected by two model features: (i) shock persistence; and (ii) a country's access to international financial markets, proxied by the debt-elastic premium. Transitory shocks generally lead to an improved fiscal stance, as reflected in a positive correlation between the primary balance and GDP. The cyclicality of tax policy closely mirrors that of the primary balance, becoming countercyclical for less persistent shocks. As the shock persistence decreases, government spending becomes less procyclical and can even turn acyclical, particularly when financial frictions are low. This is because less persistent shocks induce a more muted response in private consumption, leading to a weaker fiscal response. Importantly, higher financial frictions tend to amplify the procyclicality of government spending, as they limit the government's ability to smooth spending over time. This effect is particularly pronounced for less persistent shocks, where the need for consumption smoothing is higher.

Composition Channel. In a multi-good economy with imperfect substitution, the composition of government spending does not need to reflect the one of private consumption. Public services such as education, healthcare, national defense, and transportation predominantly exhibit non-tradable characteristics (as argued by Bianchi et al., 2023). Divergent preferences for non-tradable expenditures between private consumption and government spending influence the fiscal response, generating a *relative composition channel*.

When government spending is tilted towards non-tradable goods, its demand increases

between private consumption and government spending.

relative to the baseline, pushing up both the relative price of non-tradable goods and that of government spending. The opposite holds true for private consumption. The magnitude of this effect is proportional to the share of non-tradable goods in aggregate government spending (and in consumption), with a higher share of the former leading to a greater convergence between the dynamics of government spending's relative price and that of non-tradable goods (and vice versa). A higher relative price of public goods provision increases the overall level of spending, expressed in units of private consumption. Therefore, even when households' valuation towards consumption and government spending is the same, government spending increases by more than consumption. Thus, the tax rate must increase to fund this additional expenditure, which in turn depresses private consumption, weakening the base effect. The composition channel has the potential to accentuate the procyclicality of government spending, while also making the tax rate more countercyclical.

6.4 Model Correlations

In this Section, we concentrate on the implied correlation of fiscal variables as predicted by our theoretical model. Columns (1) and (2) of Table 4 present correlations conditional on export price shocks. When the model is only affected by commodity price shocks, we observe a pronounced procyclicality in spending and countercyclicality in tax rates. The overall fiscal stance, represented by the primary balance, displays a positive correlation with GDP. An analogous correlation pattern is present with respect to export prices. These correlations mirror the conditional correlations obtained in our empirical analysis, as reported in Table 3.

In our empirical analysis, we highlight the significant differences between unconditional and conditional correlations. These differences reflect the presence of multiple shocks, each exhibiting distinct conditional correlations. Indeed, optimal policy suggests that the response of fiscal instruments depends on the overall transmission of different shocks to the economy, resulting in correlations between these fiscal variables, and GDP and export prices that can vary significantly depending on the nature of the shocks. To illustrate this point, columns (3) and (4) extend the analysis to include a broader range of shocks–both foreign, such as shocks to export and import prices and global financing costs, and domestic, including total factor productivity and preference shocks.[42] Incorporating multiple shocks yields notable changes in the correlation patterns between fiscal variables, GDP, and export prices. Specifically, focusing on the baseline calibration (columns 3 and 4), the correlation between government spending and GDP decreases notably, alongside a transition to negative correlations for spending/GDP, revenue/GDP, and the tax rate with GDP.

Although the sign of the correlations with export prices remains unchanged, the magnitude experiences a marked decline. The numbers in brackets reproduce the same exercises but consider alternative shock combinations, achieved by randomly varying the standard

[42]Specifically, we introduce a stochastic TFP shock in each sector (equation E.15), a preference shock, specified as a shifter to the parameter ϕ, in equation (E.2), a shock to the foreign interest rate (equations E.9 and E.19), and a shock to the price of imported goods (equation E.6). The shock processes are calibrated as AR(1) processes, with parameters set according to existing literature. The reported results remain robust to reasonable variations in their calibration. Further details are provided in Appendix F.2.

Table 4: Model Correlations

	Export Price Shock		Multiple Shocks			
	Corr w/GDP (1)	Corr w/Px (2)	Corr w/GDP (3)		Corr w/Px (4)	
Spending	0.92	0.96	0.19	[-0.39, 0.94]	0.47	[0.04, 0.91]
Consumption Tax	0.30	0.18	-0.37	[-0.62, 0.32]	0.03	[0.02, 0.17]
Primary Balance	0.32	0.19	0.28	[0.07, 0.39]	0.15	[0.01, 0.20]
Revenue	0.99	0.99	0.24	[-0.36, 0.99]	0.50	[0.04, 0.95]
Spending/GDP	0.46	0.57	-0.34	[-0.66, 0.39]	0.13	[0.02, 0.54]
Revenue/GDP	0.90	0.95	-0.29	[-0.64, 0.78]	0.16	[0.03, 0.90]
Spread	-0.87	-0.91	-0.57	[-0.80, 0.09]	-0.82	[-0.88, -0.06]

Notes: For each variable listed, we report the correlations conditional on the export price shock (left panel) and multiple shocks (right panel). For the latter, we consider a combination of foreign shocks (shocks to export and import prices in foreign currency and shocks to global financing costs, r^*), and domestic shocks (both sectoral TFP shocks and preference shocks). The range of the minimum and maximum correlations for alternative model settings where we allow the volatility to each of the shocks to be in between 0 and twice the baseline volatility is shown in square brackets. Columns (1) and (3) measure the correlation with respect to detrended GDP, while columns (2) and (4) measure the correlation with respect to detrended export prices.

deviations of the various shocks within a range from 0 to twice the original value under the baseline calibration. We report the minimum and maximum correlations from 100 alternative random calibrations. These results highlight that the same optimal policy may lead to substantially varying patterns of unconditional correlations. Indeed, the correlation of many fiscal instruments with GDP can shift from positive to negative, depending on the relative importance of the shocks within the same theoretical model.

Since the same optimal policy results in different responses of fiscal instruments to underlying shocks, conditional correlations between fiscal variables and economic indicators are inherently shock-dependent. Therefore, a simple analysis of the procyclicality of fiscal instruments or indicators based on unconditional correlations may provide a misleading picture of the cyclicality of fiscal policy.

6.5 Optimal Fiscal Rules

While the Ramsey optimal policy provides a theoretical benchmark, its applicability in real-world settings remains unclear. Instead, many countries opt for fiscal rules as a more feasible approach to fiscal management (see, e.g., Bova et al., 2014). This raises several important questions: To what extent can simple, implementable rules approximate optimal policy outcomes? What specific prescriptions do these rules entail? And what are the welfare losses associated with using rules rather than fully optimal policies?

To address these questions, we evaluate the effectiveness of simple, implementable fiscal rules within our model framework. Our analysis centers on revenue-to-GDP ($\frac{\mathcal{R}_t}{\mathcal{Y}_t}$) and spending-to-GDP ratio ($\frac{\mathcal{G}_t}{\mathcal{Y}_t}$) rules:

$$\frac{\mathcal{R}_t}{\mathcal{Y}_t} \quad = \quad \frac{\mathcal{R}}{\mathcal{Y}} + \gamma_{r,y} \log\left(\frac{\mathcal{Y}_t}{\mathcal{Y}}\right) + \gamma_{r,x} \log\left(\frac{p_t^{x,\$}}{p^{x,\$}}\right) + \gamma_{r,d} \frac{q_t\left(d_t^g - d^g\right)}{\mathcal{Y}_t}, \tag{12}$$

Table 5: Welfare Evaluation Under Different Policy Regimes

	Policy Rule	$\gamma_{j,y}$	$\gamma_{j,x}$	$\gamma_{j,d}$	$\lambda \times 100$
Fiscal Rules	Revenue-to-GDP	0.041	0.011	0.003	0.009
	Spending-to-GDP	-0.235	0.049	0.142	

Notes: The table reports the optimal rule coefficients, $\gamma_{j,y}$, $\gamma_{j,x}$ and $\gamma_{j,d}$ for $j = \{g, r\}$. In all cases, the optimal rules restrict the policy parameters $\gamma_{j,y}$ and $\gamma_{j,x}$ to the interval $(-3, 3)$, and $\gamma_{j,d}$ in the interval $(0, 3)$. A positive coefficient $\gamma_{g,d}$ implies that government spending as a share to GDP falls when debt increases (see, eq. 13). The welfare costs, λ, are defined as the percentage drop in the Ramsey-optimal consumption process necessary to equate the level of welfare under the Ramsey policy and the alternative policy (calculated following Schmitt-Grohé and Uribe, 2007).

$$\frac{\mathcal{G}_t}{\mathcal{Y}_t} = \frac{\mathcal{G}}{\mathcal{Y}} + \gamma_{g,y} \log\left(\frac{\mathcal{Y}_t}{\mathcal{Y}}\right) + \gamma_{g,x} \log\left(\frac{p_t^{x,\$}}{p^{x,\$}}\right) - \gamma_{g,d} \frac{q_t\,(d_t^g - d^g)}{\mathcal{Y}_t}. \tag{13}$$

These rules allow for a cyclical adjustment in the fiscal stance with systematic responses to deviations in output ($\frac{\mathcal{Y}_t}{\mathcal{Y}}$) and export prices ($\frac{p_t^{x,\$}}{p^{x,\$}}$) from their steady-state levels, as well as for a response to the debt-to-GDP ratio. The magnitude and direction of these responses is given by the γ coefficients. We constrain the debt-to-GDP ratio coefficients to be positive, implying that spending decreases and revenues increase in response to rising debt levels. In addition, the debt components in the rules prescribe fiscal consolidation in response to increasing debt levels and help mitigate extreme debt volatility (see Heresi, 2024).

The rules coefficients are selected to maximize expected lifetime utility, while ensuring the local uniqueness of the rational expectations equilibrium. When considered in combination, these rules effectively define an implied rule for the primary balance-to-GDP ratio, consistent with the approach outlined by Kumhof and Laxton (2013). This framework allows us to systematically evaluate the performance of these fiscal rules against the optimal policy benchmark.

Our analysis reveals that the optimal fiscal rules closely align with the Ramsey optimal policy, resulting in minimal welfare losses.[43] As shown in Table 5, the welfare cost in consumption terms relative to the Ramsey optimal consumption process is 0.009%. The rules imply a procyclical spending-to-GDP response and a countercyclical response of the revenue-to-GDP ratio to an export price shock. Taken together, the rules also imply a countercyclical fiscal stance, with the primary balance-to-GDP ratio increasing following a rise in export prices. Therefore, our results reveal a deviation from balanced budget behavior, which generally characterize the fiscal stance of EMDEs (see, e.g., Ilzetzki and Végh, 2008). Moreover, the rules reflect a distinctly "prudent" approach, where the primary balance is required to improve as debt rises (in line with Mendoza and Ostry, 2008). Specifically, the rules suggest that this fiscal prudence is achieved primarily by reducing spending rather than by significantly increasing revenues (as a share of GDP).

By allowing the government to react differently to GDP and export price fluctuations, we find that the welfare maximizing fiscal rules imply a strong countercyclical response to GDP

[43] Appendix H reports additional details on the model under fiscal rules as well as the IRFs in this setting.

of $\mathcal{G}_t/\mathcal{Y}_t$ (and, implicitly, the primary balance over GDP, $\mathcal{PB}_t/\mathcal{Y}_t$), leaning against the wind of the domestic business cycle while maintaining a procyclical response to fluctuations in international commodity prices. These findings underscore an important insight: the coefficients in the policy rule alone do not allow for a clear indication of the cyclicality in a fiscal policy instrument or the overall stance.

6.6 Rationalizing the Role of Institutions

In Section 5.4, we established that countries with lower institutional quality exhibit quantitatively smaller reactions in fiscal variables to commodity price shocks. We also found that the government spending-to-GDP ratio decreases in countries with higher institutional quality while rising in those with weaker institutional frameworks. This Section explores whether these empirical observations can be rationalized by optimal policy within an alternative theoretical framework that more closely captures key structural features of countries with low levels of institutional quality.

To this end, we expand the theoretical model in two directions. First, we propose that economies with weaker institutions experience inefficiencies associated with the production of final output, possibly due to rent-seeking behavior.[44] Second, we assume that countries with lower institutional quality encounter more difficulties in accessing international financial markets. Therefore, when they do, they face more stringent borrowing conditions in the form of higher sensitivity of spreads to debt and no reaction of spreads to export prices. Otherwise, we continue to assume that the decisions regarding government spending, tax rates, and public debt are optimal, and institutional quality itself is treated as given (i.e., not modeled as endogenous). The model equations and solution are presented in Appendix I.

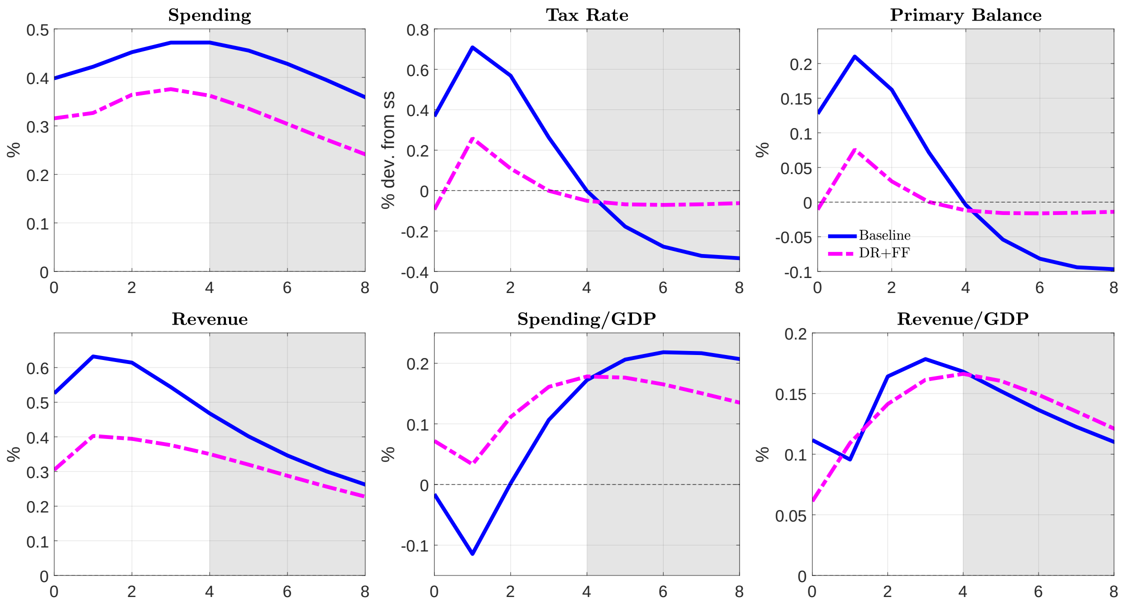
Figure 8 compares the responses of fiscal variables after an export price shock across two different model specifications: the previously described baseline (in blue), and a model integrating decreasing returns, heightened financial frictions, and no response of spreads to export price shocks (depicted in magenta). Qualitatively, we can associate the responses of the baseline economy with the empirical responses under high-quality institutions and the responses of the alternative model with the empirical responses under low-quality institutions.

In line with our empirical findings, both models show positive initial responses in government spending, the tax rate, revenues, and the primary balance, while differing in the initial response of spending as a share of GDP. Higher borrowing costs, driven by greater financial frictions, reduce the incentive to smooth consumption over time for countries with weaker institutional quality. As a result, these countries are less inclined to raise taxes or improve their fiscal position when a commodity price boom enhances their terms of trade.[45] Consequently, spending as a share of GDP increases alongside revenues, contrasting with

[44] We model production inefficiencies and the possibility of rent-seeking behavior in the economy by assuming that countries with lower institutional quality operate under a decreasing returns technology across production sectors. In this setting, positive profits are present in equilibrium, which can also be thought of as the accrual of rents in the presence of economic inefficiencies.

[45] This aligns with the results under financial autarky discussed in the previous section.

Figure 8: The Role of Institutions



Notes: This Figure shows the impulse responses of the fiscal variables to (a one standard deviation) export price shock in the baseline model (in solid blue); and a model with decreasing returns (DR), high financial frictions (FF), and without the credit spread channel (in magenta). Responses of the tax rate are expressed in percentage deviations from steady state values. Responses of the primary balance, spending, revenue, and fiscal ratios are in percentage points. Gray areas denote the periods outside of the empirical horizon. Vertical axes denote years.

the temporary decline observed in the baseline scenario. At the same time, the presence of rents limits the benefits to the broader economy, leading to lower government spending and revenues than in the baseline case. This framework highlights a potential challenge faced by countries with weaker institutions, as they gain less from positive external shocks, partly due to limited access to international financial markets.

7 Conclusion

This paper investigates how fiscal policy reacts to commodity price shocks in EMDEs. Given their significant impact on the terms of trade, these shocks are key drivers of the business cycles of EMDEs. We show that, contrary to the procyclical fiscal stance indicated by unconditional correlations and common literature perspectives, a conditional approach uncovers a different interaction between fiscal policies and commodity price shocks. Government spending is procyclical, while tax rates are countercyclical. However, the overall fiscal stance suggests that fiscal policy mitigates the transmission of external commodity price shocks. Additionally, the increase in tax rates during commodity booms, coupled with an expanded tax base, strengthens the primary balance–contrasting with the acyclical pattern implied by unconditional correlations. Our findings align with the optimal fiscal policy response predicted by a multi-good SOE model that incorporates export price shocks. The model successfully reproduces the conditional correlations following exogenous shifts in the terms of trade and reveals significantly different patterns of "unconditional" correlations when multiple shocks are introduced.

Additionally, we revisit the role that institutions play in shaping fiscal policy responses to commodity price shocks. Countries with weaker institutions display a more pronounced procyclical fiscal stance: spending increases more than proportionally with respect to GDP, the improvements in the primary balance are more muted as revenue does not benefit from any increase in the tax rates. These responses are consistent with optimal Ramsey policy in a framework where inefficiencies in production result in economic rents and where lower credibility limits the ability of accessing international markets and therefore produces higher sensitivity of borrowing costs to the current level of debt.

Our study underlines the critical importance of distinguishing between conditional and unconditional correlations when analyzing the (pro)cyclicality of fiscal policy. Relying on unconditional correlations overlooks how fiscal instruments and the overall policy stance respond differently to various structural shocks. Consequently, unconditional correlations are inadequate indicators of the appropriateness of fiscal policy.

References

ALESINA, A., F. R. CAMPANTE, AND G. TABELLINI (2008): "Why is Fiscal Policy Often Procyclical?" *Journal of the European Economic Association*, 6, 1006–1036.

AREZKI, R. AND M. BRÜCKNER (2012): "Commodity Windfalls, Democracy and External Debt," *Economic Journal*, 122, 848–866.

ARROYO MARIOLI, F. AND C. A. VÉGH (2023): "Fiscal Procyclicality in Commodity Exporting Countries: How Much Does it Pour and Why?" NBER Working Papers 31431, National Bureau of Economic Research, Inc.

AZZIMONTI, M. AND N. MITRA (2023): "Sovereign Default and Tax-smoothing in the Shadow of Corruption and Institutional Weakness," NBER Working Papers 31943, National Bureau of Economic Research, Inc.

BARRO, R. J. (1979): "On the Determination of the Public Debt," *Journal of Political Economy*, 87, 940–971.

BARTIK, T. J. (1991): *Who Benefits from State and Local Economic Development Policies?*, W.E. Upjohn Institute.

BIANCHI, J., P. OTTONELLO, AND I. PRESNO (2023): "Fiscal Stimulus under Sovereign Risk," *Journal of Political Economy*, 131, 2328–2369.

BLANCHARD, O. AND L. KATZ (1992): "Regional Evolutions," *Brookings Papers on Economic Activity*, 23, 1–76.

BORUSYAK, K., P. HULL, AND X. JARAVEL (2022): "Quasi-Experimental Shift-Share Research Designs," *The Review of Economic Studies*, 89, 181–213.

——— (2024): "A Practical Guide to Shift-Share Instruments," NBER Working Papers 33236, National Bureau of Economic Research, Inc.

BOVA, M. E., N. CARCENAC, AND M. M. GUERGUIL (2014): "Fiscal Rules and the Procyclicality of Fiscal Policy in the Developing World," IMF Working Papers 2014/122, International Monetary Fund.

CÉSPEDES, L. F. AND A. VELASCO (2014): "Was this Time Different?: Fiscal Policy in Commodity Republics," *Journal of Development Economics*, 106, 92–106.

CHARI, V. V. AND P. J. KEHOE (1993): "Sustainable Plans and Mutual Default," *The Review of Economic Studies*, 60, 175–195.

CLOYNE, J. (2013): "Discretionary Tax Changes and the Macroeconomy: New Narrative Evidence from the United Kingdom," *American Economic Review*, 103, 1507–28.

CLOYNE, J., O. JORDÀ, AND A. M. TAYLOR (2023): "State-Dependent Local Projections: Understanding Impulse Response Heterogeneity," NBER Working Papers 30971, National Bureau of Economic Research, Inc.

CUADRA, G., J. SANCHEZ, AND H. SAPRIZA (2010): "Fiscal Policy and Default Risk in Emerging Markets," *Review of Economic Dynamics*, 13, 452–469.

DAVID, A. AND D. LEIGH (2018): "A New Action-based Dataset of Fiscal Consolidation in Latin America and the Caribbean," IMF Working Papers 2018/094, International Monetary Fund.

DEGASPERI, R., S. HONG, AND G. RICCO (2020): "The Global Transmission of U.S. Monetary Policy," CEPR Discussion Papers 14533.

DEN HAAN, W. J. (2000): "The Comovement Between Output and Prices," *Journal of Monetary Economics*, 46, 3–30.

DI PACE, F., L. JUVENAL, AND I. PETRELLA (2024): "Terms-of-Trade Shocks are Not all Alike," *American Economic Journal: Macroeconomics*, Forthcoming.

DRECHSEL, T., M. MCLEAY, AND S. TENREYRO (2019): "Monetary Policy for Commodity Booms and Busts," *Proceedings - Economic Policy Symposium, Jackson Hole, Kansas City, Federal Reserve Bank*.

DRECHSEL, T. AND S. TENREYRO (2018): "Commodity Booms and Busts in Emerging Economies," *Journal of International Economics*, 112, 200–218.

FERNÁNDEZ, A., A. GONZÁLEZ, AND D. RODRÍGUEZ (2018): "Sharing a Ride on the Commodities Roller Coaster: Common Factors in Business Cycles of Emerging Economies," *Journal of International Economics*, 111, 99–121.

FERNÁNDEZ, A., D. GUZMAN, R. E. LAMA, AND C. A. VÉGH (2021): "Procyclical Fiscal Policy and Asset Market Incompleteness," NBER Working Papers 29149, National Bureau of Economic Research, Inc.

FERNÁNDEZ, A., S. SCHMITT-GROHÉ, AND M. URIBE (2017): "World Shocks, World Prices, and Business Cycles: An Empirical Investigation," *Journal of International Economics*, 108, 2–14.

FRANKEL, J. A., C. A. VÉGH, AND G. VULETIN (2013): "On Graduation from Fiscal Procyclicality," *Journal of Development Economics*, 100, 32–47.

GAVIN, M. AND R. PEROTTI (1997): "Fiscal Policy in Latin America," *NBER Macroeconomics Annual*, 12, 11–61.

GOLDSMITH-PINKHAM, P., I. SORKIN, AND H. SWIFT (2020): "Bartik Instruments: What, When, Why, and How," *The American Economic Review*, 110, pp. 2586–2624.

HALAC, M. AND P. YARED (2024): "A Theory of Fiscal Responsibility and Irresponsibility," *Journal of Political Economy*, Forthcoming.

HAMANN, F., J. C. MENDEZ-VIZCAINO, E. G. MENDOZA, AND P. RESTREPO-ECHAVARRIA (2023): "Natural Resources and Sovereign Risk in Emerging Economies: A Curse and a Blessing," Working Paper 31058, National Bureau of Economic Research.

HAMILTON, J. D. (2003): "What is an Oil Shock?" *Journal of Econometrics*, 113, 363–398.

HERESI, R. (2024): "From Macroeconomic Stability to Welfare: Optimizing Fiscal Rules in Commodity-Dependent Economies," Mimeo.

ILZETZKI, E. (2011): "Rent-Seeking Distortions and Fiscal Procyclicality," *Journal of Development Economics*, 96, 30–46.

ILZETZKI, E. AND C. A. VÉGH (2008): "Procyclical Fiscal Policy in Developing Countries: Truth or Fiction?" NBER Working Papers 14191, National Bureau of Economic Research, Inc.

IMBENS, G. W. AND J. D. ANGRIST (1994): "Identification and Estimation of Local Average Treatment Effects," *Econometrica*, 62, 467–475.

IMF (2009): *Export and Import Price Index Manual: Theory and Practice*, International Monetary Fund. https://www.imf.org/en/Publications/Manuals-Guides/Issues/2016/12/31/Export-and-Import-Price-Index-Manual-Theory-and-Practice-19587.

——— (2016): *How to Adjust to a Large Fall in Commodity Prices*, IMF How to Do Notes, International Monetary Fund.

——— (2024): *Guidelines for Fiscal Adjustment*, International Monetary Fund. https://www.imf.org/external/pubs/ft/pam/pam49/pam4902.htm.

JORDÀ, O. (2005): "Estimation and Inference of Impulse Responses by Local Projections," *American Economic Review*, 95, 161–182.

JORDÀ, , M. SCHULARICK, AND A. M. TAYLOR (2020): "The Effects of Quasi-Random Monetary Experiments," *Journal of Monetary Economics*, 112, 22–40.

JUVENAL, L. AND I. PETRELLA (2024): "Unveiling the Dance of Commodity Prices and the Global Financial Cycle," *Journal of International Economics*, 150, 103913.

KAMINSKY, G. L. (2010): "Terms of Trade Shocks and Fiscal Cycles," in *Inflation in an Era of Relative Price Shocks*, ed. by R. Fry, C. Jones, and C. Kent, Reserve Bank of Australia, RBA Annual Conference Volume.

KAMINSKY, G. L., C. M. REINHART, AND C. A. VÉGH (2004): "When It Rains, It Pours: Procyclical Capital Flows and Macroeconomic Policies," *NBER Macroeconomics Annual*, 19, 11–53.

KUMHOF, M. AND D. LAXTON (2013): "Simple Fiscal Policy Rules for Small Open Economies," *Journal of International Economics*, 91, 113–127.

LANE, P. R. AND G. M. MILESI-FERRETTI (2007): "The External Wealth of Nations Mark II: Revised and Extended Estimates of Foreign Assets and Liabilities, 1970-2004," *Journal of International Economics*, 73, 223–250.

LANE, P. R. AND A. TORNELL (1999): "The Voracity Effect," *American Economic Review*, 89, 22–46.

MENDOZA, E. G. (1995): "The Terms of Trade, the Real Exchange Rate, and Economic Fluctuations," *International Economic Review*, 36, 101–137.

MENDOZA, E. G. AND J. D. OSTRY (2008): "International evidence on fiscal solvency: Is fiscal policy 'responsible'?" *Journal of Monetary Economics*, 55, 1081–1093.

PANIZZA, U. AND D. JAIMOVICH (2007): "Procyclicality or Reverse Causality?" IDB Publications (Working Papers) 1949, Inter-American Development Bank.

RIASCOS, A. AND C. VÉGH (2003): "Procyclical Government Spending in Developing Countries: The Role of Capital Market Imperfections," Tech. rep., Working paper, University of Maryland.

RIGOBON, R. (2004): "Comment on: When It Rains, It Pours: Procyclical Capital Flows and Macroeconomic Policies," in *NBER Macroeconomics Annual 2004, Volume 19*, National Bureau of Economic Research, Inc, NBER Chapters, 80–82.

ROMER, C. D. AND D. H. ROMER (2010): "The Macroeconomic Effects of Tax Changes: Estimates Based on a New Measure of Fiscal Shocks," *American Economic Review*, 100, 763–801.

SCHMITT-GROHÉ, S. AND M. URIBE (2003): "Closing Small Open Economy Models," *Journal of International Economics*, 61, 163–185.

——— (2007): "Optimal Simple and Implementable Monetary and Fiscal Rules," *Journal of Monetary Economics*, 54, 1702–1725.

——— (2018): "How Important are Terms-of-Trade Shocks?" *International Economic Review*, 59, 85–111.

TALVI, E. AND C. A. VEGH (2005): "Tax Base Variability and Procyclical Fiscal Policy in Developing Countries," *Journal of Development Economics*, 78, 156–190.

VÉGH, C. A. AND G. VULETIN (2015): "How Is Tax Policy Conducted over the Business Cycle?" *American Economic Journal: Economic Policy*, 7, 327–70.

WORLD BANK (2024): *Fiscal Policy in Commodity Exporters: An Enduring Challenge*, Global Economic Prospects, The World Bank Group.

YARED, P. (2019): "Rising Government Debt: Causes and Solutions for a Decades-Old Trend," *Journal of Economic Perspectives*, 33, 115–140.

A Data

Our sample comprises a group of 54 emerging and low income countries. The split by income category is summarized in Table A.1. The data set includes information on output, capital flows, spreads, and export prices. The sources of data are described in section A.1. Tables A.2 and A.3 provide a comprehensive summary of the macro and spreads coverage. The fiscal data coverage is detailed in Table A.4. Section A.2 describes the criteria to select the countries into our sample.

A.1 Data Sources

Macro data:

- Real GDP in local currency units. Source: World Bank's World Development Indicators (WDI) database. Indicator code: NY.GDP.MKTP.KN

- Nominal GDP in local currency units. Source: World Bank's World Development Indicators (WDI) database. Indicator code: NY.GDP.MKTP.CN

- Spreads are obtained combining data based on:
  - EMBI. Description: JP Morgan EMBI global diversified index and JP Morgan EMBI global index, in bps. Sources: Datastream, Bloomberg, and JP Morgan.
  - Interest rate spreads. Description: domestic rate over U.S. rate (lending rate or t-bill), in %. Source: International Financial Statistics.

  Emerging market sovereign spreads are mainly derived from the J.P. Morgan EMBI global diversified index, which measures the spread over Treasuries. To expand coverage for certain countries, we supplemented this with the J.P. Morgan EMBI global index. Moreover, to further extend coverage, we also used interest rate spreads, which are calculated as the difference between the domestic t-bill and the U.S. t-bill. In cases where this data was not available, we used the domestic lending rate over the U.S. lending rate instead. A comprehensive breakdown of our calculations can be found in Table A.3.

- Export Prices. Export price index, 2010=100. Sources: Authors' calculation using as inputs data from MIT Observatory of Economic Complexity, World Bank, Federal Reserve Economic Data (FRED, Federal Reserve Bank of St. Louis).

Quality of Institutions:

- Institutional Quality Index refers to the political risk rating which ranges from 0 (highest risk) to 100 (lowest risk). It evaluates the political stability of a country on a comparable basis with other countries by assessing risk points in government stability, socioeconomic conditions, investment profile, internal conflict, external conflict, corruption, military in politics, religious tensions, law and order, ethnic tensions, democratic accountability, and bureaucracy quality. Source: Political risk rating sourced from the International Country Risk Guide (ICRG).

Fiscal data:

- Government Total Expenditure. Source: World Economic Outlook. Indicator code: GGX spliced with GGENL.
- Government Revenue. Source: World Economic Outlook. Indicator code: GGR spliced with GGRG.
- Primary Balance. Source: World Economic Outlook. Indicator code: GGXONLB spliced with GGBXI.
- Government Interest Expense. Source: World Economic Outlook. Indicator code: GGEI spliced with GCEI.
- Fiscal variables are calculated in terms of trend GDP and nominal GDP using nominal GDP in local currency units. Source: World Bank's World Development Indicators (WDI) database. Indicator code: NY.GDP.MKTP.CN

- VAT tax rate: VAT Tax rates are from Végh and Vuletin (2015). To expand the coverage, for the following countries we used data from the Inter-American Center of Tax Administrators: Bolivia (1990-1993), Brazil (2011-2019), Costa Rica (1990-2019), Dominican Republic (1990-1991), Guatemala (1990-2019), Panama (1990-2019).

A.2 Country Selection

We restrict the set of countries that we study to ensure the availability of data for the variables analyzed. First, we focus on emerging countries according to the definition of the IMF World Economic Outlook. Second, we drop large economies such as China and India. Finally, we drop economies which are classified as emerging but are part of the European Union such as Poland. After applying these filters, our sample consists of 54 emerging economies.

Table A.1: Country Coverage by Income Classification

Upper Middle Income	Lower Middle Income	Low Income
Argentina	Algeria	Cameroon
Armenia	Angola	Côte d'Ivoire
Azerbaijan	Bolivia	Ghana
Belarus	Egypt	Kenya
Belize	El Salvador	Mozambique
Brazil	Indonesia	Nigeria
Chile	Lebanon	Tanzania
Colombia	Mongolia	Vietnam
Costa Rica	Morocco	Zambia
Dominican Republic	Pakistan	
Ecuador	Philippines	
Gabon	Sri Lanka	
Georgia	Tunisia	
Guatemala	Ukraine	
Iraq		
Jamaica		
Kazakhstan		
Kuwait		
Malaysia		
Mexico		
Panama		
Peru		
Qatar		
Russia		
Serbia		
South Africa		
Thailand		
Türkiye		
Trinidad and Tobago		
Uruguay		
Venezuela		

Notes: This Table shows the country classification by income group. The low income classification is from the IMF while the breakdown between upper middle income and lower middle income is sourced from the World Bank.

Table A.2: Macro and Institutional Quality Data Coverage

Country	Real GDP	Nominal GDP	Spreads	Institutions	Export Prices
Algeria	1990-2019	1990-2019	1999-2019	1990-2019	1990-2019
Angola	1990-2019	1990-2019	2006-2019	1990-2019	1990-2019
Argentina	1990-2019	1990-2019	1993-2019	1990-2019	1990-2019
Armenia	1990-2019	1990-2019	2000-2019	1998-2019	1993-2019
Azerbaijan	1990-2019	1990-2019	1998-2019	1998-2019	1993-2019
Belarus	1990-2019	1990-2019	2004-2019	1998-2019	1993-2019
Belize	1990-2019	1990-2019	2007-2019	-	1990-2019
Bolivia	1990-2019	1990-2019	1997-2019	1990-2019	1990-2019
Brazil	1990-2019	1990-2019	1994-2019	1990-2019	1990-2019
Cameroon	1990-2019	1990-2019	1993-2019	1990-2019	1990-2019
Chile	1990-2019	1990-2019	1999-2019	1990-2019	1990-2019
Colombia	1990-2019	1990-2019	1997-2019	1990-2019	1990-2019
Costa Rica	1990-2019	1990-2019	2002-2019	1990-2019	1990-2019
Côte d'Ivoire	1990-2019	1990-2019	1998-2019	1990-2019	1990-2019
Dominican Republic	1990-2019	1990-2019	2001-2019	1990-2019	1990-2019
Ecuador	1990-2019	1990-2019	1995-2019	1990-2019	1990-2019
Egypt	1990-2019	1990-2019	2001-2019	1990-2019	1990-2019
El Salvador	1990-2019	1990-2019	2002-2019	1990-2019	1990-2019
Gabon	1990-2019	1990-2019	2007-2019	1990-2019	1990-2019
Georgia	1990-2019	1990-2019	2007-2019	-	1993-2019
Ghana	1990-2019	1990-2019	2007-2019	1990-2019	1990-2019
Guatemala	1990-2019	1990-2019	2002-2019	1990-2019	1990-2019
Indonesia	1990-2019	1990-2019	2004-2019	1990-2019	1990-2019
Iraq	1990-2019	1990-2019	2006-2019	1990-2019	1990-2019
Jamaica	1990-2019	1990-2019	2007-2019	1990-2019	1990-2019
Kazakhstan	1990-2019	1990-2019	2007-2019	1998-2019	1993-2019
Kenya	1990-2019	1990-2019	1993-2019	1990-2019	1990-2019
Kuwait	1995-2019	1990-2019	2004-2019	1990-2019	1993-2019
Lebanon	1990-2019	1990-2019	1998-2019	1990-2019	1990-2019
Malaysia	1990-2019	1990-2019	1996-2019	1990-2019	1990-2019
Mexico	1990-2019	1990-2019	1993-2019	1990-2019	1990-2019
Mongolia	1990-2019	1990-2019	2012-2019	1990-2019	1990-2019
Morocco	1990-2019	1990-2019	1997-2019	1990-2019	1990-2019
Mozambique	1990-2019	1991-2019	2000-2019	1990-2019	1990-2019
Nigeria	1990-2019	1990-2019	1993-2019	1990-2019	1990-2019
Pakistan	1990-2019	1990-2019	2001-2019	1990-2019	1990-2019
Panama	1990-2019	1990-2019	1996-2019	1990-2019	1990-2019
Peru	1990-2019	1990-2019	1997-2019	1990-2019	1990-2019
Philippines	1990-2019	1990-2019	1993-2019	1990-2019	1990-2019
Qatar	2000-2019	1990-2019	2001-2019	1990-2019	1990-2019
Russia	1990-2019	1990-2019	1997-2019	1990-2019	1993-2019
Serbia	1995-2019	1995-2019	2005-2019	1990-2019	2007-2019
South Africa	1990-2019	1990-2019	1994-2019	1990-2019	1990-2019
Sri Lanka	1990-2019	1990-2019	2007-2019	1990-2019	1990-2019
Tanzania	1990-2019	1990-2019	1993-2019	1990-2019	1990-2019
Thailand	1990-2019	1990-2019	1997-2005	1990-2019	1990-2019
Trinidad and Tobago	1990-2019	1990-2019	2007-2019	1990-2019	1990-2019
Tunisia	1990-2019	1990-2019	2002-2019	1990-2019	1990-2019
Türkiye	1990-2019	1990-2019	1996-2019	1990-2019	1990-2019
Ukraine	1990-2019	1990-2019	2000-2019	1998-2019	1993-2019
Uruguay	1990-2019	1990-2019	2001-2019	1990-2019	1990-2019
Venezuela	1990-2014	1990-2014	1993-2019	1990-2019	1990-2019
Vietnam	1990-2019	1990-2019	2005-2019	1990-2019	1990-2019
Zambia	1990-2019	1990-2019	1990-2019	1990-2019	1990-2019

Notes: This Table presents the macro and institutional quality data coverage for each country included in our sample.

Table A.3: Spreads Data

Country	Years	Notes
Algeria	1999-2019	1999-2002 uses EMBI GD index. Coverage extended by splicing using the African EMBI GD index for 2003-2019.
Angola	2006-2019	2012-2019 uses EMBI G and GD indices. Spliced backwards using interest rate spread based on domestic t-bill over U.S. t-bill for 2006-2011.
Argentina	1993-2019	EMBI G and EMGI GD indices
Armenia	2000-2019	2013-2019 uses EMBI GD index. Spliced backwards using interest rate spread on domestic lending rate over U.S. lending rate for 2000-2012
Azerbaijan	1998-2019	2012-2019 uses EMBI GD index. Spliced backwards using interest rate spread on domestic lending rate over U.S. lending rate for 1998-2011.
Belarus	2004-2019	2010-2019 uses EMBI GD index. Spliced backwards using interest rate spread on domestic lending rate over U.S. lending rate for 2004-2009.
Belize	2007-2019	EMBI GD index.
Bolivia	1997-2019	2012-2019 uses EMBI G and GD indices. Spliced backwards using interest rate spread on domestic lending rate over U.S. lending rate for 1997-2011.
Brazil	1994-2019	EMBI GD index.
Cameroon	1993-2019	2015-2019 uses EMBI G and GD indices. Spliced backwards using the Sub-Saharan Africa GD index for 2003-2014.
Chile	1999-2019	EMBI GD index.
Colombia	1997-2019	EMBI GD index.
Costa Rica	2002-2019	2012-2018 uses EMBI GD index. Spliced backwards using the CACI index for Costa Rica for 2002-2011.
Côte d'Ivoire	1998-2019	EMBI GD index.
Dominican Republic	2001-2019	EMBI G and EMGI GD indices
Ecuador	1995-2019	EMBI GD index.
Egypt	2001-2019	EMBI GD index.
El Salvador	2002-2019	EMBI GD index.
Gabon	2007-2019	EMBI G and EMGI GD indices
Georgia	2007-2019	2008-2019 uses EMBI GD index. Spliced backwards using interest rate spread on domestic lending rate over U.S. lending rate for 2007.
Ghana	2007-2019	EMBI GD index.
Guatemala	2002-2019	2012-2019 uses EMBI GD index. Spliced backwards using the CACI index for Guatemala for 2002-2011.
Indonesia	2004-2019	EMBI GD index.
Iraq	2006-2019	EMBI GD index.
Jamaica	2007-2019	EMBI GD index.
Kazakhstan	2007-2019	EMBI GD index.
Kenya	1993-2019	2014-2019 uses EMBI GD index. Spliced backwards using the Sub-Saharan Africa GD index for 2003-2013.
Kuwait	2004-2019	Due to lack of EMBI data corresponds to MECI spread.
Lebanon	1998-2019	EMBI GD index.
Malaysia	1996-2019	EMBI GD index.
Mexico	1993-2019	EMBI G and EMGI GD indices.
Mongolia	2012-2019	EMBI GD index.
Morocco	1997-2019	EMBI GD index.
Mozambique	2000-2019	2012-2014 uses EMBI G index. Spliced backwards using interest rate spread based on domestic t-bill over U.S. t-bill for 2000-2011.
Nigeria	1993-2019	EMBI GD index.

to be continued in the next page . . .

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Country	Years	Notes
Pakistan	2001-2019	EMBI GD index.
Panama	1996-2019	EMBI GD index.
Peru	1997-2019	EMBI GD index.
Philippines	1993-2019	EMBI GD index.
Qatar	2001-2019	Due to lack of EMBI data corresponds to MECI spread.
Russia	1997-2019	EMBI G and EMGI GD indices
Serbia	2005-2019	EMBI GD index.
South Africa	1994-2019	EMBI G and EMGI GD indices
Sri Lanka	2007-2019	EMBI G and EMGI GD indices
Tanzania	1993-2019	2013-2019 uses EMBI G and GD indices. Spliced backwards using interest rate spread based on domestic t-bill over U.S. t-bill for 1993-2011.
Thailand	1997-2005	EMBI GD index.
Trinidad and Tobago	2007-2019	EMBI GD index.
Tunisia	2002-2019	EMBI GD index.
Türkiye	1996-2019	EMBI GD index.
Ukraine	2000-2019	EMBI GD index.
Uruguay	2001-2019	EMBI GD index.
Venezuela	1993-2019	EMBI G and EMGI GD indices
Vietnam	2005-2019	EMBI G and EMGI GD indices
Zambia	1990-2019	2012-2014 uses EMBI GD index. Spliced backwards using interest rate spread based on domestic t-bill over U.S. t-bill for 1990-2011.

Notes: This Table displays the coverage for the country spreads data along with the specific indices used for each country.

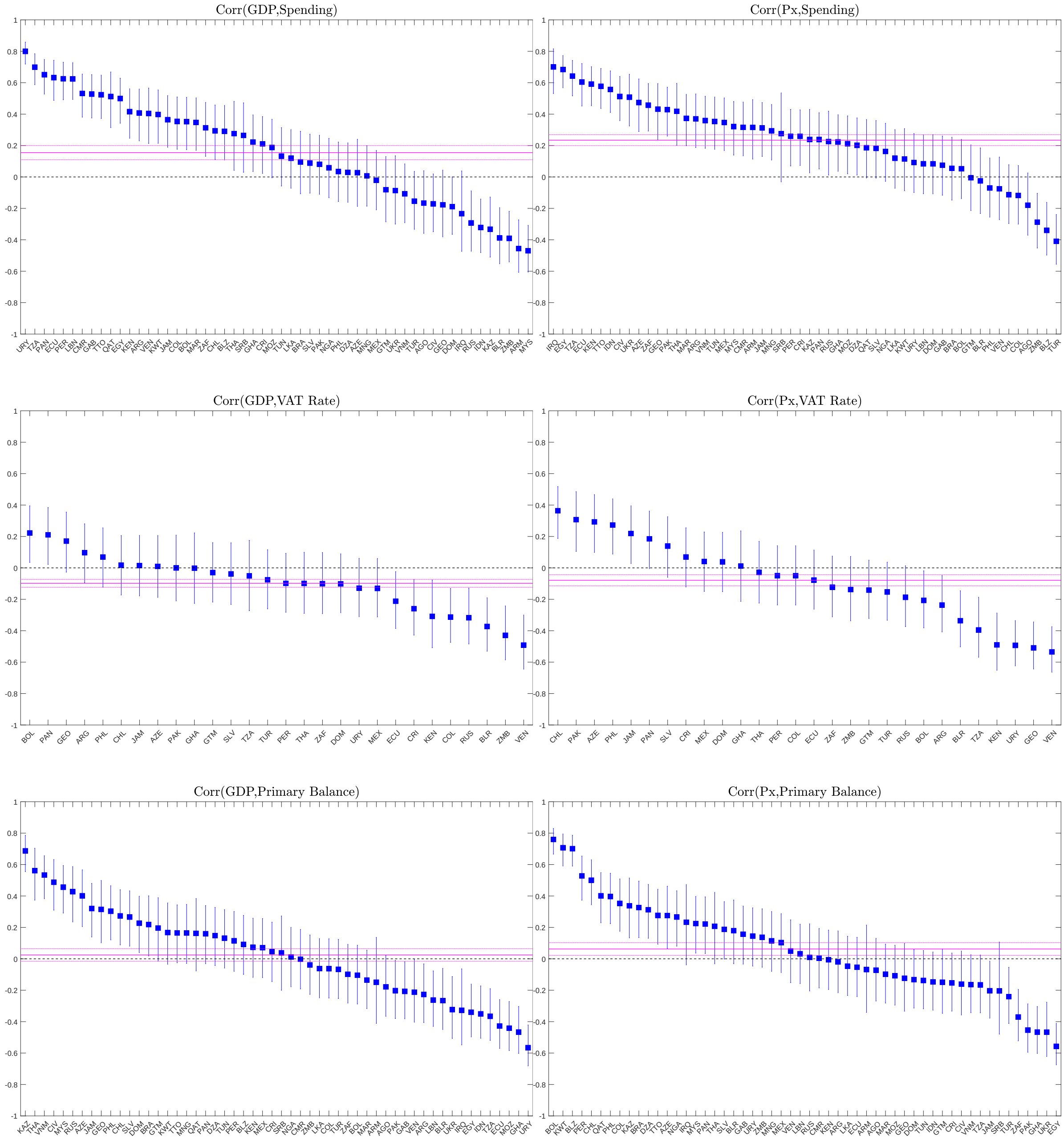
Table A.4: Fiscal Data Coverage

Country	Expenditure	Revenue	Primary Balance	Interest Expenditure	VAT rate
Algeria	1990-2019	1990-2019	1990-2019	1990-2019	-
Angola	1990-2019	1990-2019	1990-2019	1991-2019	-
Argentina	1990-2019	1990-2019	1992-2019	1992-2019	1990-2019
Armenia	1990-2019	1990-2019	2005-2019	1993-2019	-
Azerbaijan	1994-2019	1994-2019	1994-2019	1994-2019	1992-2019
Belarus	1990-2019	1990-2019	1993-2019	1993-2019	-
Belize	1990-2019	1990-2019	1990-2019	1990-2019	-
Bolivia	1990-2019	1990-2019	1990-2019	1990-2019	1990-2019
Brazil	1990-2019	1990-2019	1990-2019	1990-2019	2011-2019
Cameroon	1990-2019	1990-2019	1990-2019	1990-2019	-
Chile	1990-2019	1990-2019	1990-2019	1990-2019	1990-2019
Colombia	1990-2019	1990-2019	1990-2019	1990-2019	1990-2019
Costa Rica	1990-2019	1990-2019	1990-2019	1990-2019	1990-2019
Côte d'Ivoire	1990-2019	1990-2019	1995-2019	1990-2019	-
Dominican Republic	1990-2019	1990-2019	1990-2019	1990-2019	1991-2019
Ecuador	1990-2019	1990-2019	1990-2019	1995-2019	1990-2019
Egypt	1990-2019	1990-2019	1990-2019	1990-2019	-
El Salvador	1990-2019	1990-2019	1990-2019	1990-2019	1992-2019
Gabon	1990-2019	1990-2019	1990-2019	1990-2019	1995-2019
Georgia	1990-2019	1990-2019	1994-2019	1994-2019	1992-2019
Ghana	1990-2019	1990-2019	1990-2019	1990-2019	1998-2019
Guatemala	1995-2019	1995-2019	1995-2019	1990-2019	1990-2019
Indonesia	1990-2019	1990-2019	1990-2019	1990-2019	-
Iraq	2004-2019	2004-2019	2004-2019	2004-2019	-
Jamaica	1990-2019	1990-2019	1990-2019	1990-2019	1991-2019
Kazakhstan	1990-2019	1990-2019	1994-2019	1995-2019	-
Kenya	1990-2019	1990-2019	1990-2019	1990-2019	2000-2019
Kuwait	1990-2019	1990-2019	1990-2019	1990-2019	1990-2019
Lebanon	1990-2019	1990-2019	1990-2019	1990-2019	-
Malaysia	1990-2019	1990-2019	1990-2019	1990-2019	-
Mexico	1990-2019	1990-2019	1990-2019	1990-2019	1990-2019
Mongolia	1990-2019	1990-2019	1991-2019	1991-2019	-
Morocco	1990-2019	1990-2019	1990-2019	1990-2019	-
Mozambique	1990-2019	1990-2019	1990-2019	1990-2019	-
Nigeria	1990-2019	1990-2019	1990-2019	1990-2019	1994-2019
Pakistan	1990-2019	1990-2019	1991-2019	1990-2019	1995-2019
Panama	1990-2019	1990-2019	1990-2019	1990-2019	1990-2019
Peru	1990-2019	1990-2019	1991-2019	1991-2019	1990-2019
Philippines	1990-2019	1990-2019	1990-2019	1990-2019	1990-2019
Qatar	1990-2019	1990-2019	1990-2019	1991-2019	1990-2019
Russia	1990-2019	1990-2019	1996-2019	1992-2019	1992-2019
Serbia	2000-2019	2000-2019	2000-2019	2000-2019	-
South Africa	1990-2019	1990-2019	1990-2019	1990-2019	1992-2019
Sri Lanka	1990-2019	1990-2019	1990-2019	1990-2019	-
Tanzania	1990-2019	1990-2019	1990-2019	1990-2019	1998-2019
Thailand	1990-2019	1990-2019	2000-2019	2000-2019	1992-2019
Trinidad and Tobago	1990-2019	1990-2019	1990-2019	1990-2019	-
Tunisia	1990-2019	1990-2019	1990-2019	1990-2019	-
Türkiye	1990-2019	1990-2019	1990-2019	1990-2019	1990-2019
Ukraine	1990-2019	1990-2019	1994-2019	1995-2019	-
Uruguay	1990-2019	1990-2019	1990-2019	1990-2019	1990-2019
Venezuela	1990-2016	1990-2016	1990-2016	1990-2016	1993-2019
Vietnam	1990-2019	1990-2019	1990-2019	1990-2019	-
Zambia	1990-2019	1990-2019	1990-2019	1990-2019	1995-2019

Notes: This Table presents the fiscal data coverage for each country included in our sample.

B Additional Stylized Facts

Figure B.1: Unconditional Correlations



Notes: Each chart in this Figure displays the unconditional correlations between GDP (left panels) and government spending, the VAT rate, and the primary balance; and between export prices (Px, right panels) and the same set of fiscal variables. For each country, we present the estimated correlation coefficients (blue markers) along with their corresponding 68% confidence intervals (blue lines). The countries are arranged according to the magnitude of their correlation coefficients. The horizontal magenta lines indicate the overall estimated mean of these correlation coefficients, together with their 68% confidence intervals.

Table B.1: Unconditional Correlations by Institutional Quality

	Corr w/ GDP				Corr w/ Px			
	Low QI		High QI		Low QI		High QI	
Spending	0.094	[0.086]	**0.262**	[0.101]	**0.231**	[0.091]	**0.254**	[0.046]
VAT Rate	**-0.219**	[0.062]	-0.046	[0.045]	**-0.218**	[0.108]	0.064	[0.082]
Primary Balance	-0.073	[0.063]	**0.149**	[0.082]	-0.018	[0.059]	**0.179**	[0.076]
Revenue	0.057	[0.087]	**0.427**	[0.083]	**0.215**	[0.088]	**0.488**	[0.064]
Spending/GDP	-0.019	[0.076]	0.022	[0.082]	**0.256**	[0.060]	0.079	[0.080]
Revenue/GDP	-0.059	[0.076]	**0.141**	[0.085]	**0.245**	[0.056]	**0.363**	[0.055]
EMBI Spread	**-0.218**	[0.067]	-0.042	[0.110]	**-0.163**	[0.095]	-0.003	[0.100]

Notes: For each variable listed, we report the average (unconditional) correlation and its associated standard error (in square brackets). The left column measures correlation with respect to detrended GDP, whereas the right column measures correlation with respect to detrended export prices. Values in bold denote that the average correlation is significant at a 10% level. "Low QI" refers to countries in the bottom quartile of the institutional quality indicator distribution. Similarly, "High QI" refers to countries in the top quartile.

C Commodity Events

This appendix summarizes the methodology adopted to identify events tied to substantial commodity price fluctuations, which we use in building the instrument for export prices. Our approach involved examining historical documents, reports, and newspaper articles to pinpoint significant commodity price shifts, independent of global economic conditions. Following this, we classified each event into positive or negative price shocks, contingent on the price change trajectory. This categorization eventually influences the characterization of a country's export price shock as positive or negative, contingent on its role as an exporter of the particular commodity in question.

The series were constructed by using a number of sources: Food and Agriculture Organization (FAO) reports, publications from the International Monetary Fund (IMF) and the World Bank (WB), newspaper articles, academic papers and a number of online sources. In order to establish some rules at the time of selecting the dates, we followed the criteria listed below.

1. The event has to be important enough to affect a commodity market at a global level. Examples of these are natural disasters or weather related shocks in key areas where the commodity is produced, major geopolitical events, and unanticipated news on the volume of global production or demand of commodities.
2. The event should have an unambiguous effect on the price of the commodity.
3. The event has to be unrelated to important macroeconomic developments such as the global financial crisis or a U.S. recession. This aims at eliminating endogenous responses of commodity prices to the state of the economy.

By using this criteria we were able to identify 24 episodes of exogenous commodity price shocks that are unrelated to business cycle fluctuations. Of these events, 16 are favorable commodity price shocks and 8 are negative price shocks.

Figure C.1 summarizes the number of selected events, (i.e., the instances where $z_{it} \neq 0$), for each year and country in our sample.

C.1 Agriculture: Food and Beverage Commodities

i. Coffee

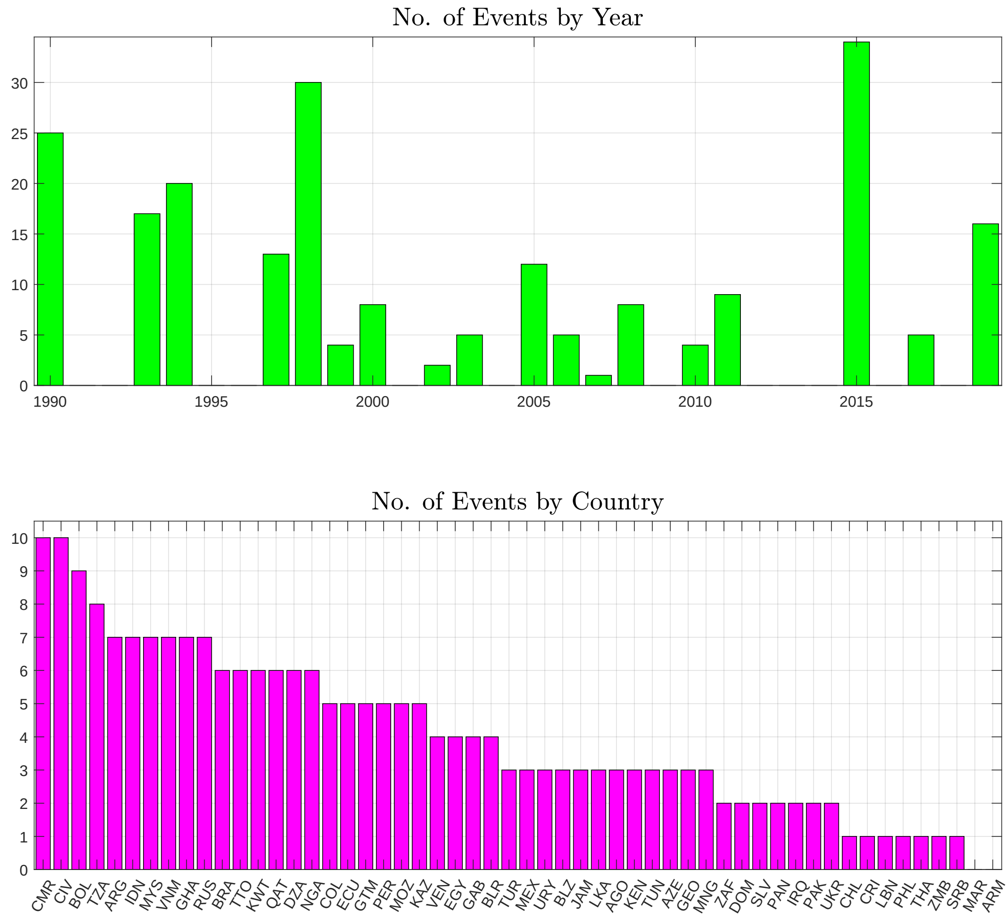
Year of Event: 1994.
Type of Event: Positive price shock.

According to a report from the International Coffee Organization (ICO), climate shocks which affected coffee prices were recorded in Brazil in 1994.[1] Our data are in line with this observation given that we observe that Arabica coffee prices increased from 1.56 dollars per kilo in 1993 to 3.31 in 1994.

Newspaper Articles. A newspaper article from the New York Times documents that the

[1] Report available at: `http://www.ico.org/news/icc-111-5-r1e-world-coffee-outlook.pdf`.

Figure C.1: Summary of Events Coverage



Notes: Count of instances where $z_{it} \neq 0$, for each year (top panel) and country (bottom panel) in our sample.

climate shock of 1994 in Brazil is related to a frost. Some important aspects of the article are quoted in what follows.

New Frost Hits Brazilian Coffee, New York Times (July 11, 1994):[2]

"Frost struck in Brazil's biggest coffee-growing state early today, and farmers said the effects were harsher than a freeze that hit two weeks ago."

"(...)Coffee prices soared after the previous cold snap late last month, which destroyed one-third of next year's crop. Brazil is the largest coffee producer, accounting for about a quarter of world production. A threat to its crop can drastically affect world coffee prices(...)."

ii. Cereal

Year of Event: 1997.

Type of Event: Negative price shock.

As documented in De Winne and Peersman (2016), in 1996 the FAO issued a favorable fore-

[2] Article available at: `https://www.nytimes.com/1994/07/11/business/new-frost-hits-brazil-coffee.html`.

cast for world 1996 cereal output.[3] The largest increase was expected in coarse grains output, mostly in developed countries. Overall, global cereal production increased by 7.8 percent that year and this translated into lower prices. Our data show that the cereal price index experienced a sharp reduction from 1996 to 1997, going from 83.61 to 64.76.

Year of Event: 2010.
Type of Event: Positive price shock.

De Winne and Peersman (2016) report that cereal output was seriously affected by adverse weather conditions in key producing countries in Europe. A group of countries that includes the Russian Federation, Kazakhstan and Ukraine suffered from a heatwave and droughts while the Republic of Moldova had floods. According to a report from the FAO, "International prices of grain have surged since the beginning of July in response to drought-reduced crops in CIS exporting countries and a subsequent decision by the Russian Federation to ban exports."[4]

iii. Cocoa

Year of Event: 1999.
Type of Event: Negative price shock.
According to a report from FAO, the drop in cocoa prices during 1999 was primarily attributed to a surplus in supply resulting from a rise in production levels across major cocoa-producing nations.

Newspaper Articles. An article from the New York Times documents the cocoa price decline in 1999.
The Market: Commodities, New York Times (November 3, 1999):[5]

"COCOA FALLS. Cocoa fell as shippers in the Ivory Coast, the world's largest supplier, begin exporting newly harvested beans at a time of weak demand. In New York, cocoa for December delivery fell $38, to $840 a metric ton."

Year of Event: 2002.
Type of Event: Positive price shock.

According to a report from the International Cocoa Organization, the increase in cocoa prices in 2002 was largely due to an attempted coup on 19th September in Côte d'Ivoire, which is the leading cocoa producing country. Uncertainty over potential disruptions emanating from the sociopolitical crisis and civil war pushed prices to a 16-year high at 2.44 dollars per tonne in October 2002.[6] Our data show that between 2001 and 2002 cocoa prices increased from

[3]The FAO document is available at: http://www.fao.org/docrep/004/w1690e/w1690e02.htm#I2.
[4]Available at: http://www.fao.org/docrep/012/ak354e/ak354e00.pdf.
[5]Article available at: https://www.nytimes.com/1999/11/03/business/the-markets-commodities.html?searchResultPosition=24.
[6]https://www.icco.org/about-us/international-cocoa-agreements/cat_view/30-related-documents/45-statistics-other-statistics.html.

1.07 dollars per kilo to 1.78 dollars per kilo.

Year of Event: 2017.
Type of Event: Positive price shock.

According to a report from the International Cocoa Organization, the decline in cocoa prices in 2017 was driven by favorable weather conditions in major producing countries such as Côte d'Ivoire and Ghana.[7] Our data show that cocoa prices declined around 30 percent in 2017.

Newspaper Articles. A newspaper article from the New York Times documents the cocoa price increase originated in Cote d'Ivore in 2002. Some important aspects of the article are quoted below.
War Inflates Cocoa Prices But Leaves Africans Poor, New York Times (October 31, 2002):[8]
"As civil war raged in Ivory Coast, the world's biggest cocoa producer, speculative traders here and in New York sent prices this month to 17-year highs."

iv. Rice

Year of Event: 2008.
Type of Event: Positive price shock.

In 2008 rice prices nearly doubled. A report from the United States Department of Agriculture explains that the price increase was driven by trade restrictions of major suppliers.[9]

v. Sugar

Year of Event: 2006.
Type of Event: Positive price shock.

The sugar price increase in 2006 was caused by severe draughts in Thailand, the second largest sugar producer.[10]

vi. Soybean

Year of Event: 2008.
Type of Event: Positive price shock.

A report from the U.S. Bureau of Labor Statistics highlights that the high soybean prices in

[7] https://www.icco.org/wp-content/uploads/2019/07/ICCO-Monthly-Cocoa-Market-Review-February-2017.pdf.

[8] Article available at: https://www.nytimes.com/2002/10/31/business/war-inflates-cocoa-prices-but-leaves-africans-poor.html.

[9] https://www.ers.usda.gov/webdocs/outlooks/38489/13518_rcs09d01_1_.pdf?v=242#:~:text=Global\%20rice\%20prices\%20increased\%20nearly,through\%20the\%20spring\%20of\%202008.

[10] see https://www.aa.com.tr/en/politics/thailand-facing-its-worst-drought-in-20-years-/552381.

2008 originated in the expectation of a reduction in supply.[11] We observe an increase of 40 percent in soybean prices in our data.

C.2 Agriculture: Raw Materials

i. Cotton

Year of Event: 1994.
Type of Event: Positive price shock.

A report from the U.S. International Trade Commission describes that the 1994 cotton price increase was driven by a decline in production in key production areas such as China, and India.[12] The decline in production in China is explained by bad weather and a bollworm infestation. A study from the National Cotton Council of America explains that the price increase is also partly due to a recovery in world cotton consumption following the stagnation that resulted from the dissolution of the Soviet Union in the early 1990s.[13] Our data indicate that cotton prices declined from 1.28 dollars per kilo in 1993 to 1.76 dollars per kilo in 1994.

Year of Event: 2003.
Type of Event: Positive price shock.

MacDonald and Meyer (2018) analyze the challenges faced when forecasting cotton prices in the long run. The article highlights that in 2003 there was a severe weather damage to cotton crops in China which resulted in a surge in cotton prices. In addition, an article from the National Cotton Council of America highlights that in the 2003 season, "(...) USDA's forecast put world sticks at their lowest level since 1994/95, raising the specter of a world cotton shortage for the first time in nearly a decade."[14] Our data show that cotton prices increased from 1.02 dollars per kilo in 2002 to 1.40 dollars per kilo in 2003.

Year of Event: 2010.
Type of Event: Positive price shock.

Janzen et al. (2018) analyze the extent to which cotton price movements can be attributed to comovement with other commodities vis-à-vis cotton specific developments. They point at the fact that in 2010-2011 cotton was scarce as a consequence of a negative supply shock generated by lower than average planted crops and negative weather shocks in the USA and Pakistan. This led to an increase in the price of cotton. The authors explain that this boom-

[11] https://www.bls.gov/opub/btn/volume-9/a-historical-look-at-soybean-price-increases-what-happened-since-the-year-2000.htm

[12] Article available at: https://books.google.com/books?id=OZFDf6qLEosC&pg=SA3-PA5&lpg=SA3-PA5&dq=cotton+prices+1994&source=bl&ots=vi6JuOeGer&sig=DX9iSSIDP__dPIGTNKEfB03FkSA&hl=en&sa=X&ved=2ahUKEwiJkOOWztneAhVkneAKHWFOCWs4ChDoATADegQIBRAB#v=onepage&q=cotton\%20prices\%201994&f=false.

[13] Article available at: https://www.cotton.org/issues/2005/upload/WorldCottonMarket.pdf.

[14] Article available at: https://www.cotton.org/issues/2005/upload/WorldCottonMarket.pdf.

bust appears to be cotton-specific, unlike other cases in which a set of macroeconomic factors drive the price of a broad range of commodities. Our data confirm the findings of the paper. In fact, cotton prices increased from 1.38 dollars per kilo in 2009 to 2.28 dollars per kilo in 2010.

ii. Timber

Year of Event: 1993.
Type of Event: Positive price shock.

Sohngen and Hayne (1994) explain that the 1993 price spike was driven by the environmentally friendly policies that President Clinton issued to protect forests which limited the timber harvests.[15] The application of such policies is confirmed in the list of environmental actions taken by President Clinton and Vice President Al Gore and is documented in the White House Archives.[16] Our data reveal that the timber price index increased from 72 in 1992 to 101 in 1993.

Newspaper Articles. A newspaper article from the Washington Post documents this episode and describes how the environmental policy was viewed as a threat to the woods product industry.
Clinton to Slash Logging (July 2, 1993):[17]

"To protect the region's wildlife and old-growth forests, the administration plan will allow for average timber harvests over the next decade of 1.2 billion board feet per year. That is about half the level of the last two years, and only a third of the average rate between 1980 and 1992, when annual harvests swelled as high as 5.2 billion board feet."

iii. Tobacco

Year of Event: 1993.
Type of Event: Negative price shock.

A report from the FAO highlights that the worldwide increase in competition for exports in 1993 led to a substantial fall in tobacco prices.[18] Our data reveal that tobacco prices declined 22 percent between 1992 and 1993.

iv. Rubber

Year of Event: 2010.
Type of Event: Positive price shock.

In 2010 rubber prices almost doubled in 2010. This is due to severe draughts in Thailand

[15]Article available at: `https://www.fs.fed.us/pnw/pubs/pnw_rp476.pdf`.
[16]Available here `https://clintonwhitehouse4.archives.gov/CEQ/earthday/ch13.html`.
[17]`https://www.washingtonpost.com/archive/politics/1993/07/02/clinton-to-slash-logging/f2266e63-f45f-4f88-bd1f-5f1a1edd820f/`
[18]Commodity Review and Outlook 1993-1994, Food and Agriculture Organization of the United Nations, page 156.

and India, major rubber producers.

Newspaper Articles. A newspaper article from the Financial Times documents this.
Rubber price breaks 58-year record (March 31, 2010):[19]

"The price surge comes on the back of the worst drought in north Thailand in a decade, which meteorologists blame on the lingering impact of the El Niño weather phenomenon. Drought has also hit India, the world's fourth-largest producer."

C.3 Energy Commodities

i. Combined Energy Commodities

Year of Event: 2015.
Type of Event: Negative price shock.
The booming U.S. shale oil production played a significant role in the oil price plummet in 2015. However, this event has affected the prices of the main fossil fuels commodities. Our data shows that crude oil prices declined 47 percent, while coal and natural gas prices contracted 16 and 26 percent, respectively, between 2014 and 2015.

Year of Event: 2019.
Type of Event: Negative price shock.

This is the first time that the United States became a net energy exporter following the development of shale technology (EIA, 2020). Therefore, this event can be understood as an event affecting crude oil, natural gas, and coal prices. However, it is not visible in crude oil price because there were attacks to Saudi Arabia oil facilities which disrupted oil exports (World Bank, 2021). This effect partially offset the price reduction from shale technology in the United States. In our dataset we observe that natural gas prices declined 25 percent in 2019 while coal declined 15 percent.

Newspaper Articles. A newspaper article explains the dimension of the oil price plunge.
How the U.S. and OPEC Drive Oil Prices, New York Times (October 5, 2015)[20]

"The global price of a barrel of oil remains near its lowest point since the depths of the 2009 recession–a result of a supply glut and battle for market share between the OPEC oil cartel and the United States, which has shifted toward the role of global swing producer."

iii. Crude Oil

Year of Event: 1998.
Type of Event: Negative price shock.

[19] https://www.ft.com/content/636c534c-3ce1-11df-bbcf-00144feabdc0
[20] https://www.nytimes.com/interactive/2015/09/30/business/how-the-us-and-opec-drive-oil-prices.html?searchResultPosition=28.

Känzig (2021) highlights the role played by oil supply expectations in driving the plunge in oil prices in 1998. Our dataset indicate that oil prices declined 32 percent in 1998.

iii. Natural Gas

Year of Event: 2000.
Type of Event: Positive price shock.

The Energy Information Administration (EIA) documents the California energy crisis of 2000-2001.[21] In terms of natural gas, a report from the Task Force on Natural Gas Market Stability finds that "the 2000-2001 California natural gas crisis resulted in major part from a perfect storm of sudden demand increase, impaired physical capacity, natural gas diversion, and inadequate storage fill. The quick summary is as follows: Low hydroelectric availability in 2000, coupled with a modest increase in overall power needs resulted in a substantial increase in gas-fired generation usage, with little preparation."[22] A study from the Federal Reserve Bank of San Francisco documents the natural gas price increase in 2000.[23] Our data show that the natural gas price index jumped from 39.78 in 1999 to 73.85 in 2000.

Year of Event: 2005.
Type of Event: Positive price shock.

An article from the "Oil and Gas Journal" highlights that the effects of Hurricanes Katrina and Rita were the main source of the price increase. Some details of the article are quoted below.[24]

"The combined effects of the 2004 and 2005 hurricane seasons had an impact across all sectors of the US gas industry. Hurricane Ivan, which made landfall in September 2004, caused more long-term gas production interruptions than any previous hurricane, but its impacts were dwarfed by Hurricanes Katrina (landfall Aug. 29, 2005) and Rita (Sept. 24, 2005). The combined effects of Hurricanes Katrina and Rita were by far the most damaging in the history of the US petroleum industry."

A report from the Federal Energy Regulatory Commission highlights the following:[25]

"The pump was primed for significant energy price effects well before Hurricanes Katrina and Rita hit the Gulf Coast production areas in September. The Gulf storms exacerbated already tight supply and demand conditions, increasing prices for fuels in the United States further after steady upward pressure on prices throughout the summer of 2005. Most of

[21] https://www.eia.gov/electricity/policies/legislation/california/subsequentevents.html.

[22] http://bipartisanpolicy.org/wp-content/uploads/sites/default/files/Introduction\%20to\%20North\%20American\%20Natural\%20Gas\%20Markets_0.pdf.

[23] https://www.frbsf.org/economic-research/publications/economic-letter/2001/february/economic-impact-of-rising-natural-gas-prices/#subhead3.

[24] https://www.ogj.com/articles/print/volume-104/issue-36/general-interest/us-gas-market-responds-to-hurricane-disruptions.html.

[25] https://www.ferc.gov/EventCalendar/Files/20051020121515-Gaspricereport.pdf.

this was due to increased electric generation demand for natural gas caused by years of investment in gas-fired generation and a significantly warmer-than-average summer. Supply showed some weakness despite increasing numbers of active drilling rigs. The result was broadly higher energy prices."

Our natural gas index data shows a clear spike in 2005, going up from 95.39 in 2004 to 142.40 in 2005.

Newspaper Articles. The increase in natural gas prices in the aftermath of the hurricanes received media attention. An example from NBC News is included in what follows.[26]

"Gas prices in cities across the United States soared by as much as 40 cents a gallon from Tuesday to Wednesday, a surge blamed on disruptions by Hurricane Katrina in Gulf of Mexico oil production."

C.4 Metals and Mineral Commodities

i. Aluminum

Year of Event: 1994.
Type of Event: Positive price shock.
According to the "Commodity Markets and Developing Countries" report from the World Bank, aluminum prices increased in 1994 due to a reduction in stocks, attributed primarily to the cutbacks in production by major producers.[27] Our data reveal that aluminum prices went up 30 percent in 1994.

Newspaper Articles. A newspaper article illustrates the cuts in supply.
A Loose Plan On Output of Aluminum, New York Times (January 31, 1994):[28]

"Six leading aluminum producers have agreed on ways to reduce a serious oversupply that has depressed prices on world markets."

ii. Iron ore

Year of Event: 2019.
Type of Event: Positive price shock.
The collapse of a mining dam in Brazil, the largest iron ore producer, led the price increase. Our data reveal that iron ore prices increased around 35 percent in 2019.[29]

iii. Lead

Year of Event: 2017.

[26] http://www.nbcnews.com/id/9146363/ns/business-local_business/t/pump-prices-jump-across-us-after-katrina/#.W3NQbehKiUk.
[27] http://https://thedocs.worldbank.org/en/doc/475131464184948121-0050022016/original/CMO1994November.pdf.
[28] Article available at: https://www.nytimes.com/1994/01/31/business/a-loose-plan-on-output-of-aluminum.html?.
[29] https://www.ft.com/content/8c2f26f6-72b0-11e9-bf5c-6eeb837566c5.

Type of Event: Negative price shock.

According to the ”Commodity Markets Review” from the World Bank, prices declined due to rising stocks and expectation that suspended production from the Magellan mine in Australia will be allowed to resume in the first quarter of 2008. In our data lead prices declined 32%.[30]

iv. Nickel

Year of Event: 2001.

Type of Event: Positive price shock.

According to World Bank (2001), various supply problems contributed to the tight market, particularly technical problems bringing on new capacity in Australia and labor strikes in Canada.[31] In our data nickel prices increased by 44%.

C.5 Country-Specific Assumptions

Our approach requires the omission of certain events when they are a result of unique weather conditions or political incidents exclusive to a specific country. The following exclusions have been implemented:

- The cocoa price surge of 2002, instigated by an attempted coup in Côte d'Ivoire amidst an ongoing civil war and escalating tensions, is omitted for this particular country.
- The sugar price shock in 2006, which was due to drought conditions in Thailand, is not considered in our analysis for this country.
- The 2010 spike in cereal prices, precipitated by weather conditions in Russia, Kazakhstan, and Ukraine, results in these countries' exclusion from the event.
- The cotton price shock in 2010, induced by weather-related incidents in Pakistan, is disregarded for Pakistan in our analysis.
- The rubber price shock in 2010, triggered by droughts in Thailand, leads to Thailand's exclusion from this event in our analysis.
- The cereal shock in 2010 took place later in the year and was more reflected in 2011 prices. We therefore use this shock for 2011.
- The 2019 disruption to iron ore prices, attributable to the collapse of a mining dam in Brazil, is specifically excluded for Brazil in our study.

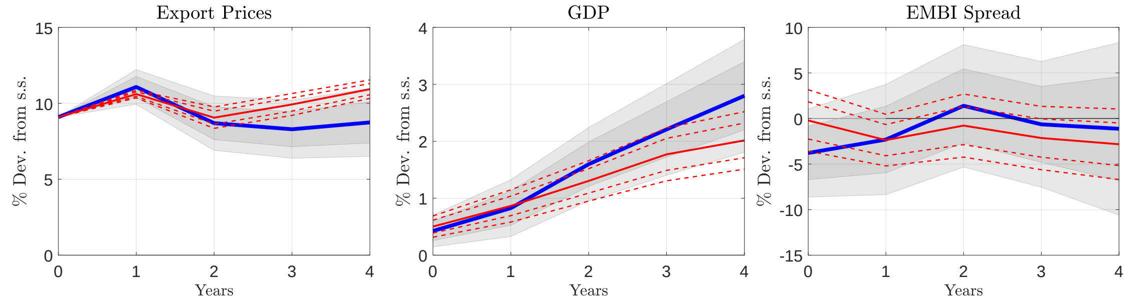
[30] https://thedocs.worldbank.org/en/doc/324111462981400952-0050022016/original/CMO2007December.pdf.

[31] https://thedocs.worldbank.org/en/doc/398441462978606788-0050022016/original/CMO2001GEP.pdf.

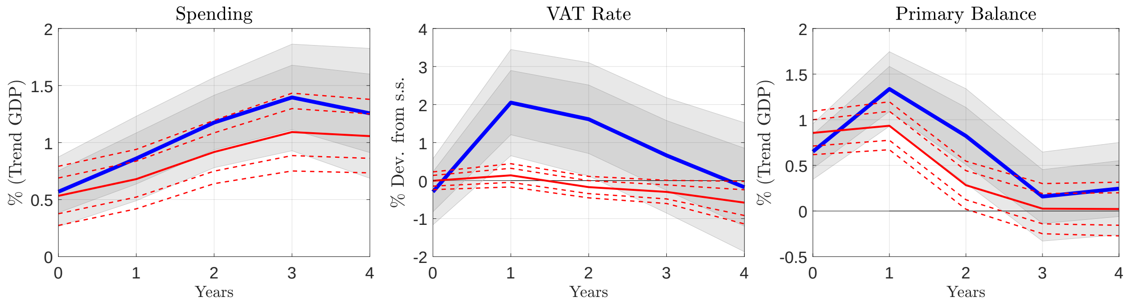
D Additional Results

D.1 LATE vs. OLS

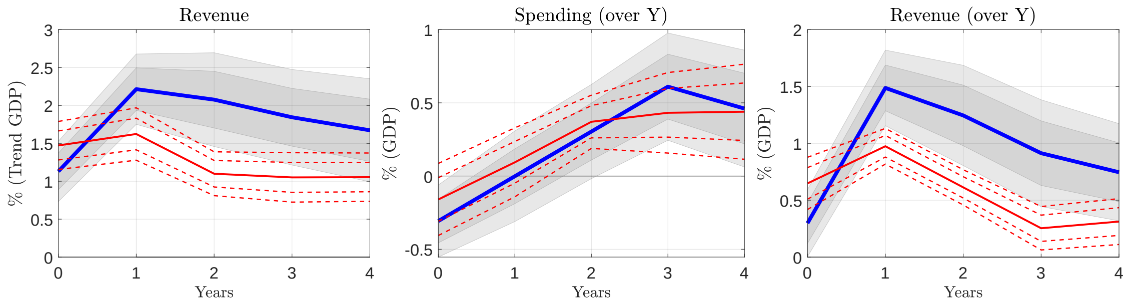
Figure D.1: LATE vs. OLS



(a) Macro Effects



(b) Fiscal Instruments (c) Fiscal Stance

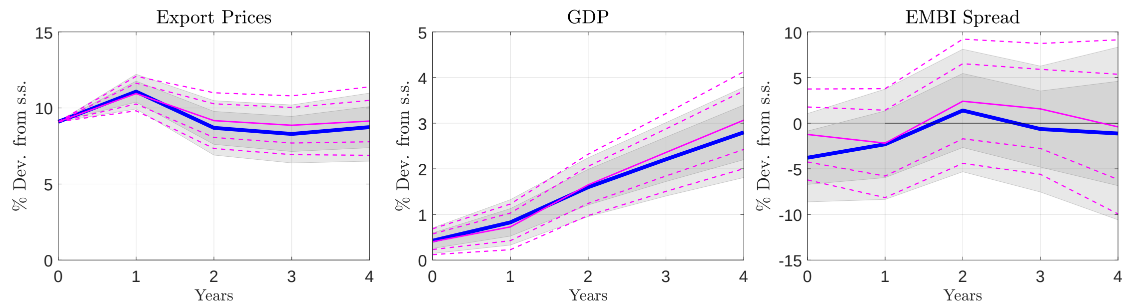


(d) Other Fiscal Outcomes

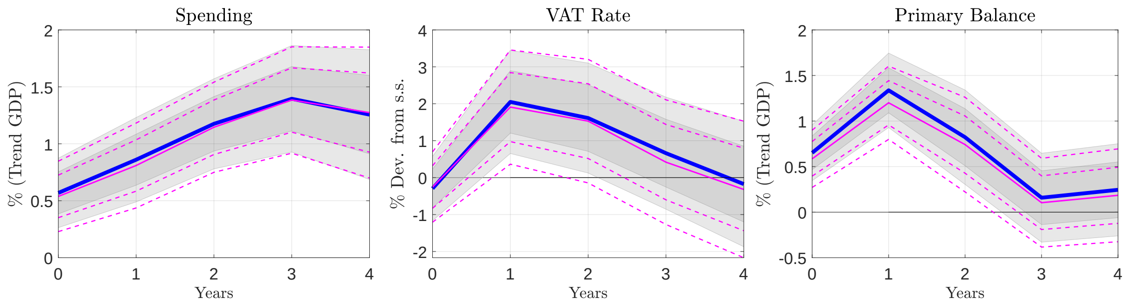
Notes: The Impulse Responses show the LATE (in blue) of one standard deviation increase in Px driven by commodity price shocks. OLS results are in red. Areas and dashed lines denote 68% and 90% confidence intervals.

D.2 Alternative Instruments

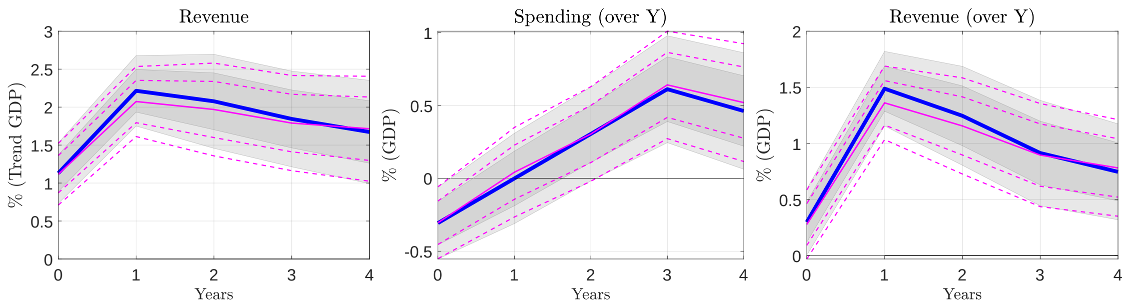
Figure D.2: Alternative Instrument: Using Only the Sign of the Event



(a) Macro Effects



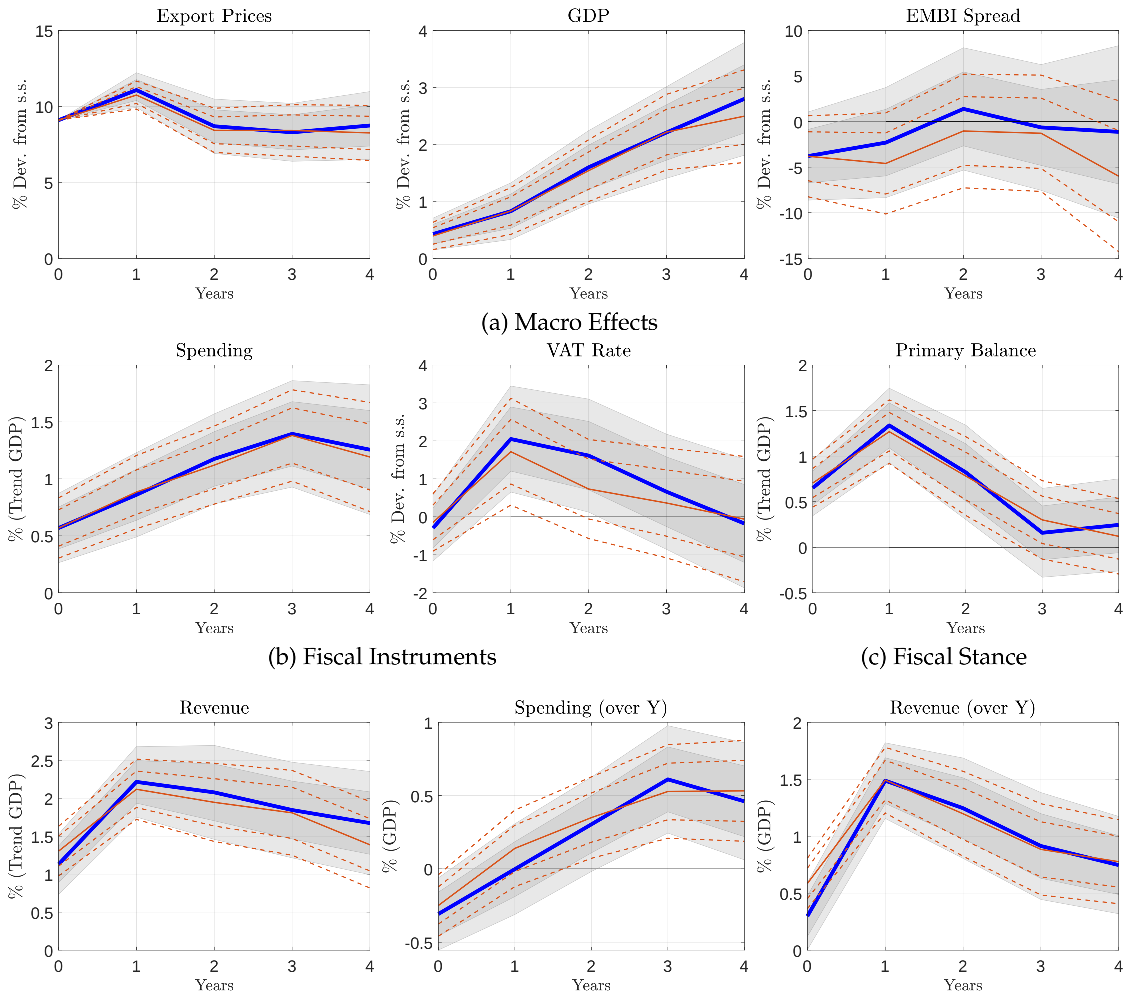
(b) Fiscal Instruments (c) Fiscal Stance



(d) Other Fiscal Outcomes

Notes: The Impulse Responses show the baseline LATE (in blue) of one standard deviation increase in Px driven by commodity price shocks. In magenta we report alternative estimates constructed with an instrument that only accounts for the information on the sign of the shock for each event and differing country exposure to it. Specifically, $\tilde{z}_{i,t} = \sum_j \mathbf{1}(\mathrm{w}_{i,c} > \underline{\mathrm{w}})\mathrm{w}_{i,c}\mathrm{Sign}_{j,t}$, where the signs are the ones reported in Table 2. Areas and dashed lines denote 68% and 90% confidence intervals.

Figure D.3: Alternative Instrument: Time-Varying Weights



(a) Macro Effects

(b) Fiscal Instruments

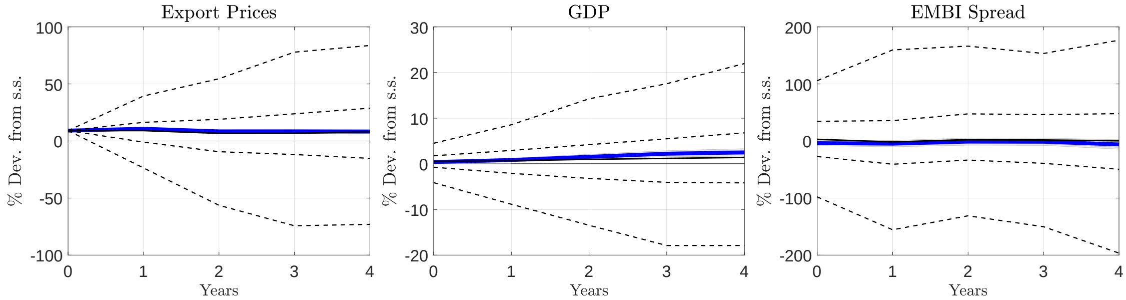
(c) Fiscal Stance

(d) Other Fiscal Outcomes

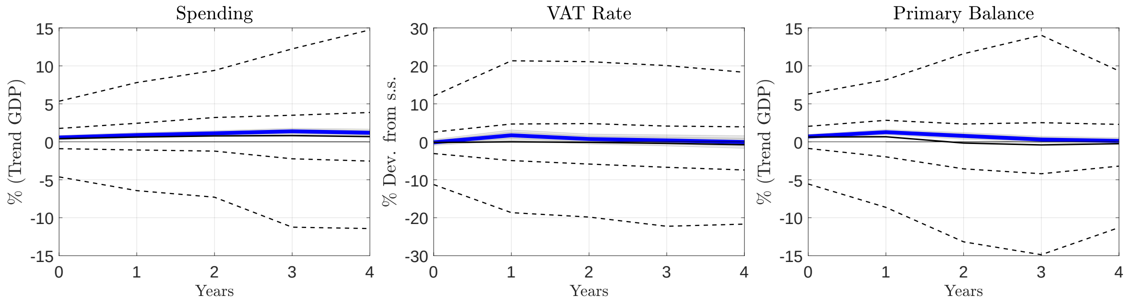
Notes: The Impulse Responses show the baseline LATE (in blue) of one standard deviation increase in Px driven by commodity price shocks. In magenta we report alternative estimates constructed with an instrument that uses time varying export shares. Specifically, $\tilde{z}_{i,t} = \sum_j \mathbf{1}(\mathbb{w}_{i,c,t-1} > \underline{\mathbb{w}})\mathbb{w}_{i,c,t-1}\mathfrak{q}_{j,t}$, where $\mathbb{w}_{i,c,t-1}$ is the export weight of commodity $\mathfrak{c}$ (associated with event j) for country i at time $t-1$. Areas and dashed lines denote 68% and 90% confidence intervals.

D.3 Placebo Exercise

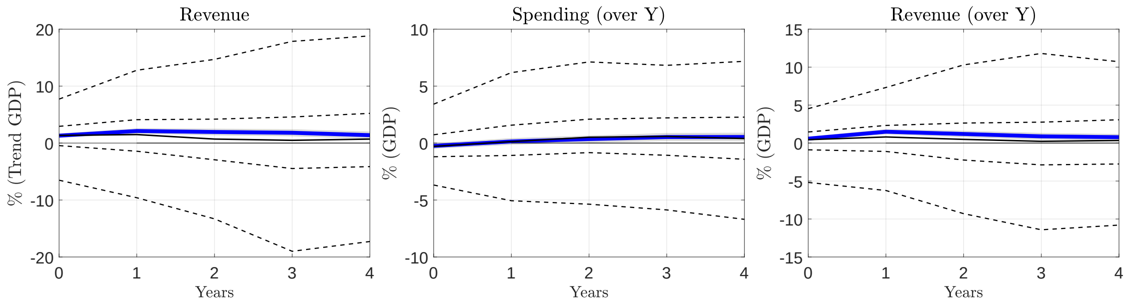
Figure D.4: Placebo With Reshuffled Events



(a) Macro Effects



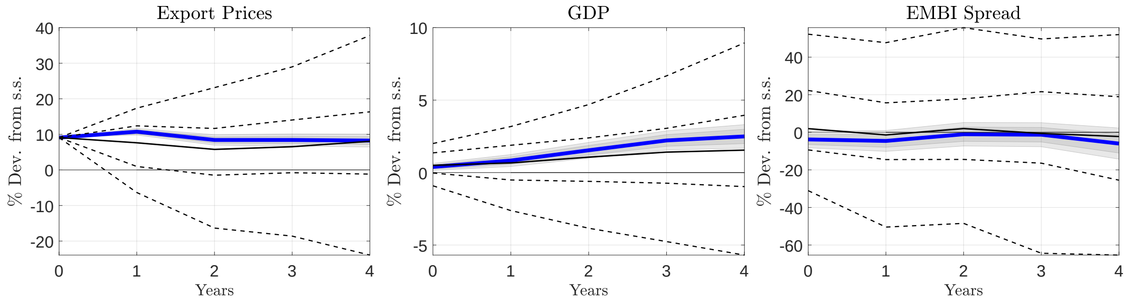
(b) Fiscal Instruments (c) Fiscal Stance



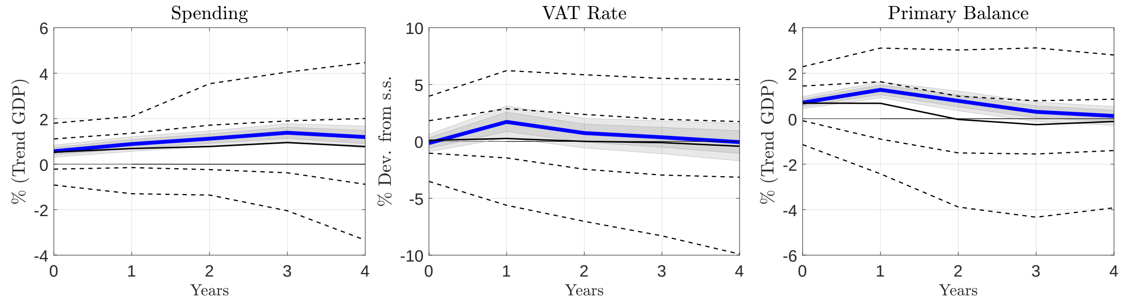
(d) Other Fiscal Outcomes

Notes: The Impulse Responses show the baseline LATE (in blue) of one standard deviation increase in Px driven by commodity price shocks (grey areas denote 68% and 90% confidence intervals). The responses in black report median as well as the 68% and 90% point-wise quantiles from the estimates resulting from 10,000 placebo instruments. The latter are constructed by keeping the same subset of commodity events, but allocating them at random dates.

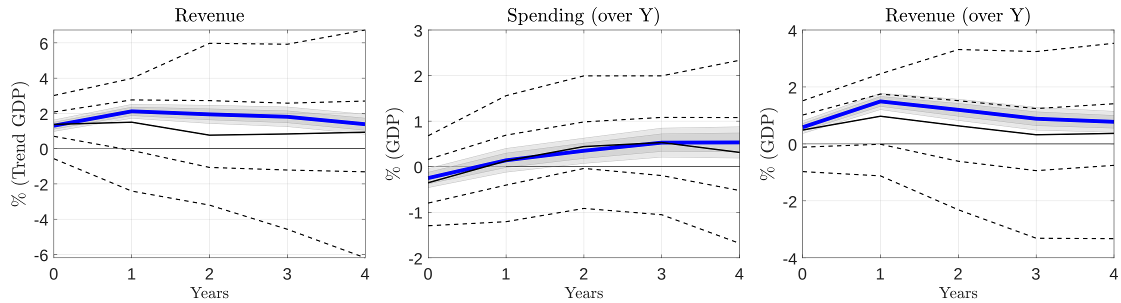
Figure D.5: Placebo With Reshuffled Events with Large F-statistic



(a) Macro Effects



(b) Fiscal Instruments (c) Fiscal Stance

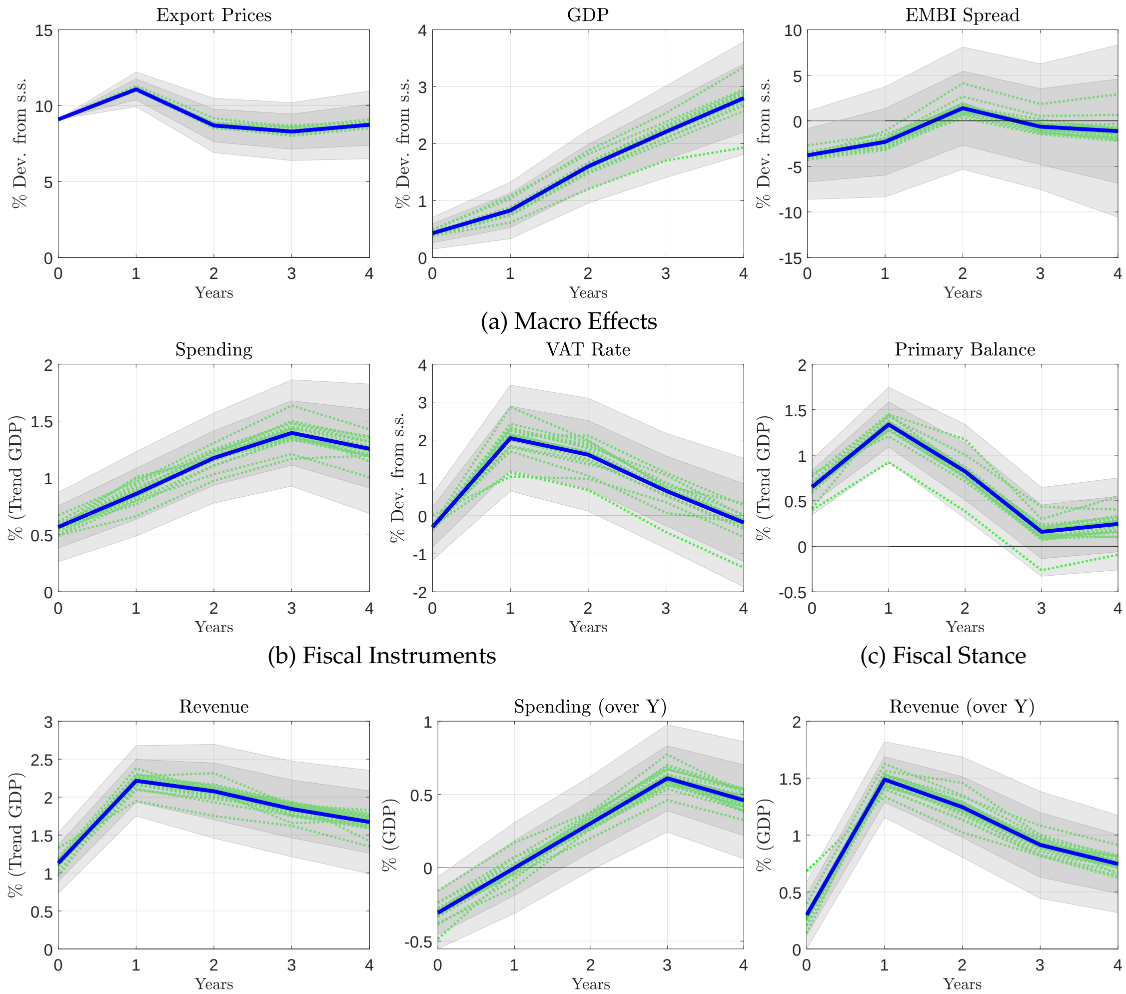


(d) Other Fiscal Outcomes

Notes: The Impulse Responses show the baseline LATE (in blue) of one standard deviation increase in Px driven by commodity price shocks (grey areas denote 68% and 90% confidence intervals). The responses in black report median as well as the 68% and 90% point-wise quantiles from the estimates resulting from the selection of placebo instruments used in Figure D.4 with an F-statistic in the first stage larger than 100.

D.4 Dropping Countries and Events

Figure D.6: Individual Country Drop



(a) Macro Effects

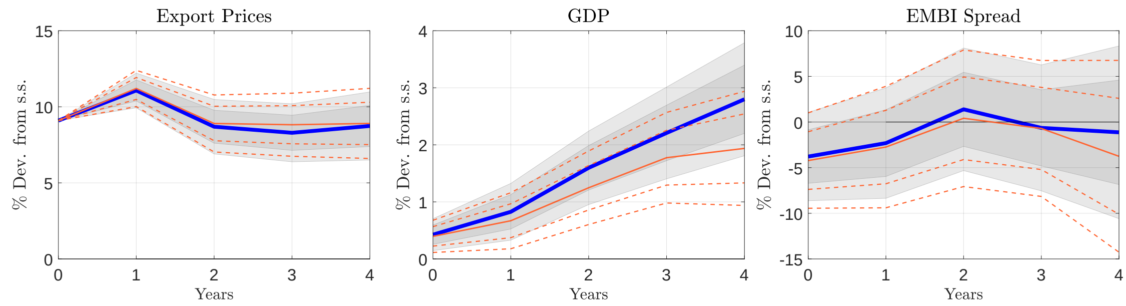
(b) Fiscal Instruments

(c) Fiscal Stance

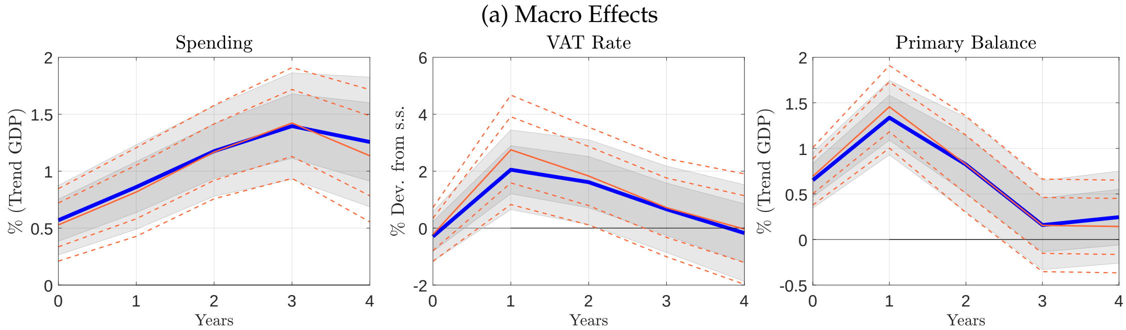
(d) Other Fiscal Outcomes

Notes: The Impulse Responses show the baseline LATE (in blue) of one standard deviation increase in Px driven by commodity price shocks. Lines in green correspond to the LATE mean estimates dropping from the dataset one country at the time. Gray areas denote 68% and 90% confidence intervals.

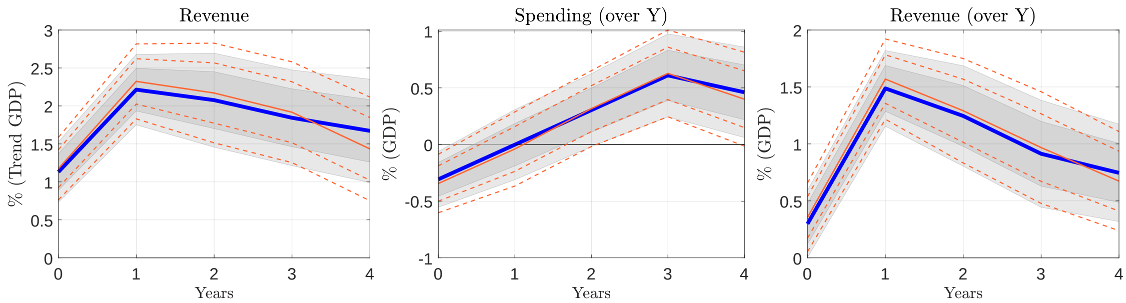
Figure D.7: Robustness to Dropping Major Macro Events



(a) Macro Effects



(b) Fiscal Instruments (c) Fiscal Stance

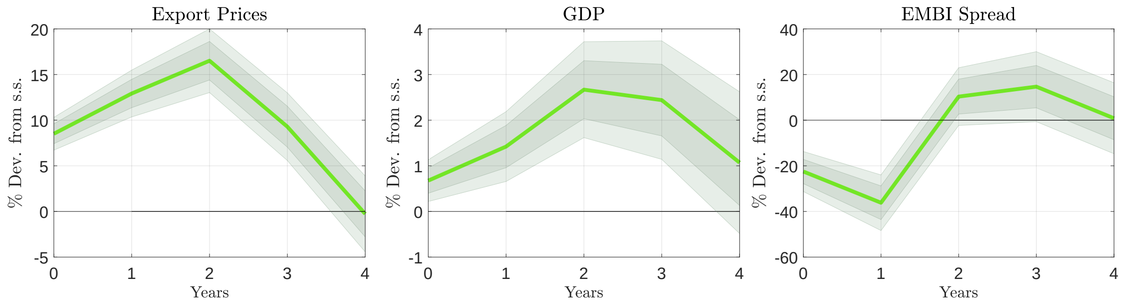


(d) Other Fiscal Outcomes

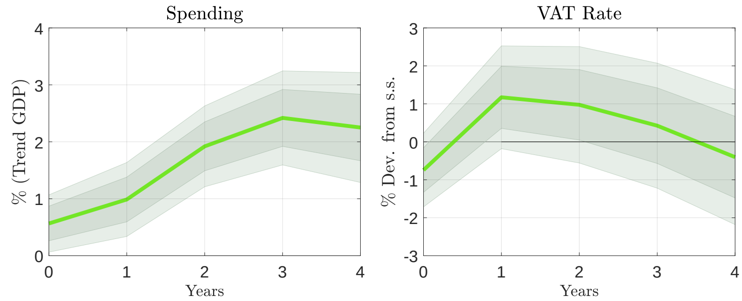
Notes: The Impulse Responses show the baseline LATE (in blue) of one standard deviation increase in Px driven by commodity price shocks. Lines in orange correspond to the LATE mean estimates eliminating from the set of events the ones coinciding with major global events (the Asian Crisis in 1997 and the Great Recession the associated recovery 2008-10). Gray areas denote 68% and 90% confidence intervals.

D.5 Additional Shocks

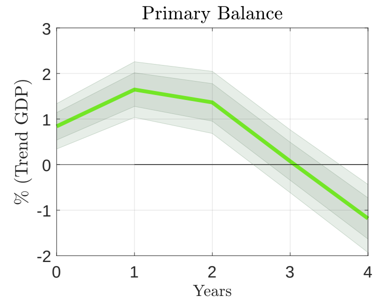
Figure D.8: Decline in the BAA Spread Driven by U.S. Monetary Policy Shocks



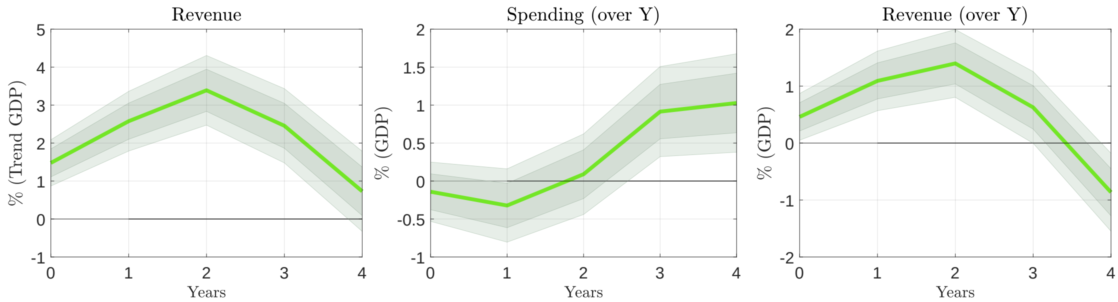
(a) Macro Effects



(b) Fiscal Instruments



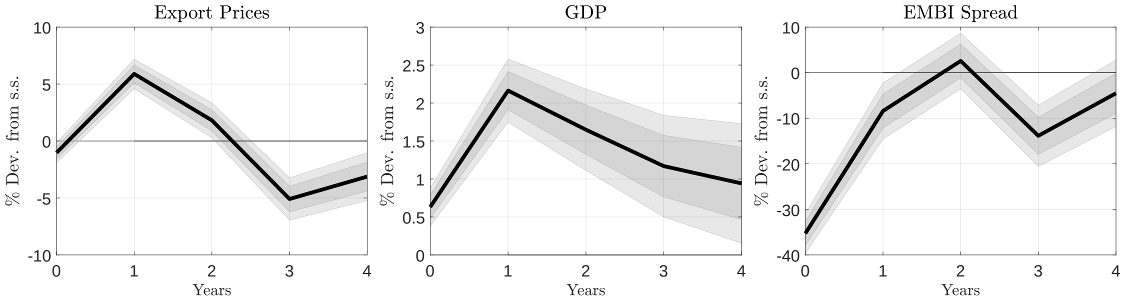
(c) Fiscal Stance



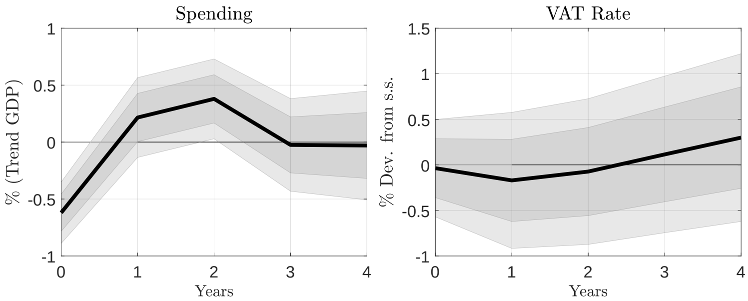
(d) Other Fiscal Outcomes

Notes: The Impulse Responses show the LATE of one standard deviation decline in the BAA spread driven by a more accommodative U.S. monetary policy. Shaded areas denote 68% and 90% confidence intervals.

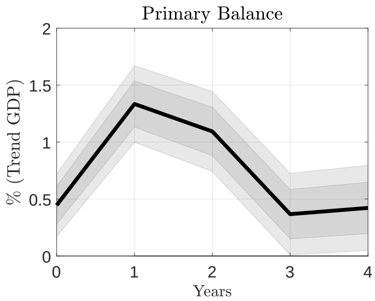
Figure D.9: Decline in the BAA Spread driven by Reduction in Global Risk Appetite



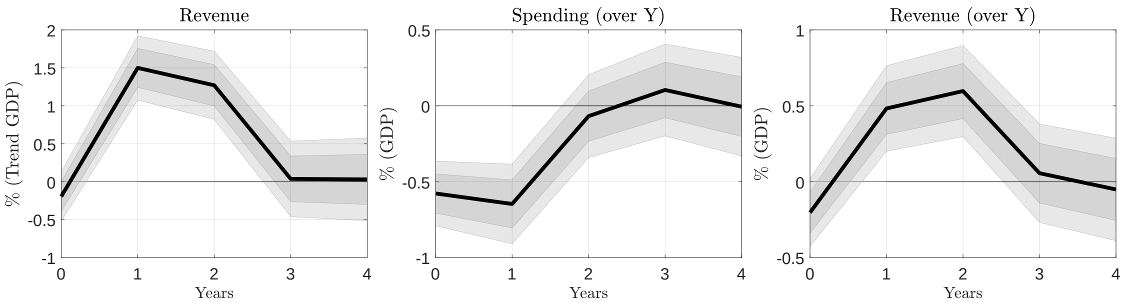
(a) Macro Effects



(b) Fiscal Instruments



(c) Fiscal Stance



(d) Other Fiscal Outcomes

Notes: The Impulse Responses show the LATE of one standard deviation decline in the BAA spread driven by shifts in global risk appetite. Shaded areas denote 68% and 90% confidence intervals.

E Theoretical Framework

E.1 Households

The economy is populated by a representative household that maximizes life-time utility

$$\mathcal{U}_0 = \mathbb{E}_0 \sum_{t=0}^{\infty} \beta^t \left\{ U\left(c_t, \bar{c}_{t-1}, \bar{g}_t, \bar{g}_{t-1}, h_t^x, h_t^n\right) \right\}, \tag{E.1}$$

with c_t denoting consumption, $\bar{g}_t$ government spending, h_t^x hours worked in the exportable sector, h_t^n hours worked in the non-tradable sector, and β the discount factor. The representative household values consumption in relation to an (external) habit level, calculated based on the level of (past) aggregate consumption. This is akin to the concept of "keeping up with the Joneses."[32] We introduce government in the utility function (as in, e.g., Barro, 1981; Christiano and Eichenbaum, 1992). To preserve a symmetric treatment of private and public consumption, we incorporate habit formation in government spending, in the spirit of Leith, Moldovan, and Rossi (2015). In maximizing utility, each household takes aggregate variables, denoted with a bar (i.e., the habit level of consumption and government spending) as given.

The period utility function featuring separable preferences is defined as

$$U = \phi \frac{\left(c_t - b\bar{c}_{t-1}\right)^{1-\sigma_c} - 1}{1-\sigma_c} + (1-\phi) \frac{\left(\bar{g}_t - b\bar{g}_{t-1}\right)^{1-\sigma_g} - 1}{1-\sigma_g} - \sum_{j=\{n,x\}} \frac{\left(h_t^j\right)^{1+\varphi}}{1+\varphi}, \tag{E.2}$$

where ϕ is the weight attached to utility derived from habit-adjusted consumption; σ_c and σ_g denote, respectively, the inverse of the intertemporal elasticities of substitution of habit adjusted consumption and government spending; b indicates the overall degree of habit formation; and φ is the inverse of the Frisch elasticity of labor supply, which we assume to be common across non-tradable (n) and exportable sectors (x).

Aggregate consumption combines tradable (c_t^τ) and non-tradable consumption (c_t^n):

$$c_t = \left(\chi^{\frac{1}{\epsilon}} \left(c_t^\tau\right)^{1-\frac{1}{\epsilon}} + (1-\chi)^{\frac{1}{\epsilon}} \left(c_t^n\right)^{1-\frac{1}{\epsilon}}\right)^{\frac{1}{1-\frac{1}{\epsilon}}}, \tag{E.3}$$

where χ denotes the share of tradable consumption in aggregate consumption and ϵ is the elasticity of substitution between tradable and non-tradable consumption. Tradable consumption bundles exportable (c_t^x) and importable consumption (c_t^m) following:

$$c_t^\tau = \left(\nu^{\frac{1}{\eta}} \left(c_t^m\right)^{1-\frac{1}{\eta}} + (1-\nu)^{\frac{1}{\eta}} \left(c_t^x\right)^{1-\frac{1}{\eta}}\right)^{\frac{1}{1-\frac{1}{\eta}}}, \tag{E.4}$$

where ν is the share of importable goods in tradable consumption and η is the elasticity of substitution between importable and exportable goods. There are two aggregators for

[32] See, for example, Abel (1990) and Constantinides (1990). The inclusion of habits introduces endogenous persistence, and leads to a smoother response of consumption, aligning with observations from empirical analysis. But otherwise does not affect, in any way, the qualitative response of the model.

government spending, following the same form as equations (E.3) and (E.4). We assume that the parameters in these equations are identical for both the government spending aggregator and the government tradable aggregator.

By substituting the individual demands into the consumption aggregators, we can derive expressions for aggregate prices (which we take as a numéraire) and tradable prices,

$$1 = \left[\chi \left(p_t^{\tau}\right)^{1-\epsilon} + (1-\chi)\left(p_t^{n}\right)^{1-\epsilon}\right]^{\frac{1}{1-\epsilon}}, \tag{E.5}$$

$$p_t^{\tau} = \left[\nu \left(p_t^{m}\right)^{1-\eta} + (1-\nu)\left(p_t^{x}\right)^{1-\eta}\right]^{\frac{1}{1-\eta}}. \tag{E.6}$$

Let q_t denote the real exchange rate. We assume that the LOOP holds separately for export and import prices: $p_t^x = p_t^{x,\$} q_t$ and $p_t^m = p^{m,\$} q_t$. Export and import prices (denominated in real U.S. dollars), $p_t^{x,\$}$ and $p^{m,\$}$, are taken as exogenous to the domestic country given that they are determined in global markets. In line with the empirical analysis, we focus on the transmission of commodity price shocks though their impact on export prices and assume a constant $p^{m,\$}$.

Households maximize lifetime utility subject to a sequence of budget constraints of the form:

$$q_t\left(1+r_{t-1}\right) d_{t-1} - q_t d_t = \sum_{j=\{n,x\}} \left(w_t^j h_t^j + \pi_t^j\right) - \theta_t c_t, \tag{E.7}$$

where d_t represents the stock of private debt (expressed in terms of foreign output), r_t the associated interest rate, w_t^j the sectoral real wage rate, π_t^j sectoral profits, and θ_t is the gross consumption tax rate. In the baseline model, we assume that international asset markets are incomplete. Specifically, the private and public sectors can only insure themselves by trading (one period) risk-free bonds.

The interest rate paid on private debt is given by:

$$r_t = r^* + s_t, \tag{E.8}$$

where s_t is a spread paid by the private sector above the foreign interest rate, r^*. The process for the shock is specified later. The spread is assumed to be an increasing function of debt and a decreasing function of export price shocks,

$$s_t = s + \psi\left[\exp\left(\bar{d}_t - d\right) - 1\right] - \rho_x \ln\left(p_t^{x,\$}\right). \tag{E.9}$$

The assumption that the interest rate depends on debt follows Schmitt-Grohé and Uribe (2003), which in turn ensures that the model has a stationary solution. Here, s denotes the value of the spread in the steady state, ψ a debt elastic premium, d the steady state value of private debt, and $\bar{d}_t$ the aggregate level of foreign debt.[33] We also explore an additional mechanism through which export price shocks influence credit spreads, with ρ_x denoting

[33] In line with Schmitt-Grohé and Uribe (2003) we assume that individual agents do not internalize the effects of their decisions on the level of the spreads.

the sensitivity of the credit spread to export price shocks. We follow the approach by Drechsel and Tenreyro (2018), who document a pronounced comovement between interest rates and commodity prices and show the importance of this channel through the lens of an SOE model.

Taking $\bar{c}_t$, $\bar{g}_t$, $\bar{d}_t$ and prices as given, the first order conditions of the household problems with respect to c_t, c_t^n, c_t^τ, c_t^m, c_t^x, d_t, h_t^n and h_t^x are:

$$q_t \lambda_t = \mathbb{E}_t \left\{ \beta \left(1 + r_t\right) q_{t+1} \lambda_{t+1} \right\}; \tag{E.10}$$

$$\theta_t \lambda_t = \phi_t \left(c_t - b\bar{c}_{t-1}\right)^{-\sigma_c}; \tag{E.11}$$

$$\begin{aligned} c_t^n &= \left(1-\chi\right)\left(p_t^n\right)^{-\epsilon} c_t, & c_t^\tau &= \chi \left(p_t^\tau\right)^{-\epsilon} c_t, \\ c_t^m &= \nu \left(\frac{p_t^m}{p_t^\tau}\right)^{-\eta} c_t^\tau; & c_t^x &= \left(1-\nu\right)\left(\frac{p_t^x}{p_t^\tau}\right)^{-\eta} c_t^\tau; \end{aligned} \tag{E.12}$$

$$\lambda_t w_t^n = \left(h_t^n\right)^\varphi, \qquad \lambda_t w_t^x = \left(h_t^x\right)^\varphi. \tag{E.13}$$

Equation (E.10) is the uncovered interest rate parity condition, equation (E.11) the marginal utility of consumption, equations in (E.12) are the individual consumption demands of non-tradable, tradable, importable and exportable goods, respectively, and equations in (E.13) the labor supplies of the non-tradable and exportable sectors. Note that the individual demands of government spending are of the same form as individual consumption demands.

E.2 Firms' Problems

Firms in sector $j = \{n, x\}$ maximize profits:

$$\pi_t^j = p_t^j y_t^j - w_t^j h_t^j - p_t^m m_t^j, \tag{E.14}$$

subject to the following technological constraint:

$$y_t^j = a^j \left(h_t^j\right)^{\alpha_j} \left(m_t^j\right)^{1-\alpha_j}, \tag{E.15}$$

where m_t^j denotes intermediate imported inputs by sector j, $1-\alpha_j$ is the share of imported intermediates used in production, and a^j is sectoral productivity. The first order conditions are:

$$\frac{w_t^j}{p_t^j} = \alpha_j \frac{y_t^j}{h_t^j}, \tag{E.16}$$

$$\frac{p_t^m}{p_t^j} = \left(1-\alpha_j\right)\frac{y_t^j}{m_t^j}. \tag{E.17}$$

Equations (E.16) and (E.17) define the sectoral labor demand and the demand for imported intermediates for each sector. Specifically, sectoral wage rates equate the marginal product of labor, and the demand for imported intermediates is set by matching their marginal products to the relative prices between imported intermediates and sectoral prices.

E.3 Government and Market Clearing

The government's intertemporal budget constraint is given by:

$$q_t \left(1 + r_{t-1}^g\right) d_{t-1}^g - q_t d_t^g = (\theta_t - 1) c_t - g_t, \tag{E.18}$$

where g_t is aggregate government spending. The variable r_t^g is the interest rate paid on public debt, which takes the following form:

$$r_t^g = r^* + s_t^g, \tag{E.19}$$

where s_t^g is a spread paid by the government above and beyond the foreign interest rate, r^*. Consistent with prior analysis, the spread on public debt is a function of debt and export price shocks:

$$s_t^g = s + \psi_g \left[\exp\left(\bar{d}_t^g - d^g\right) - 1\right] - \rho_{x,g} \ln\left(p_t^{x,\$}\right), \tag{E.20}$$

where s denotes the value of the spread in the steady state, ψ_g a debt elastic premium, d^g the steady-state value of government debt, and $\rho_{x,g}$ the sensitivity of the government spread to export price shocks. The last mechanism can be understood in terms of, for example, a decrease in sovereign risk, associated with an increase in government revenues and foreign reserves following an increase in the main exporting commodity price as in Hamann et al. (2023).

Adding up the budget constraints of the representative household and the government yields the net foreign asset position:

$$q_t (1 + r_{t-1}) d_{t-1} - q_t d_t + q_t \left(1 + r_{t-1}^g\right) d_{t-1}^g - q_t d_t^g = tb_t, \tag{E.21}$$

where the trade balance is given by the following expression:

$$tb_t = p_t^x x_t - p_t^m \mathfrak{m}_t. \tag{E.22}$$

Exports are defined as $x_t = y_t^x - c_t^x - g_t^x$ and imports as $\mathfrak{m}_t = m_t^x + m_t^n + c_t^m + g_t^m$. Market clearing in the non-tradable goods market requires that:

$$y_t^n = c_t^n + g_t^n. \tag{E.23}$$

Gross value added is defined as:

$$y_t = \alpha_n p_t^n y_t^n + \alpha_x p_t^x y_t^x. \tag{E.24}$$

Consistent with the data (and the empirical analysis), output, consumption, government spending, fiscal revenues, the primary balance, and the trade balance in the model are defined, respectively, in constant prices as:

$$\mathcal{Y}_t = \alpha_n p^n y_t^n + \alpha_x q y_t^x, \tag{E.25}$$

$$\mathcal{C}_t = \frac{c_t}{y_t}\mathcal{Y}_t, \tag{E.26}$$

$$\mathcal{G}_t = \frac{g_t}{y_t}\frac{\mathcal{Y}_t}{y}, \tag{E.27}$$

$$\mathcal{R}_t = \frac{(\theta_t - 1)\, c_t}{y_t}\frac{\mathcal{Y}_t}{y}, \tag{E.28}$$

$$\mathcal{PB}_t = \frac{(\theta_t - 1)\, c_t - g_t}{y_t}\frac{\mathcal{Y}_t}{y}, \tag{E.29}$$

$$\mathcal{TB}_t = \frac{tb_t}{y_t}\frac{\mathcal{Y}_t}{y}. \tag{E.30}$$

E.4 Optimal Ramsey Policy

The Ramsey planner chooses aggregate government spending and taxes (real prices and allocations) to maximize the utility of the representative household subject to the implementability conditions described above (i.e., first order conditions of households and firms, market clearing conditions together with budget constraints). To reduce the dimensionality of the problem, we substitute the first order conditions with respect to imported intermediates (m_t^x and m_t^n, in equation E.17) directly into the sectoral production functions (equation E.15). In addition, by combining the labor supply and demand for each sector (equations E.13 and E.16), we eliminate the wage rates (w_t^x and w_t^n). This process yields four key constraints:

$$y_t^j = (a_t)^{\frac{1}{\alpha_j}} \left(\frac{p_t^j}{p_t^m}\right)^{\frac{1-\alpha_j}{\alpha_j}} (1-\alpha_j)^{\frac{1-\alpha_j}{\alpha_j}} h_t^j, \tag{E.31}$$

$$\left(h_t^j\right)^{1+\varphi} = \lambda_t p_t^j \alpha_j y_t^j, \tag{E.32}$$

for $j = \{n, x\}$. The Ramsey planner therefore chooses λ_t, c_t, c_t^τ, c_t^m, c_t^x, h_t^n, h_t^x, y_t^n, y_t^x, p_t^n, p_t^τ, q_t, d_t^g, d_t, g_t, g_t^τ, g_t^m, g_t^x, and θ_t so as to maximize welfare, specified in equations (E.1)-(E.2), subject to the set of implementability conditions that can replicate the competitive equilibrium (equations E.5-E.6, E.10-E.12, E.16-E.32 and the corresponding government demands for non-tradable, tradable, exportable, and importable goods). Note that we assume that the Ramsey planner internalizes the externality associated with external habit formation.[34] The full problem, the planner's first order conditions, and the steady state are detailed in Appendix F.1. Although our analysis focuses on export prices as the only source of variation, the Ramsey policy is, by definition, not shock-specific.

E.5 Model Calibration

We calibrate the model economy to annual frequency. The steady-state real foreign interest rate (r^*) is set at 0.04, following Schmitt-Grohé and Uribe (2003), while the steady state spread (s) equal to 0.032 is consistent with Aguiar and Gopinath (2007). The debt elastic premia,

[34] Since the Ramsey planner internalizes the externality coming from habits, we drop the bar over the variable to denote aggregates from the households' first order conditions.

represented by the parameters ψ and ψ_g, are both set to 0.58, in line with the median point estimate of Schmitt-Grohé and Uribe (2018) for a similar sample of EMDEs. We assign the values of $\rho_x = \rho_{x,g}$ to be 0.05.[35] It is worth noting that from the optimality conditions of the Ramsey planner (under the assumption of incomplete markets), the government debt issuance (d^g) must be zero in equilibrium (for proof, see Fernández et al., 2021).

The habit persistence parameter, b, is set to 0.2. This parameter value lies within the range of parameter estimates of DSGE models estimated at quarterly frequencies (0.65 and 0.85). In our baseline calibration, the intertemporal elasticities of substitution for private and public consumption are equal to 1 ($\sigma_c = \sigma_g = 1$) and the inverse of the Frisch elasticity of labor supply (φ) is 0.455, based on Mendoza (1995). In the absence of physical capital, we calibrate the non-tradable sector's labor intensity to be greater than the tradable sector ($\alpha_n > \alpha_x$). We set the imported intermediate shares following Arseneau and Leduc (2013), which implies values for the labor share of $\alpha_n = 0.9$ and $\alpha_x = 0.8$.

Following Schmitt-Grohé and Uribe (2018), we choose the elasticities of substitution ($\eta = 0.5$, $\epsilon = 1.1$) and the share of tradable in aggregate consumption (χ). The non-tradable output's value-added share is approximately 0.55, with χ fixed at 0.35, reflecting the absence of an importable producing sector. We also calibrate tradable consumption and government spending's importable share (ν) at 0.15, to target the export-to-output ratio at 0.25. We choose the value of ϕ to match the size of the government in the data (at 30% of GDP). The debt-to-GDP ratio is calibrated at 0.335 (annual terms), following Lane and Milesi-Ferretti (2007).

We assume that the export price shock process follows an AR(2) of the form:

$$\ln\left(p_t^{x,\$}\right) = \varrho_x^1 \ln\left(p_{t-1}^{x,\$}\right) + \varrho_x^2 \ln\left(p_{t-2}^{x,\$}\right) + \zeta_x \varepsilon_x, \tag{E.33}$$

where ϱ_x^1 and ϱ_x^2 denote the persistence coefficients, which are chosen to match the average response of the export price shocks in the empirical analysis, and are equal to 1.05 and -0.15, respectively. The dispersion of the export price process, ζ_x, is set the match the response of output at peak (with a value of 0.09). Export prices ($p^{x,\$}$) are normalized to 1 in the steady state, implying equivalent values for tradable prices and the real exchange rate, under the LOOP. This normalization ensures that the steady-state demands for importable and exportable goods equal the shares of tradable goods.

F Ramsey Fiscal Policy

In this Appendix, we present the baseline model (and the model without the credit spread channel) in Section F.1. The introduction of the shock processes and their calibration are discussed in Section F.2. The multiple shocks are introduced in the discussion in Section 6.4 of the main draft. The model under financial autarky is shown in Section F.3.

[35]These values are lower than the estimates reported by Drechsel and Tenreyro (2018) for three reasons: (i) their analysis focuses on shocks specific to commodity prices, whereas the model focuses mainly on export price shocks; (ii) we find that the response of the export price shock is more persistent relative to their central estimate; and (iii) their analysis is done solely on Argentinean data, which is a country whose credit spread is highly volatile.

F.1 Incomplete Asset Markets

The Ramsey planner chooses the instruments g_t and θ_t so as to maximize the welfare of the representative household subject to a set of conditions that implement the decentralized equilibrium. Note that in equilibrium, $\bar{d}_t = d_t, \bar{c}_t = c_t$ and $\bar{g}_t = g_t$. Since the Ramsey planner internalizes the externality coming from habits, we drop the bar over the variable to denote aggregates from the households' first order conditions. The planner's problem is to choose λ_t, c_t, c_t^{τ}, c_t^m, c_t^x, h_t^n, h_t^x, y_t^n, y_t^x, p_t^n, p_t^{τ}, q_t, d_t^g, d_t, g_t, g_t^{τ}, g_t^m, g_t^x and θ_t to maximize the following Lagrangian:

$$
\begin{aligned}
\mathcal{L} =& \mathbb{E}_0 \sum_{t=0}^{\infty} \beta^t \left[\phi_t \frac{(c_t - bc_{t-1})^{1-\sigma_c} - 1}{1-\sigma_c} + (1-\phi_t) \frac{(g_t - bg_{t-1})^{1-\sigma_g} - 1}{1-\sigma_g} - \sum_{j=\{n,x\}} \frac{\left(h_t^j\right)^{1+\varphi}}{1+\varphi} \right] + \\
& \mathbb{E}_0 \sum_{t=0}^{\infty} \beta^t \mu_{1t} \left[q_t d_t - q_t (1+r_{t-1}) d_{t-1} + q_t d_t^g - q_t (1+r_{t-1}) d_{t-1}^g + \right. \\
& \left. + \alpha_x q_t p_t^{x,\$} y_t^x - (1-\alpha_n) p_t^n y_t^n - q_t p_t^{x,\$} (c_t^x + g_t^x) - q_t p_t^{m,\$} (c_t^m + g_t^m) \right] + \\
& \mathbb{E}_0 \sum_{t=0}^{\infty} \beta^t \mu_{2t} \left[q_t d_t^g - q_t (1+r_{t-1}) d_{t-1}^g + (\theta_t - 1) c_t - g_t \right] + \\
& \mathbb{E}_0 \sum_{t=0}^{\infty} \beta^t \mu_{3t} \left[\phi_t (c_t - bc_{t-1})^{-\sigma_c} - \theta_t \lambda_t \right] + \\
& \mathbb{E}_0 \sum_{t=0}^{\infty} \beta^t \mu_{4t} \left[q_t \lambda_t - \beta (1+r_t) q_{t+1} \lambda_{t+1} \right] + \\
& \mathbb{E}_0 \sum_{t=0}^{\infty} \beta^t \mu_{5t} \left[c_t^n - (1-\chi) (p_t^n)^{-\epsilon} c_t \right] + \mathbb{E}_0 \sum_{t=0}^{\infty} \beta^t \mu_{6t} \left[c_t^{\tau} - \chi (p_t^{\tau})^{-\epsilon} c_t \right] + \\
& \mathbb{E}_0 \sum_{t=0}^{\infty} \beta^t \mu_{7t} \left[c_t^m - \nu \left(\frac{q_t p_t^{m,\$}}{p_t^{\tau}} \right)^{-\eta} c_t^{\tau} \right] + \mathbb{E}_0 \sum_{t=0}^{\infty} \beta^t \mu_{8t} \left[c_t^x - (1-\nu) \left(\frac{q_t p_t^{x,\$}}{p_t^{\tau}} \right)^{-\eta} c_t^{\tau} \right] + \\
& \mathbb{E}_0 \sum_{t=0}^{\infty} \beta^t \mu_{9t} \left[y_t^n - \left(\frac{p_t^n}{q_t p_t^{m,\$}} \right)^{\frac{1-\alpha_n}{\alpha_n}} (1-\alpha_n)^{\frac{1-\alpha_n}{\alpha_n}} (a_t^n)^{\frac{1}{\alpha_n}} h_t^n \right] + \\
& \mathbb{E}_0 \sum_{t=0}^{\infty} \beta^t \mu_{10t} \left[y_t^x - \left(\frac{p_t^{x,\$}}{p_t^{m,\$}} \right)^{\frac{1-\alpha_x}{\alpha_x}} (1-\alpha_x)^{\frac{1-\alpha_x}{\alpha_x}} (a_t^x)^{\frac{1}{\alpha_x}} h_t^x \right] + \\
& \mathbb{E}_0 \sum_{t=0}^{\infty} \beta^t \mu_{11t} \left[(h_t^n)^{1+\varphi} - \lambda_t \alpha_n p_t^n y_t^n \right] + \mathbb{E}_0 \sum_{t=0}^{\infty} \beta^t \mu_{12t} \left[(h_t^x)^{1+\varphi} - \lambda_t \alpha_x q_t p_t^{x,\$} y_t^x \right] + \\
& \mathbb{E}_0 \sum_{t=0}^{\infty} \beta^t \mu_{13t} \left[\left(\chi (p_t^{\tau})^{1-\epsilon} + (1-\chi) (p_t^n)^{1-\epsilon} \right)^{\frac{1}{1-\epsilon}} - 1 \right] + \\
& \mathbb{E}_0 \sum_{t=0}^{\infty} \beta^t \mu_{14t} \left[\left(\nu \left(p_t^{m,\$} \right)^{1-\eta} + (1-\nu) \left(p_t^{x,\$} \right)^{1-\eta} \right)^{\frac{1}{1-\eta}} q_t - p_t^{\tau} \right] + \\
& \mathbb{E}_0 \sum_{t=0}^{\infty} \beta^t \mu_{15t} \left[g_t^n - (1-\chi) (p_t^n)^{-\epsilon} g_t \right] + \mathbb{E}_0 \sum_{t=0}^{\infty} \beta^t \mu_{16t} \left[g_t^{\tau} - \chi (p_t^{\tau})^{-\epsilon} g_t \right] +
\end{aligned}
$$

$$\mathbb{E}_0 \sum_{t=0}^{\infty} \beta^t \mu_{17t} \left[g_t^m - \nu \left(\frac{q_t p_t^{m,\$}}{p_t^\tau} \right)^{-\eta} g_t^\tau \right] + \mathbb{E}_0 \sum_{t=0}^{\infty} \beta^t \mu_{18t} \left[g_t^x - (1-\nu) \left(\frac{q_t p_t^{x,\$}}{p_t^\tau} \right)^{-\eta} g_t^\tau \right] +$$
$$\mathbb{E}_0 \sum_{t=0}^{\infty} \beta^t \mu_{19t} \left[y_t^n - c_t^n - g_t^n \right],$$

where

$$r_t = r^* + s + \psi \left[\exp \left(d_t - d \right) - 1 \right] - \rho_x \ln \left(p_t^{x,\$} \right), \tag{F.1}$$
$$r_t^g = r^* + s + \psi_g \left[\exp \left(d_t^g - d^g \right) - 1 \right] - \rho_{x,g} \ln \left(p_t^{x,\$} \right). \tag{F.2}$$

The first order conditions are:

$$\lambda_t: \quad \mu_{4t} q_t = (1 + r_{t-1}) q_t \mu_{4t-1} + \mu_{11t} \alpha_n p_t^n y_t^n + \mu_{12t} \alpha_x q_t p_t^{x,\$} y_t^x + \mu_{3t} \theta_t, \tag{F.3}$$
$$c_t: \quad \mu_{5t} c_t^n + \mu_{6t} c_t^\tau + \sigma_c c_t \left[\mu_{3t} \phi_t \left(c_t - b c_{t-1} \right)^{-\sigma_c - 1} - \mathbb{E}_t \mu_{3t+1} \phi_{t+1} \left(c_{t+1} - b c_t \right)^{-\sigma_c - 1} \right] =$$
$$= c_t \left[\phi_t \left(c_t - b c_{t-1} \right)^{-\sigma_c} - \beta b \mathbb{E}_t \phi_{t+1} \left(c_{t+1} - b c_t \right)^{-\sigma_c} \right] + \mu_{2t} \left(\theta_t - 1 \right) c_t, \tag{F.4}$$
$$c_t^n: \quad \mu_{19t} = \mu_{5t}, \tag{F.5}$$
$$c_t^\tau: \quad \mu_{7t} c_t^m + \mu_{8t} c_t^x = \mu_{6t} c_t^\tau, \tag{F.6}$$
$$c_t^m: \quad \mu_{1t} q_t p_t^{m,\$} = \mu_{7t}, \tag{F.7}$$
$$c_t^x: \quad \mu_{1t} q_t p_t^{x,\$} = \mu_{8t}, \tag{F.8}$$
$$h_t^n: \quad \mu_{11t} (1+\varphi) \left(h_t^x \right)^{1+\varphi} = \left(h_t^n \right)^{1+\varphi} + \mu_{9t} y_t^n, \tag{F.9}$$
$$h_t^x: \quad \mu_{12t} (1+\varphi) \left(h_t^x \right)^{1+\varphi} = \left(h_t^x \right)^{1+\varphi} + \mu_{10t} y_t^x, \tag{F.10}$$
$$y_t^n: \quad \mu_{11t} \lambda_t \alpha_n p_t^n + \mu_{1t} \left(1 - \alpha_n \right) p_t^n = \mu_{9t} + \mu_{19t}, \tag{F.11}$$
$$y_t^x: \quad \mu_{12t} \lambda_t \alpha_x q_t p_t^{x,\$} = \mu_{1t} \alpha_x q_t p_t^{x,\$} + \mu_{10t}, \tag{F.12}$$
$$p_t^n: \quad \epsilon \left(\mu_{5t} c_t^n + \mu_{15t} g_t^n \right) + \mu_{13t} (1-\chi) \left(p_t^n \right)^{1-\epsilon} = \mu_{1t} \left(1 - \alpha_n \right) p_t^n y_t^n +$$
$$+ \mu_{11t} \lambda_t \alpha_n p_t^n y_t^n + \mu_{9t} \frac{1-\alpha_n}{\alpha_n} y_t^n, \tag{F.13}$$
$$p_t^\tau: \quad \epsilon \left[\mu_{6t} c_t^\tau + \mu_{16t} g_t^\tau \right] + \mu_{13t} \chi \left(p_t^\tau \right)^{1-\epsilon} = \eta \left[\mu_{7t} c_t^m + \mu_{8t} c_t^x + \mu_{17t} g_t^m + \mu_{18t} g_t^x \right] +$$
$$+ \mu_{14t} p_t^\tau, \tag{F.14}$$
$$q_t: \quad \mu_{12t} \lambda_t \alpha_x q_t p_t^{x,\$} y_t^x + \mu_{4t-1} q_t \left(1 + r_{t-1} \right) \lambda_{t-1} - q_t \mu_{4t} \lambda_t =$$
$$= \eta \left[\mu_{7t} c_t^m + \mu_{8t} c_t^x + \mu_{17t} g_t^m + \mu_{18t} g_t^x \right] + q_t \mu_{2t} \left[d_t^g - \left(1 + r_{t-1}^g \right) d_{t-1}^g \right] +$$
$$+ q_t \mu_{1t} \Big[d_t - \left(1 + r_{t-1} \right) d_{t-1} + d_t^g - \left(1 + r_{t-1}^g \right) d_{t-1}^g +$$
$$+ \alpha_x p_t^{x,\$} y_t^x - p_t^{x,\$} \left(c_t^x + g_t^x \right) - p_t^{m,\$} \left(c_t^m + g_t^m \right) \Big] +$$
$$+ \mu_{14t} p_t^\tau + \mu_{9t} \frac{1-\alpha_n}{\alpha_n} y_t^n, \tag{F.15}$$
$$d_t: \quad q_t \mu_{1t} - \mu_{4t} \mathbb{E}_t \beta \psi \left(e^{d_t - d} \right) q_{t+1} \lambda_{t+1} = \beta \mu_{1t+1} q_{t+1} \left[1 + r_t + \psi \left(e^{d_t - d} \right) d_t \right], \tag{F.16}$$
$$d_t^g: \quad \left(\mu_{1t} + \mu_{2t} \right) q_t = \beta \mathbb{E}_t \left(\mu_{1t+1} + \mu_{2t+1} \right) q_{t+1} \left[1 + r_t^g + \psi_g \left(e^{d_t^g - d^g} \right) d_t^g \right], \tag{F.17}$$
$$g_t: \quad \left[(1-\phi_t) \left(g_t - b g_{t-1} \right)^{-\sigma_g} - \beta b \mathbb{E}_t \left(1 - \phi_{t+1} \right) \left(g_{t+1} - b g_t \right)^{-\sigma_g} \right] g_t = \mu_{2t} g_t +$$

$$+\mu_{15t}g_t^n+\mu_{16t}g_t^\tau, \tag{F.18}$$

$$g_t^n: \quad \mu_{15t}=\mu_{19t}, \tag{F.19}$$

$$g_t^\tau: \quad \mu_{16t}g_t^\tau=\mu_{17t}g_t^m+\mu_{18t}g_t^x, \tag{F.20}$$

$$g_t^m: \quad q_t p_t^{m,\$}\mu_{1t}=\mu_{17t}, \tag{F.21}$$

$$g_t^x: \quad q_t p_t^{x,\$}\mu_{1t}=\mu_{18t}, \tag{F.22}$$

$$\theta_t: \quad \mu_{2t}c_t=\mu_{3t}\lambda_t, \tag{F.23}$$

And 19 Lagrange multipliers associated with the implementability constraints.

F.1.1 Steady State

The steady state consists of the following set of equations:

$$1=\beta\left(1+r\right), \tag{F.24}$$

$$c^n=\left(1-\chi\right)\left(p^n\right)^{-\epsilon}c, \tag{F.25}$$

$$c^\tau=\chi\left(p^\tau\right)^{-\epsilon}c, \tag{F.26}$$

$$c^m=\nu c^\tau, \tag{F.27}$$

$$c^x=\left(1-\nu\right)c^\tau, \tag{F.28}$$

$$g^n=\left(1-\chi\right)\left(p^n\right)^{-\epsilon}g, \tag{F.29}$$

$$g^\tau=\chi\left(p^\tau\right)^{-\epsilon}g, \tag{F.30}$$

$$g^m=\nu g^\tau, \tag{F.31}$$

$$g^x=\left(1-\nu\right)g^\tau, \tag{F.32}$$

$$\lambda=\phi\frac{\left[\left(1-b\right)c\right]^{-\sigma_c}}{\theta}, \tag{F.33}$$

$$q=p^\tau, \tag{F.34}$$

$$\mu_1 p^\tau=\mu_6, \tag{F.35}$$

$$\mu_1 q=\mu_8, \tag{F.36}$$

$$\mu_1 q=\mu_7, \tag{F.37}$$

$$\mu_{19}=\mu_{15}, \tag{F.38}$$

$$\mu_{15}=\mu_5, \tag{F.39}$$

$$\mu_{17}=p^\tau\mu_1, \tag{F.40}$$

$$\mu_{18}=p^\tau\mu_1, \tag{F.41}$$

$$\mu_{16}=p^\tau\mu_1, \tag{F.42}$$

$$r\frac{d}{y}\left(\alpha_n p^n y^n+\alpha_x p^\tau y^x\right)=\alpha_x p^\tau y^x-\left(1-\alpha_n\right)p^n y^n-\chi\left(p^\tau\right)^{1-\epsilon}\left(c+g\right), \tag{F.43}$$

$$\left(\theta-1\right)c=g, \tag{F.44}$$

$$y^n=\left(\frac{p^n}{p^\tau}\right)^{\frac{1-\alpha_n}{\alpha_n}}\left(1-\alpha_n\right)^{\frac{1-\alpha_n}{\alpha_n}}h^n, \tag{F.45}$$

$$y^x = (1-\alpha_x)^{\frac{1-\alpha_x}{\alpha_x}} h^x, \tag{F.46}$$

$$\theta (h^n)^{1+\varphi} = \phi [(1-b)c]^{-\sigma_c} \alpha_n p^n y^n, \tag{F.47}$$

$$\theta (h^x)^{1+\varphi} = \phi [(1-b)c]^{-\sigma_c} \alpha_x p^\tau y^x, \tag{F.48}$$

$$1 = \chi (p^\tau)^{1-\epsilon} + (1-\chi)(p^n)^{1-\epsilon}, \tag{F.49}$$

$$y^n = (1-\chi)(p^n)^{-\epsilon}(c+g), \tag{F.50}$$

$$-rp^\tau \mu_4 = \mu_{11}\alpha_n p^n y^n + \mu_{12}\alpha_x p^\tau y^x + \mu_3\theta, \tag{F.51}$$

$$\begin{aligned}\phi [(1-b)c]^{-\sigma_c}(1-b\beta) + \mu_2(\theta-1) &= \mu_5(1-\chi)(p^n)^{-\epsilon} + \mu_6\chi(p^\tau)^{-\epsilon} + \\ &+ \mu_3\phi\sigma_c [(1-b)c]^{-\sigma_c-1}(1-b\beta),\end{aligned} \tag{F.52}$$

$$\mu_{11}(1+\varphi)(h^x)^{1+\varphi} = (h^n)^{1+\varphi} + \mu_9 y^n, \tag{F.53}$$

$$\mu_{12}(1+\varphi)(h^x)^{1+\varphi} = (h^x)^{1+\varphi} + \mu_{10} y^x, \tag{F.54}$$

$$\mu_{11}\phi [(1-b)c]^{-\sigma_c}\alpha_n p^n = \theta(\mu_9+\mu_5) - \theta\mu_1(1-\alpha_n)p^n, \tag{F.55}$$

$$\mu_{12}\phi [(1-b)c]^{-\sigma_c}\alpha_x p^\tau = \theta(\alpha_x\mu_1 p^\tau + \mu_{10}), \tag{F.56}$$

$$\begin{aligned}\mu_{11}\phi [(1-b)c]^{-\sigma_c}\alpha_n p^n y^n &= \theta\epsilon\mu_5(1-\chi)(p^n)^{-\epsilon}(c+g) + \theta\mu_{13}(1-\chi)(p^n)^{1-\epsilon} - \\ &- \theta\mu_1(1-\alpha_n)p^n y^n - \theta\mu_9\frac{1-\alpha_n}{\alpha_n}y^n,\end{aligned} \tag{F.57}$$

$$(\epsilon-\eta)\mu_1(c+g) + \mu_{13} = \mu_{14}\frac{(p^\tau)^\epsilon}{\chi}, \tag{F.58}$$

$$\begin{aligned}(\mu_{12}\alpha_x p^\tau y^x + r^* p^\tau \mu_4)\phi [(1-b)c]^{-\sigma_c} &= \theta\eta\mu_1\chi(p^\tau)^{1-\epsilon}(c+g) + \\ &+ \theta\mu_1\left[-r\frac{d}{y}(\alpha_n p^n y^n + \alpha_x p^\tau y^x) + p^\tau\alpha_x y^x - \chi(p^\tau)^{1-\epsilon}(c+g)\right] + \\ &+ \theta\left[\mu_{14}p^\tau + \mu_9\frac{1-\alpha_n}{\alpha_n}y^n\right],\end{aligned} \tag{F.59}$$

$$\mu_1\frac{d}{y}(\alpha_n p^n y^n + \alpha_x p^\tau y^x)\theta = -\mu_4 p^\tau\phi [(1-b)c]^{-\sigma_c}, \tag{F.60}$$

$$(1-\phi)[(1-b)g]^{-\sigma_g}(1-b\beta) = \mu_2 + \mu_{15}(1-\chi)(p^n)^{-\epsilon} + \mu_1\chi(p^\tau)^{1-\epsilon}, \tag{F.61}$$

$$\mu_2\theta c = \mu_3\phi [(1-b)c]^{-\sigma_c}, \tag{F.62}$$

$$g = \frac{g}{y}(\alpha_n p^n y^n + \alpha_x p^\tau y^x). \tag{F.63}$$

F.2 Calibration of Multiple Shocks

We assume that sectoral TFP (a_t^n and a_t^x), the gross foreign interest rate ($1+r_t^*$), the import price ($p_t^{m,\$}$) and the preference ($\phi_t$) shocks follow exogenous processes of the AR(1) form:

$$\ln(\mathrm{x}_t/\mathrm{x}) = \varrho_\mathrm{x}\ln(\mathrm{x}_{t-1}/\mathrm{x}) + \zeta_\mathrm{x}\varepsilon_\mathrm{x} \quad \text{for} \quad \mathrm{x} = \left\{a^n, a^x, 1+r^*, p^{m,\$}, \phi\right\},$$

where x denotes the steady state value of each of the exogenous variables ($a^n = 1$, $a^x = 1$, $1+r^* = 1.04$, $p^{m,\$} = 1$, $\phi = 0.69$). The calibration of the persistence and dispersion of the TFP process follows Fernández et al. (2021). The corresponding values are $\varrho_a = 0.8145$ and $\zeta_a = 0.024$, respectively. The persistence and dispersion of the foreign interest rate are in line with Uribe and Yue (2006), with corresponding values $\rho_r = 0.4746$ and $\zeta_r = 0.01$. We set the persistence of import price and preference shocks to $\varrho_m = \varrho_\phi = 0.9$. We calibrate the dispersion of the import price shock to be half as large as the export price shock in line with Di Pace et al. (2024). We calibrate the dispersion of the preference shock conservatively to target the lower value of the standard deviation of government spending. Unconditional correlations are computed by drawing random innovations for the other shocks and simulating the model for 25,000 observations.

F.3 Financial Autarky

The Ramsey planner chooses the instruments g_t and θ_t so as to maximize the welfare of the representative household. Note that in equilibrium, $\bar{d}_t = d_t$, $\bar{c}_t = c_t$ and $\bar{g}_t = g_t$. This is not internalized by households but internalized by the Ramsey planner. Financial autarky requires that:

$$d_t + d_t^g = 0.$$

This implies that there is one market for debt and that the interest rate (r_t) clears this market (i.e., the interest rate is no longer exogenously determined). This also implies that market clearing in the tradable goods market is given by:

$$\alpha_x q_t p_t^{x,\$} y_t^x = (1-\alpha_n) p_t^n y_t^n + q_t p_t^{x,\$} \left(c_t^x + g_t^x\right) + q_t p_t^{m,\$} \left(c_t^m + g_t^m\right).$$

Given the equilibrium condition in the asset market, the government budget constraint can be therefore re-written as,

$$d_t - (1+r_{t-1}) d_{t-1} = (\theta_t - 1) c_t - g_t + \Pi_t - \Phi(d_t).$$

In studying this problem, we assume that the representative households faces a quadratic portfolio transaction cost, $\Phi(d_t) = \frac{\psi}{2}\left(d_t - \bar{d}\right)^2$, and that these services are provided by a government agency at zero administrative cost.[36] Profits are transferred to households in order to remove any associated wealth effects; i.e., $\Pi_t = \Phi(d_t)$. In turn, the portfolio costs are internalized by both the household and the Ramsey planner. This is different to the incomplete market case in which only the Ramsey planner internalizes the upward sloping supply of funds.

The Ramsey planner chooses λ_t, r_t, c_t, c_t^τ, c_t^m, c_t^x, h_t^n, h_t^x, y_t^n, y_t^x, p_t^n, p_t^τ, q_t, d_t, g_t, g_t^τ, g_t^m, g_t^x and θ_t to maximize:

$$\begin{aligned}
\mathcal{L} =& \mathbb{E}_0 \sum_{t=0}^{\infty} \beta^t \left[\phi_t \frac{(c_t - bc_{t-1})^{1-\sigma_c} - 1}{1-\sigma_c} + (1-\phi_t) \frac{(g_t - bg_{t-1})^{1-\sigma_g} - 1}{1-\sigma_g} - \sum_{j=\{n,x\}} \frac{\left(h_t^j\right)^{1+\varphi}}{1+\varphi} \right] + \\
& \mathbb{E}_0 \sum_{t=0}^{\infty} \beta^t \mu_{1t} \left[\alpha_x q_t p_t^{x,\$} y_t^x - (1-\alpha_n) p_t^n y_t^n - q_t p_t^{x,\$}\left(c_t^x + g_t^x\right) - q_t p_t^{m,\$}\left(c_t^m + g_t^m\right) \right] +
\end{aligned}$$

[36] The parameter for the portfolio adjustment cost is fixed at 0.24, which makes the value of the borrowing spread nearly twice as large as in the baseline model. All remaining parameters are selected to align with the specifications of the baseline model.

$$\mathbb{E}_0 \sum_{t=0}^{\infty} \beta^t \mu_{2t} \left[(1 + r_{t-1}) d_{t-1} - d_t + (\theta_t - 1) c_t - g_t \right] +$$
$$\mathbb{E}_0 \sum_{t=0}^{\infty} \beta^t \mu_{3t} \left[\phi_t (c_t - b c_{t-1})^{-\sigma_c} - \theta_t \lambda_t \right] +$$
$$\mathbb{E}_0 \sum_{t=0}^{\infty} \beta^t \mu_{4t} \left[\lambda_t - \beta \frac{(1 + r_t)}{1 - \psi (d_t - d)} \lambda_{t+1} \right] +$$
$$\mathbb{E}_0 \sum_{t=0}^{\infty} \beta^t \mu_{5t} \left[c_t^n - (1 - \chi) (p_t^n)^{-\epsilon} c_t \right] + \mathbb{E}_0 \sum_{t=0}^{\infty} \beta^t \mu_{6t} \left[c_t^\tau - \chi (p_t^\tau)^{-\epsilon} c_t \right] +$$
$$\mathbb{E}_0 \sum_{t=0}^{\infty} \beta^t \mu_{7t} \left[c_t^m - \nu \left(\frac{q_t p_t^{m,\$}}{p_t^\tau} \right)^{-\eta} c_t^\tau \right] + \mathbb{E}_0 \sum_{t=0}^{\infty} \beta^t \mu_{8t} \left[c_t^x - (1 - \nu) \left(\frac{q_t p_t^{x,\$}}{p_t^\tau} \right)^{-\eta} c_t^\tau \right] +$$
$$\mathbb{E}_0 \sum_{t=0}^{\infty} \beta^t \mu_{9t} \left[y_t^n - \left(\frac{p_t^n}{q_t p_t^{m,\$}} \right)^{\frac{1-\alpha_n}{\alpha_n}} (1 - \alpha_n)^{\frac{1-\alpha_n}{\alpha_n}} (a_t^n)^{\frac{1}{\alpha_n}} h_t^n \right] +$$
$$\mathbb{E}_0 \sum_{t=0}^{\infty} \beta^t \mu_{10t} \left[y_t^x - \left(\frac{p_t^{x,\$}}{p_t^{m,\$}} \right)^{\frac{1-\alpha_x}{\alpha_x}} (1 - \alpha_x)^{\frac{1-\alpha_x}{\alpha_x}} (a_t^x)^{\frac{1}{\alpha_x}} h_t^x \right] +$$
$$\mathbb{E}_0 \sum_{t=0}^{\infty} \beta^t \mu_{11t} \left[(h_t^n)^{1+\varphi} - \lambda_t \alpha_n p_t^n y_t^n \right] + \mathbb{E}_0 \sum_{t=0}^{\infty} \beta^t \mu_{12t} \left[(h_t^x)^{1+\varphi} - \lambda_t \alpha_x q_t p_t^{x,\$} y_t^x \right] +$$
$$\mathbb{E}_0 \sum_{t=0}^{\infty} \beta^t \mu_{13t} \left[\left(\chi (p_t^\tau)^{1-\epsilon} + (1 - \chi) (p_t^n)^{1-\epsilon} \right)^{\frac{1}{1-\epsilon}} - 1 \right] +$$
$$\mathbb{E}_0 \sum_{t=0}^{\infty} \beta^t \mu_{14t} \left[\left(\nu \left(p_t^{m,\$} \right)^{1-\eta} + (1 - \nu) \left(p_t^{x,\$} \right)^{1-\eta} \right)^{\frac{1}{1-\eta}} q_t - p_t^\tau \right] +$$
$$\mathbb{E}_0 \sum_{t=0}^{\infty} \beta^t \mu_{15t} \left[g_t^n - (1 - \chi) (p_t^n)^{-\epsilon} g_t \right] + \mathbb{E}_0 \sum_{t=0}^{\infty} \beta^t \mu_{16t} \left[g_t^\tau - \chi (p_t^\tau)^{-\epsilon} g_t \right] +$$
$$\mathbb{E}_0 \sum_{t=0}^{\infty} \beta^t \mu_{17t} \left[g_t^m - \nu \left(\frac{q_t p_t^{m,\$}}{p_t^\tau} \right)^{-\eta} g_t^\tau \right] + \mathbb{E}_0 \sum_{t=0}^{\infty} \beta^t \mu_{18t} \left[g_t^x - (1 - \nu) \left(\frac{q_t p_t^{x,\$}}{p_t^\tau} \right)^{-\eta} g_t^\tau \right] +$$
$$\mathbb{E}_0 \sum_{t=0}^{\infty} \beta^t \mu_{19t} \left[y_t^n - c_t^n - g_t^n \right].$$

The first order conditions are:

$$\lambda_t: \quad \mu_{4t} = \frac{(1 + r_{t-1})}{1 - \psi (d_{t-1} - d)} \mu_{4t-1} + \mu_{11t} \alpha_n p_t^n y_t^n + \mu_{12t} \alpha_x q_t p_t^{x,\$} y_t^x + \mu_{3t} \theta_t, \tag{F.64}$$
$$c_t: \quad \mu_{5t} c_t^n + \mu_{6t} c_t^\tau + \sigma_c c_t \left[\mu_{3t} \phi_t (c_t - b c_{t-1})^{-\sigma_c - 1} - \mathbb{E}_t \mu_{3t+1} \phi_{t+1} (c_{t+1} - b c_t)^{-\sigma_c - 1} \right] =$$
$$= c_t \left[\phi_t (c_t - b c_{t-1})^{-\sigma_c} - \beta b \mathbb{E}_t \phi_{t+1} (c_{t+1} - b c_t)^{-\sigma_c} \right] + \mu_{2t} (\theta_t - 1) c_t, \tag{F.65}$$
$$c_t^n: \quad \mu_{19t} = \mu_{5t}, \tag{F.66}$$
$$c_t^\tau: \quad \mu_{7t} c_t^m + \mu_{8t} c_t^x = \mu_{6t} c_t^\tau, \tag{F.67}$$
$$c_t^m: \quad \mu_{1t} q_t p_t^{m,\$} = \mu_{7t}, \tag{F.68}$$
$$c_t^x: \quad \mu_{1t} q_t p_t^{x,\$} = \mu_{8t}, \tag{F.69}$$

$$h_t^n: \quad \mu_{11t}\left(1+\varphi\right)\left(h_t^x\right)^{1+\varphi} = \left(h_t^n\right)^{1+\varphi} + \mu_{9t} y_t^n, \tag{F.70}$$

$$h_t^x: \quad \mu_{12t}\left(1+\varphi\right)\left(h_t^x\right)^{1+\varphi} = \left(h_t^x\right)^{1+\varphi} + \mu_{10t} y_t^x, \tag{F.71}$$

$$y_t^n: \quad \mu_{11t}\lambda_t\alpha_n p_t^n + \mu_{1t}\left(1-\alpha_n\right)p_t^n = \mu_{9t} + \mu_{19t}, \tag{F.72}$$

$$y_t^x: \quad \mu_{12t}\lambda_t\alpha_x q_t p_t^{x,\$} = \mu_{1t}\alpha_x q_t p_t^{x,\$} + \mu_{10t}, \tag{F.73}$$

$$p_t^n: \quad \epsilon\left(\mu_{5t}c_t^n + \mu_{15t}g_t^n\right) + \mu_{13t}(1-\chi)\left(p_t^n\right)^{1-\epsilon} = \mu_{1t}\left(1-\alpha_n\right)p_t^n y_t^n + $$
$$+ \mu_{11t}\lambda_t\alpha_n p_t^n y_t^n + \mu_{9t}\frac{1-\alpha_n}{\alpha_n}y_t^n, \tag{F.74}$$

$$p_t^\tau: \quad \epsilon\left(\mu_{6t}c_t^\tau + \mu_{16t}g_t^\tau\right) + \mu_{13t}\chi\left(p_t^\tau\right)^{1-\epsilon} = \eta\left[\mu_{7t}c_t^m + \mu_{8t}c_t^x + \mu_{17t}g_t^m + \mu_{18t}g_t^x\right] +$$
$$+ \mu_{14t}p_t^\tau, \tag{F.75}$$

$$q_t: \quad \mu_{12t}\lambda_t\alpha_x q_t p_t^{x,\$}y_t^x = \eta\left[\mu_{7t}c_t^m + \mu_{8t}c_t^x + \mu_{17t}g_t^m + \mu_{18t}g_t^x\right] + \tag{F.76}$$

$$+ \mu_{1t}q_t\left\{\alpha_x p_t^{x,\$}y_t^x - p_t^{x,\$}\left(c_t^x + g_t^x\right) - p_t^{m,\$}\left(c_t^m + g_t^m\right)\right\} + \mu_{14t}p_t^\tau + \mu_{9t}\frac{1-\alpha_n}{\alpha_n}y_t^n, \tag{F.77}$$

$$d_t: \quad \beta\mu_{2t+1}\left(1+r_t\right) = \mu_{2t} + \mu_{4t}\beta\psi\frac{\left(1+r_t\right)}{\left[1-\psi\left(d_t-d\right)\right]^2}\lambda_{t+1}, \tag{F.78}$$

$$r_t: \quad \mu_{2t+1}\beta d_t = \mu_{4t}\beta\frac{1}{1-\psi\left(d_t-d\right)}\lambda_{t+1}, \tag{F.79}$$

$$g_t: \quad \left[\left(1-\phi_t\right)\left(g_t - bg_{t-1}\right)^{-\sigma_g} - \beta b\mathbb{E}_t\left(1-\phi_{t+1}\right)\left(g_{t+1}-bg_t\right)^{-\sigma_g}\right]g_t = \mu_{2t}g_t +$$
$$+ \mu_{15t}g_t^n + \mu_{16t}g_t^\tau, \tag{F.80}$$

$$g_t^n: \quad \mu_{15t} = \mu_{19t}, \tag{F.81}$$

$$g_t^\tau: \quad \mu_{16t}g_t^\tau = \mu_{17t}g_t^m + \mu_{18t}g_t^x, \tag{F.82}$$

$$g_t^m: \quad q_t p_t^{m,\$}\mu_{1t} = \mu_{17t}, \tag{F.83}$$

$$g_t^x: \quad q_t p_t^{x,\$}\mu_{1t} = \mu_{18t}, \tag{F.84}$$

$$\theta_t: \quad \mu_{2t}c_t = \mu_{3t}\lambda_t. \tag{F.85}$$

And 19 Lagrange multipliers with their associated implementability conditions.

F.3.1 Steady State

The steady state can be summarized by the following set of equations:

$$c^n = (1-\chi)\left(p^n\right)^{-\epsilon}c, \tag{F.86}$$

$$c^\tau = \chi\left(p^\tau\right)^{-\epsilon}c, \tag{F.87}$$

$$c^m = \nu c^\tau, \tag{F.88}$$

$$c^x = (1-\nu)\,c^\tau, \tag{F.89}$$

$$g^n = (1-\chi)\left(p^n\right)^{-\epsilon}g, \tag{F.90}$$

$$g^\tau = \chi\left(p^\tau\right)^{-\epsilon}g, \tag{F.91}$$

$$g^m = \nu g^\tau, \tag{F.92}$$

$$g^x = (1-\nu)\,g^\tau, \tag{F.93}$$

$$\lambda = \phi\frac{\left[(1-b)\,c\right]^{-\sigma_c}}{\theta}, \tag{F.94}$$

$$q = p^{\tau}, \tag{F.95}$$

$$\mu_5 = \mu_{19}, \tag{F.96}$$

$$\mu_6 = \mu_1 p^{\tau}, \tag{F.97}$$

$$\mu_7 = \mu_1 p^{\tau}, \tag{F.98}$$

$$\mu_8 = \mu_1 p^{\tau}, \tag{F.99}$$

$$\mu_{19} = \mu_{15}, \tag{F.100}$$

$$\mu_{17} = p^{\tau}\mu_1, \tag{F.101}$$

$$\mu_{18} = p^{\tau}\mu_1, \tag{F.102}$$

$$\mu_{16} = p^{\tau}\mu_1, \tag{F.103}$$

$$1 - \psi\left[d - \frac{d}{y}\left(\alpha_n p^n y^n + \alpha_x p^x y^x\right)\right] = \beta\left(1 + r\right), \tag{F.104}$$

$$\alpha_x p^{\tau} y^x = \chi\left(p^{\tau}\right)^{1-\epsilon}\left(c + g\right) + \left(1 - \alpha_n\right) p^n y^n, \tag{F.105}$$

$$rd = \left(\theta - 1\right) c - g, \tag{F.106}$$

$$y^n = \left(\frac{p^n}{p^{\tau}}\right)^{\frac{1-\alpha_n}{\alpha_n}}\left(1 - \alpha_n\right)^{\frac{1-\alpha_n}{\alpha_n}} h^n, \tag{F.107}$$

$$y^x = \left(1 - \alpha_x\right)^{\frac{1-\alpha_x}{\alpha_x}} h^x, \tag{F.108}$$

$$\theta\left(h^n\right)^{1+\varphi} = \phi\left[\left(1 - b\right) c\right]^{-\sigma_c} \alpha_n p^n y^n, \tag{F.109}$$

$$\theta\left(h^x\right)^{1+\varphi} = \phi\left[\left(1 - b\right) c\right]^{-\sigma_c} \alpha_x p^{\tau} y^x, \tag{F.110}$$

$$1 = \chi\left(p^{\tau}\right)^{1-\epsilon} + \left(1 - \chi\right)\left(p^n\right)^{1-\epsilon}, \tag{F.111}$$

$$y^n = \left(1 - \chi\right)\left(p^n\right)^{1-\epsilon}\left(c + g\right), \tag{F.112}$$

$$-\mu_4 r = \mu_{11}\alpha_n p^n y^n + \mu_{12}\alpha_x p^{\tau} y^x + \mu_3\theta, \tag{F.113}$$

$$\begin{aligned}\phi\left[\left(1 - b\right) c\right]^{-\sigma_c}\left(1 - b\beta\right) + \mu_2\left(\theta - 1\right) &= \mu_5\left(1 - \chi\right)\left(p^n\right)^{-\epsilon} + \mu_6\chi\left(p^{\tau}\right)^{-\epsilon} + \\ &+ \mu_{13} + \mu_3\phi\sigma_c\left[\left(1 - b\right) c\right]^{-\sigma_c - 1}\left(1 - b\beta\right),\end{aligned} \tag{F.114}$$

$$\mu_{11}\left(1 + \varphi\right)\left(h^x\right)^{1+\varphi} = \left(h^n\right)^{1+\varphi} + \mu_9 y^n, \tag{F.115}$$

$$\mu_{12}\left(1 + \varphi\right)\left(h^x\right)^{1+\varphi} = \left(h^x\right)^{1+\varphi} + \mu_{10} y^x, \tag{F.116}$$

$$\mu_{11}\phi\left[\left(1 - b\right) c\right]^{-\sigma_c}\alpha_n p^n + \theta\mu_1\left(1 - \alpha_n\right) p^n = \theta\left(\mu_9 + \mu_5\right), \tag{F.117}$$

$$\mu_{12}\phi\left[\left(1 - b\right) c\right]^{-\sigma_c}\alpha_x p^{\tau} = \theta\left(\alpha_x\mu_1 p^{\tau} + \mu_{10}\right), \tag{F.118}$$

$$\begin{aligned}\mu_{11}\alpha_n p^n\phi\left[\left(1 - b\right) c\right]^{-\sigma_c} y^n &= \theta\epsilon\left(1 - \chi\right)\left(p^n\right)^{-\epsilon}\mu_5\left(c + g\right) + \\ &+ \theta\mu_{13}\left(1 - \chi\right)\left(p^n\right)^{(1-\epsilon)}\end{aligned} \tag{F.119}$$

$$- \theta\mu_1\left(1 - \alpha_n\right) p^n y^n + \theta\mu_9\frac{\left(1 - \alpha_n\right)}{\alpha_n} y^n, \tag{F.120}$$

$$\mu_{13} = \left(\eta - \epsilon\right)\mu_1\left(c + g\right) + \frac{\mu_{14}\left(p^{\tau}\right)^{\epsilon}}{\chi}, \tag{F.121}$$

$$\begin{aligned}\mu_{12}\alpha_x p^{\tau} y^x\phi\left[\left(1 - b\right) c\right]^{-\sigma_c} &= \theta\mu_1\eta\chi\left(p^{\tau}\right)^{(1-\epsilon)}\left(c + g\right) + \\ &+ \theta\mu_1\left(p^{\tau}\alpha_x y^x - \chi\left(p^{\tau}\right)^{1-\epsilon}\left(c + g\right)\right)\end{aligned} \tag{F.122}$$

$$+ \theta\left(\mu_{14} p^{\tau} + \mu_9\frac{\left(1 - \alpha_n\right)}{\alpha_n} y^n\right), \tag{F.123}$$

$$\mu_2 \left[\beta \left(1+r\right)-1\right] \theta = \mu_{4t} \psi \phi \left[\left(1-b\right) c\right]^{-\sigma_c}, \tag{F.124}$$

$$\mu_2 d\theta \beta \left(1+r\right) = \mu_4 \phi \left[\left(1-b\right) c\right]^{-\sigma_c} \left[1-\beta \left(1+r\right)\right], \tag{F.125}$$

$$\left(1-\phi\right) \left[\left(1-b\right) g\right]^{-\sigma_g} \left(1-b\beta\right) = \mu_2 + \mu_{15} \left(1-\chi\right) \left(p^n\right)^{-\epsilon} + \mu_1 \chi \left(p^\tau\right)^{1-\epsilon}, \tag{F.126}$$

$$\mu_2 \theta c = \mu_3 \phi \left[\left(1-b\right) c\right]^{-\sigma_c}, \tag{F.127}$$

$$g = \frac{g}{y} \left(\alpha_n p^n y^n + \alpha_x p^\tau y^x\right). \tag{F.128}$$

G Results under Alternative Parametrizations

This Section assesses the robustness of our results to alternative specifications of the baseline model through three distinct sensitivity analyses. First, we examine the impact of varying the preference parameter for government consumption. Second, we investigate the effects of altering shock persistence values. Third, we present the results for the composition channel. A summary of this exercise is described in Section 6.3 of the main text.

G.1 Varying Preference for Government Consumption

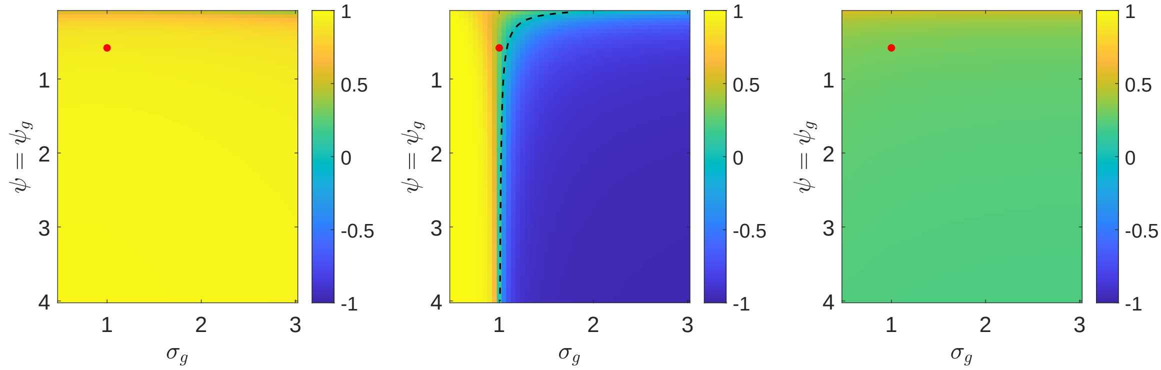
We focus on varying the intertemporal elasticity of substitution of government spending (σ_g) relative to that of private consumption (σ_c), which is held constant. Additionally, we explore the impact of changes in the degree of asset market incompleteness ($\psi = \psi_g$). All other parameters are held constant throughout the analysis. To ensure consistency in the size of the government across different calibrations, we adjust the parameter ϕ to target a value of 0.3. These parameter variations affect two main channels shaping optimal fiscal policy: consumption smoothing and consumption preference (Fernández et al., 2021; Riascos and Végh, 2003).

Let us start by defining procyclicality of government spending as a positive conditional correlation between output and government spending, denoted as $\texttt{Corr}\left(\mathcal{G}, \mathcal{Y} | p^{x,\$}\right) > 0$. The procyclicality of the tax rate are denoted as $\texttt{Corr}\left(\theta, \mathcal{Y} | p^{x,\$}\right) < 0$. As in the baseline, we keep the value of σ_c to 1. We draw a series of random export price shocks and compute the conditional correlations between fiscal variables and output for different values of σ_g (ranging from 0.5 to 3) and $\psi = \psi_g$ (from 0.1 to 4). Each time we vary the value of σ_g, we recompute the steady state.

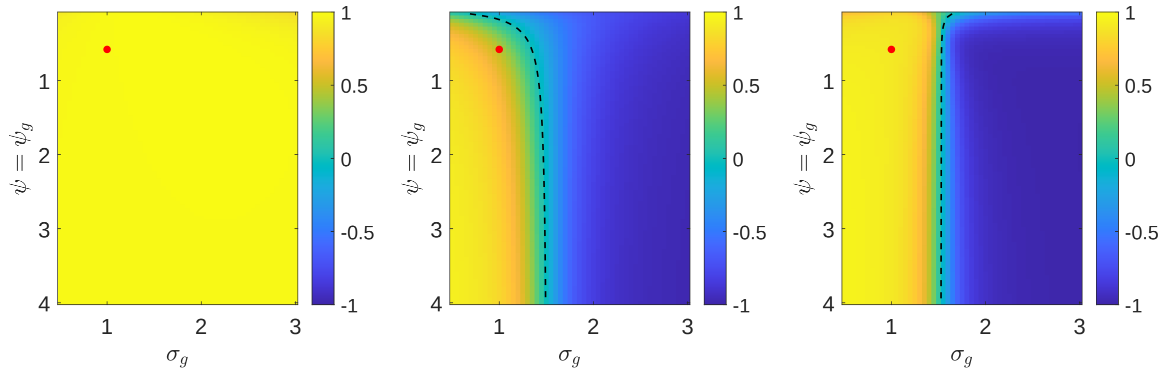
Figure G.1 shows the conditional correlations between various fiscal variables and output in response to export price shocks. The color scheme represents correlation strength and direction: yellow for positive, blue for negative, and green for near-zero correlations. In Panels (a) and (d) we observe that the correlations of goverment spening and revenue with respect to GDP are consistently high and positive across all calibrations. An expansionary shock to the price of exports increases the tax base, leading to higher fiscal revenues, regardless of the movements in the tax rate. This allows the government to increase spending while simultaneously maintaining a primary surplus.

In line with Fernández et al. (2021), the consumption preference channel, which depends on the relative values of σ_c and σ_g, is crucial in determining the cyclicality of the tax rate,

Figure G.1: Heatmaps of Fiscal Cyclicality: Preference for Government Consumption



(a) Corr $\left(\mathcal{G}, \mathcal{Y}|p^{x,\$}\right)$ **(b)** Corr $\left(\theta, \mathcal{Y}|p^{x,\$}\right)$ **(c)** Corr $\left(\text{PB}, \mathcal{Y}|p^{x,\$}\right)$



(d) Corr $\left(\mathcal{R}, \mathcal{Y}|p^{x,\$}\right)$ **(e)** Corr $\left(\mathcal{G}/\mathcal{Y}, \mathcal{Y}|p^{x,\$}\right)$ **(f)** Corr $\left(\mathcal{R}/\mathcal{Y}, \mathcal{Y}|p^{x,\$}\right)$

Notes: The subfigures **(a)**, **(b)** and **(c)** show the conditional correlations between GDP and: government spending ($\mathcal{G}$), the tax rate (θ), and the primary balance (PB),respectively, for different parametrizations of the (relative) curvature and financial frictions under the baseline. The subfigures **(d)**, **(e)** and **(f)** report the conditional correlations for other fiscal outcomes: revenue($\mathcal{R}$), spending over GDP and revenue over GDP. In all these experiments, we set $\sigma_c = 1$ and loop over the values of σ_g in the interval $[0.5, 3]$ and $\psi = \psi_g$ in the interval $[0.1, 4]$. The red dot refers to the baseline calibration and the dotted black line results denotes the locus where the conditional correlation is zero. The graph uses a color-coded scheme to represent correlations between fiscal variables and output. Yellow indicates a positive correlation, blue denotes a negative correlation, and green represents a near-zero correlation.

Panel (d). When $\sigma_g > \sigma_c$, households' preferences imply a stronger willingness to smooth public over private consumption. As a result, private consumption increases proportionately more than public consumption following an increase in the price of exports. The relatively higher tax base allows the fiscal authority to reduce the tax rate while still financing the higher public consumption, thus engaging in procyclical tax policy. Conversely, when $\sigma_g < \sigma_c$, households have a stronger preference for smoothing private over public consumption. In this scenario, private consumption increases proportionately less than public consumption in favorable economic conditions. If the fiscal authority maintains the same tax rate, tax revenues would fall short. The need to finance higher public consumption necessitates higher taxes following an expansionary shock, resulting in a countercyclical tax policy.

By the same token, for large values of σ_g, the increase in public spending is limited relative to the expansion in economic activity, leading to a countercyclical spending-to-GDP ratio following an increase in the price of exports. Requiring less revenue to cover the increase in spending means that the revenue-to-GDP ratio also falls following the same shock. These effects are amplified by greater asset market incompleteness (i.e., for large values of $\psi = \psi_g$), as spending decisions become more tightly linked to the government's ability to raise revenue through taxes.

G.2 Varying Shock Persistence

The intertemporal consumption smoothing channel is crucial in determining optimal policy in our framework. This channel is fundamentally affected by two model features: (i) shock persistence and (ii) a country's access to international financial markets, proxied by the elasticity of the spread with respect to debt. While our baseline calibration matches the export price shock persistence to empirical evidence, we now explore how variations in shock persistence together with the spread-to-debt elasticity influence optimal fiscal responses.

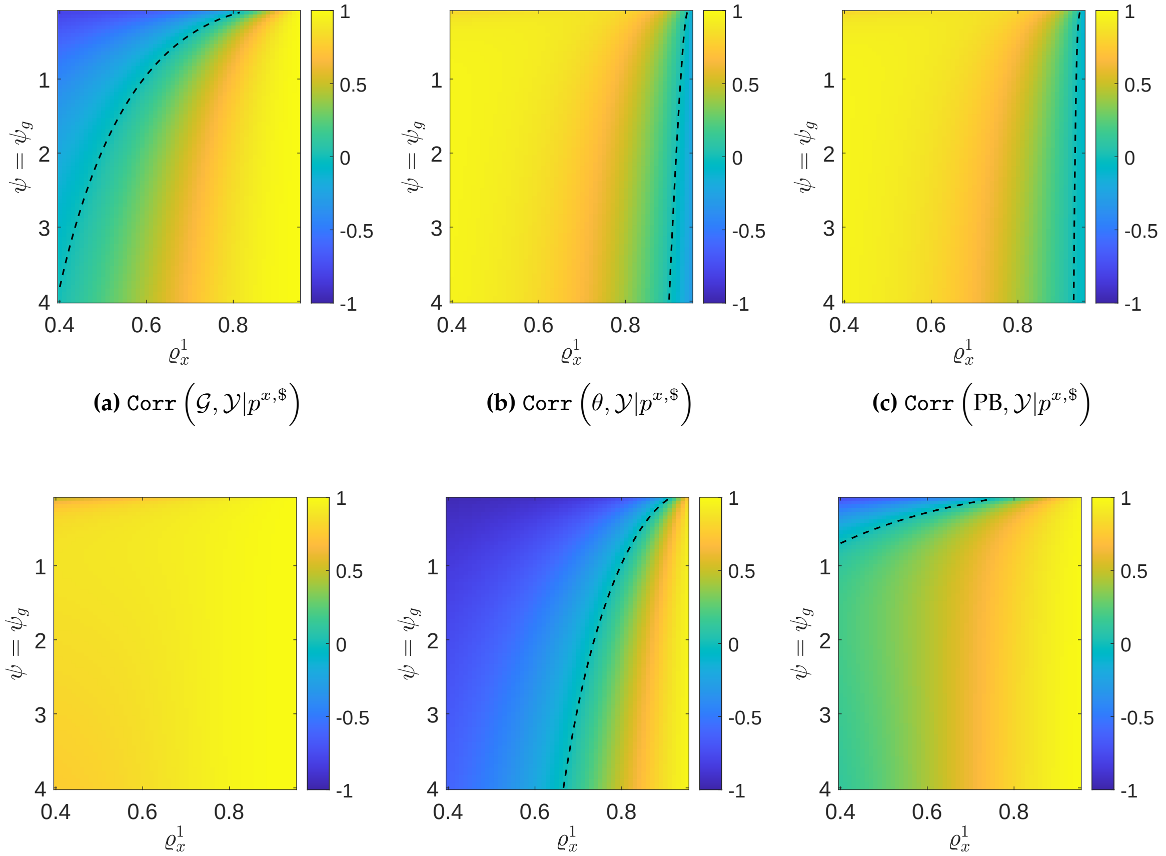
Specifically, we loop over values of the persistence parameter of the export price shock ϱ_x^1 in the interval $[0.4, 0.95]$ (and set $\varrho_x^2 = 0$) and, as in the previous exercise, $\psi = \psi_g$ (from 0.1 to 4). For each parameter pair, we compute the correlation between a fiscal indicator and GDP, reported in Figure G.2. The results underscore the critical role of commodity price shock persistence in shaping fiscal policy responses under the Ramsey policy.

Transitory shocks lead to a prolonged increase in spending, which requires a primary surplus, and therefore a $\texttt{Corr}\left(\text{PB}, \mathcal{Y}|p^{x,\$}\right) > 0$, unless the shock is close to being permanent in which case the consumption smoothing ceases to exist. Less persistent shocks tend to reduce the procyclicality of government spending (Panel a). As shock persistence declines, the response of private consumption becomes more muted, which in turn leads to a weaker fiscal response; see for example Mendes and Pennings (2020). As the shock becomes less persistent, government spending may shift from being procyclical to being more acyclical. In the case where ψ is very high, the limit to smoothing (consumption and) spending, makes spending more tightly related to the variation in revenues, therefore increasing its procyclicality.

With symmetric preferences ($\sigma_g = \sigma_c$), consumption and government spending move in tandem. Therefore, the base effect alone is insufficient to lead to public saving, and in order to obtain a positive primary balance, the government needs to pursue countercyclical tax policies. Therefore, as in the baseline case, changes in taxes follow a similar pattern to the one of the primary balance (i.e. $\texttt{Corr}\left(\theta, \mathcal{Y}|p^{x,\$}\right) > 0$ unless the shock becomes extremely persistent, Panel b). The combination of countercyclical tax changes and base effects means that for any level of persistence of the shock, we would observe that $\texttt{Corr}\left(\mathcal{R}, \mathcal{Y}|p^{x,\$}\right) > 0$ (Panel d).

The cyclicality of fiscal ratios also changes with shock persistence (Panels e and f). The government spending-to-GDP ratio becomes more countercyclical, as the increase in spending fails to outpace that of GDP, when the shock persistence decreases, but less so when saving is impaired (i.e. when $\psi = \psi_g$ is large). The revenue-to-GDP ratio generally maintains a positive correlation with the business cycle, becoming acyclical only under very low shock

Figure G.2: Heatmaps of Fiscal Cyclicality: Shock Persistence



(a) Corr $\left(\mathcal{G}, \mathcal{Y}|p^{x,\$}\right)$ **(b)** Corr $\left(\theta, \mathcal{Y}|p^{x,\$}\right)$ **(c)** Corr $\left(\text{PB}, \mathcal{Y}|p^{x,\$}\right)$

(d) Corr $\left(\mathcal{R}, \mathcal{Y}|p^{x,\$}\right)$ **(e)** Corr $\left(\mathcal{G}/\mathcal{Y}, \mathcal{Y}|p^{x,\$}\right)$ **(f)** Corr $\left(\mathcal{R}/\mathcal{Y}, \mathcal{Y}|p^{x,\$}\right)$

Notes: The subfigures **(a)**, **(b)** and **(c)** show the conditional correlations between GDP and: government spending ($\mathcal{G}$), the tax rate (θ), and the primary balance (PB), respectively, for different parametrizations of financial frictions and the persistence of the export price shock in the baseline model. The subfigures **(d)**, **(e)** and **(f)** denote the conditional correlations for other fiscal outcomes: revenue($\mathcal{R}$), spending over GDP and revenue over GDP. In all these experiments, we set $\sigma_c = \sigma_g = 1$ and loop over the values of ϱ_x^1 (setting $\varrho_x^2 = 0$) in the interval $0.4, 0.95]$ and $\psi = \psi_g$ in the interval $[0.1, 4]$. The dotted black line results denotes the locus where the conditional correlation is zero. The graph uses a color-coded scheme to represent correlations between fiscal variables and output. Yellow indicates a positive correlation, blue denotes a negative correlation, and green represents a near-zero correlation.

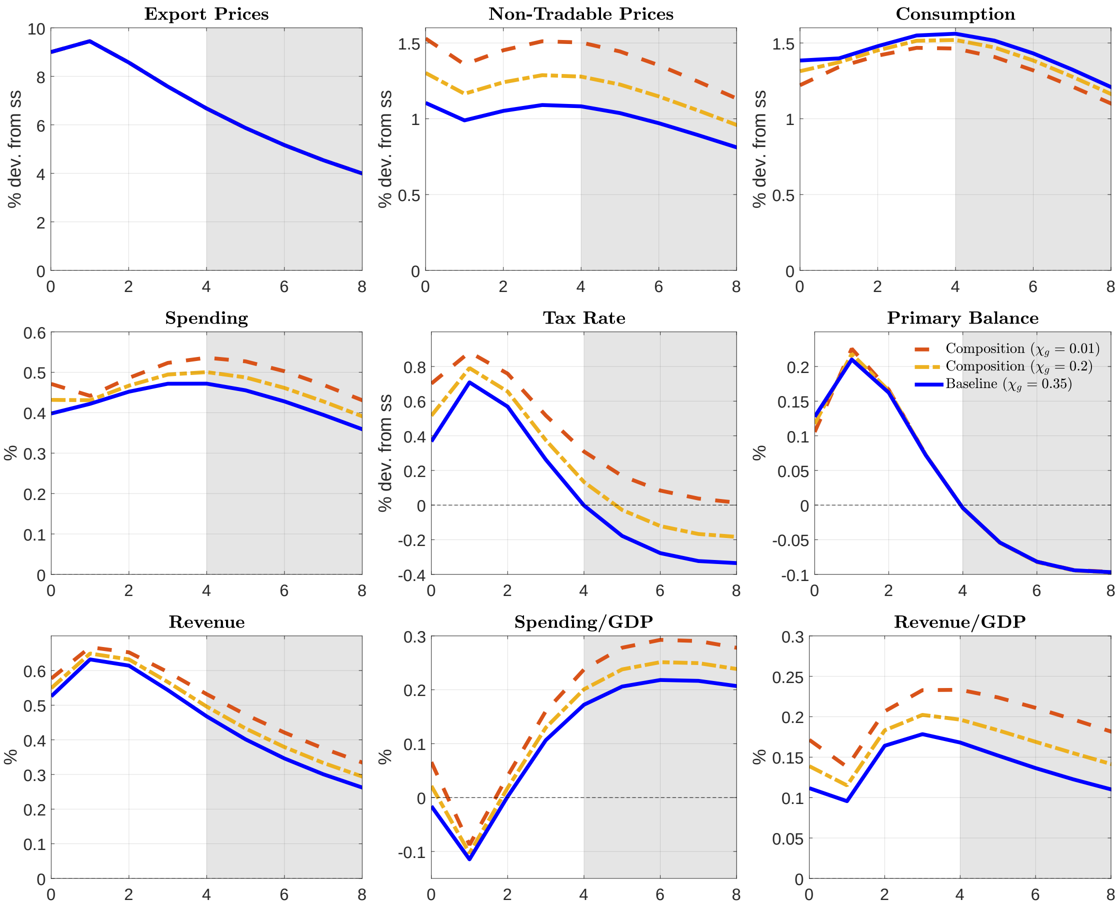
persistence, where revenue movements are roughly proportional to output fluctuations.

G.3 The Composition Channel

This section explores the relevance of the *composition channel*, which occurs when the basket of government spending is tilted towards non-tradable goods. Figure G.3 illustrates the responses of macro and fiscal variables to a one standard deviation export price shock under different model calibrations for χ_g, where $\chi_g < \chi_c$.

Specifically, modeling government consumption as a CES bundle with $\chi_g \neq \chi$ representing the share of tradable goods, the relative price of government spending diverges from that of private consumption, $p_t^g \neq p_t^c = 1$. As the share of non-tradables $(1 - \chi_g)$ in the aggregator approaches unity, the relative price of government spending converges to that of

Figure G.3: Impulse Responses: Composition Channel



Notes: This Figure illustrates the responses of macro and fiscal variables to a one standard deviation export price shock under different model calibrations. The solid blue line represents the baseline model, where the share of tradables for both government spending and consumption ($\chi_g = \chi_c = \chi$) is equal to 0.35. To demonstrate the composition channel, we present two alternative calibrations of χ_g, where $\chi_g < \chi_c$, indicating a greater share of non-tradables in government spending. The dashed orange line shows an extreme case where $\chi_g = 0.01$, while the dashed yellow line represents an intermediate case where $\chi_g = 0.2$. In these alternative scenarios, we maintain the same overall non-tradable share as in the baseline model. The macroeconomic aggregates plotted are observationally equivalent counterparts (expressed in constant prices). Non-tradable prices, consumption, and the tax rate are reported in percentage deviations from steady state values. All remaining responses are in percentage points. Gray areas denote the periods outside of the empirical horizon. Horizontal axes denote years.

non-tradable prices, $\lim_{\chi_g \to 0} p_t^g = p_t^n$.

When government spending is tilted towards non-tradable goods, its demand increases relative to the baseline, pushing up (by more than under the baseline case) the price of non-tradable goods and consequently the relative price of government spending. A higher relative price of public goods provision increases the overall level of spending, expressed in terms of consumption ($p_t^g g_t$), in absolute terms and relative to private consumption. Therefore, with the increase in revenues from the base effects being attenuated, the tax rate must increase (more than under the baseline) to fund this additional expenditure, which in turn depresses private consumption, weakening the base effect further.

Overall, the qualitative responses are in line with the baseline. The introduction of a composition channel accentuates the procyclicality of government spending while making the tax rate more countercyclical.

H Optimal Fiscal Rules

In this Section, we provide additional details on the model with fiscal rules. The first subsection details the computational approach for optimal fiscal rules. The second presents additional results of the model.

H.1 Computation of Welfare Costs

Policy evaluations are conducted by calculating the welfare cost of a fiscal policy regime compared to the time-invariant stochastic equilibrium under the Ramsey policy. The decentralized economy calibration matches the Ramsey problem, and we solve the model to a second-order approximation, focusing only on exogenous export price fluctuations.

For a policy rule regime to be implementable, it must ensure local uniqueness of the rational expectations equilibrium. The coefficients of optimal policy rules, in equations (12)-(13), are chosen so that the contingent plan for consumption, spending, and sectoral hours worked yields the highest (expected) lifetime utility:

$$\Omega_t = U_t + \beta \mathbb{E}_t \Omega_{t+1}, \tag{H.1}$$

where U_t is specified in equation (E.2).[37]

Define the welfare associated with the time-invariant stochastic allocation under the Ramsey policy, conditional on a specific state of the economy at period 0 as Ω_0^R, and the conditional welfare associated with policy regime A as Ω_0^A, where

$$\Omega_0^{\mathfrak{Z}} = \mathbb{E}_0 \sum_{t=0}^{\infty} \beta^t U\left(c_t^{\mathfrak{Z}}, c_{t-1}^{\mathfrak{Z}}, g_t^{\mathfrak{Z}}, g_{t-1}^{\mathfrak{Z}}, h_t^{x,\mathfrak{Z}}, h_t^{n,\mathfrak{Z}}\right) \quad \text{for} \quad \mathfrak{Z} = \{R, A\}. \tag{H.2}$$

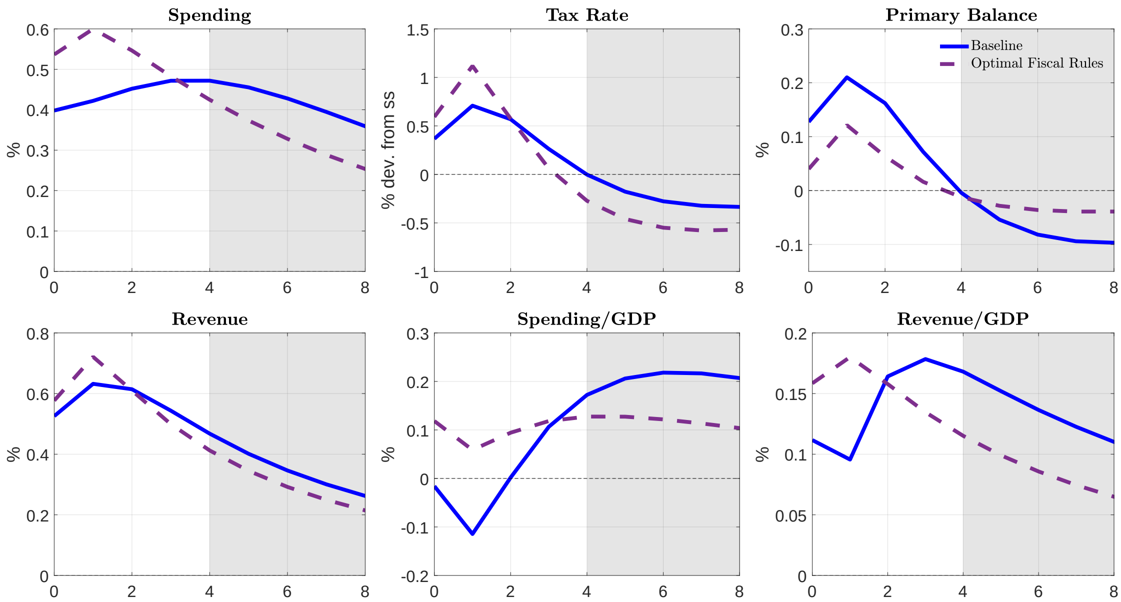
We report the unconditional welfare cost calculated in terms of steady state consumption, λ. Specifically, the unconditional welfare cost associated with the alternative policy regime, i.e. the fiscal rule, instead of the optimal Ramsey policy, is defined as

$$\mathbb{E}\Omega_0^A = \mathbb{E} \sum_{t=0}^{\infty} \beta^t U\left[(1-\lambda)\, c_t^R, c_{t-1}^R, g_t^R, g_{t-1}^R, h_t^{x,R}, h_t^{n,R}\right]. \tag{H.3}$$

We approximate λ up to the second order of accuracy (for further details see Schmitt-Grohé and Uribe, 2007). Note that we only report the unconditional welfare cost λ in the main text.

[37] Moreover, we impose some bounds for the search algorithm. Specifically, the GDP and export price policy coefficients are restricted to (-3, 3) and debt coefficients in (0,3). These restrictions are in line with Schmitt-Grohé and Uribe (2007).

Figure H.1: Ramsey Policy versus Policy Rules



Notes: This Figure shows the impulse responses of fiscal variables to (a one standard deviation) export price shock in the baseline model (in solid blue) and in a model with optimal fiscal rules (in dashed purple). Responses of the tax rate are expressed in percentage deviations from steady state values. Responses of spending, revenue, primary balance and fiscal ratios are in percentage points. Horizontal axes denote years. Gray areas denote the periods outside of the empirical horizon.

H.2 Impulse Responses under Optimal Policy Rules

In Figure H.1, we report the impulse responses of the fiscal variables to a shift in export prices when fiscal policy is conducted following the optimal fiscal rules and compare those to the one we retrieve conditional on the Ramsey policy.

The policy rules produce qualitatively similar profiles in terms of the overall fiscal stance, with some minor differences in terms of "strength". Specifically, when fiscal policy follows the optimal policy rules, spending becomes more front-loaded relative to the Ramsey policy. Higher government spending requires a stronger increase in the tax rate, which in turn increases the wedge between the marginal rate of substitution and the marginal product of labor. This leads to a somewhat weakened consumption, and therefore a welfare loss.

As a result of the weaker consumption response under the optimal fiscal rules, the base effect is slightly smaller on impact and increasing revenues requires a more pronounced hike in the tax rate. As fiscal revenues rise more than spending, the introduction of rules induces a positive (but more muted) response in the primary balance, leading to a lower response of public savings.

Despite the fact that the optimal fiscal rules imply that the primary balance over GDP features a negative coefficient associated with the price of exports, the overall response is dominated by the response to GDP and requires that the government saves part of the windfall under the optimal rules. The spending-to-GDP ratio and revenue-to-GDP ratio respond more strongly than in the baseline in the short run, but the effects are more transitory. In fact, under the optimal rules, the spending-to-GDP ratio does not fall in the short-term and remains consistently positive over the entire horizon. Consequently, while the resulting tax rate becomes more countercyclical, government spending turns out to be more procyclical than under the Ramsey policy.

I Modeling Institutions

In this Appendix, we expand the theoretical model in two directions. First, we propose that economies with weaker institutions experience inefficiencies associated with the production of final output, possibly due to rent-seeking behavior. Second, we assume that countries with lower institutional quality encounter more difficulties in accessing international financial markets. Therefore, when they do, they face more stringent borrowing conditions.

We model production inefficiencies and the possibility of rent-seeking behavior in the economy by assuming that countries with lower institutional quality operate under a decreasing returns technology across production sectors. Specifically, we introduce a parameter that affects the production scale in both the non-tradable and exportable sectors. Equation (E.15) is modified as follows:

$$y_t^j = \left[a^j \left(h_t^j\right)^{\alpha_j} \left(m_t^j\right)^{1-\alpha_j}\right]^{\xi} \quad \text{for} \quad j = \{n, x\}, \tag{I.1}$$

where $0 < \xi \leq 1$ represents the degree of inefficiency attributable to low institutional quality, with $\xi < 1$ denoting low institutional quality. The value of ξ directly modulates the shock's overall effect. This parameter reflects the concept that institutional capital is used in the production of final output, conceptualized as an aggregate rather than being sector-specific.

A decreasing returns technology generates positive profits in equilibrium,

$$\pi_t^j = (1 - \xi)\, p_t^j y_t^j, \tag{I.2}$$

which can also be thought of as the accrual of rents in the presence of economic inefficiencies. These inefficiencies work to alter the responsiveness of equilibrium prices vis-à-vis equilibrium quantities. In particular, a decreasing returns technology changes the slope of the sectoral supply curves without altering the demand schedules.

In terms of the second extension, we argue that countries with low institutional quality are confronted with higher costs of financing by both the private and public sectors. This reflects the probable reluctance of international investors to lend funds to countries with low institutional quality, who must pay a higher premium when it comes to borrowing internationally. To model this channel, we simply increase the value of the debt elastic premium for both private and public debts (relative to the baseline).[38] Moreover, we assume that the spread is insensitive to export prices, therefore spreads of countries with lower institutional quality do

[38]Specifically, we set the value of ξ to 0.65 and increase the debt elastic premia (both ψ and ψ_g) to 5.

not fall with terms-of-trade improvements. The last extension implies setting $\rho_x = \rho_{x,g} = 0$. The latter is motivated by the muted reaction of the EMBI in countries characterized by low-quality of institutions (reported in Figure 5) and is also in line with Arezki and Brückner (2012).

In the remainder of the Section, we provide a detailed mathematical formulation of our model featuring production inefficiency in the form of decreasing return to scale. We highlight in red the difference with the baseline model. If $\xi = 1$ the extension collapses to the baseline model. We also assume that countries with lower quality of institutions face a more limited ability to access international financial markets, captured by assuming a larger elasticity of the spread to debt. This incorporates the idea that countries with lower quality of institutions face a more reactive response of financial markets to debt.

Note that in equilibrium, $\bar{d}_t = d_t$, $\bar{c}_t = c_t$ and $\bar{g}_t = g_t$. This is not internalized by households but internalized by the Ramsey planner. In order to reduce the dimensionality of the problem, we replace the first order conditions with respect to m_t^x and m_t^n into the production functions. We also remove two equations by eliminating w_t^x and w_t^n. We substitute for profits in the household's budget constraint and replace for m_t^x and m_t^n using the firms' first order conditions. Therefore, we get the following five constraints:

$$
\begin{aligned}
y_t^n &= \left[a_t^n \left(h_t^n\right)^{\alpha_n \xi} \left(\left(1-\alpha_n\right) \xi \frac{p_t^n}{q_t p_t^{m,\$}} \right)^{(1-\alpha_n)\xi} \right]^{\frac{1}{1-(1-\alpha_n)\xi}}, \\
y_t^x &= \left[a_t^x \left(h_t^x\right)^{\alpha_x \xi} \left(\left(1-\alpha_x\right) \xi \frac{p_t^{x,\$}}{p_t^{m,\$}} \right)^{(1-\alpha_x)\xi} \right]^{\frac{1}{1-(1-\alpha_x)\xi}}, \\
\left(h_t^n\right)^{1+\varphi} &= \lambda_t p_t^n \alpha_n \xi y_t^n, \\
\left(h_t^x\right)^{1+\varphi} &= \lambda_t q_t p_t^{x,\$} \alpha_x \xi y_t^x, \\
q_t \left(1+r_t\right) d_{t-1} - q_t d_t + q_t \left(1+r_t^g\right) d_{t-1}^g - q_t d_t^g &= \left[1-\left(1-\alpha_x\right)\xi\right] q_t p_t^{x,\$} y_t^x - \\
&\quad - \left(1-\alpha_n\right) \xi p_t^n y_t^n - q_t p_t^{x,\$} \left(c_t^x + g_t^x\right) - q_t p_t^{m,\$} \left(c_t^m + g_t^m\right).
\end{aligned}
$$

The Ramsey planner chooses λ_t, c_t, c_t^T, c_t^m, c_t^x, h_t^n, h_t^x, y_t^n, y_t^x, p_t^n, p_t^T, q_t, d_t^g, d_t, g_t, g_t^T, g_t^m, g_t^x and θ_t so as to maximize:

$$
\begin{aligned}
\mathcal{L} =& \mathbb{E}_0 \sum_{t=0}^{\infty} \beta^t \left[\phi_t \frac{\left(c_t - b c_{t-1}\right)^{1-\sigma_c} - 1}{1-\sigma_c} + \left(1-\phi_t\right) \frac{\left(g_t - b g_{t-1}\right)^{1-\sigma_g} - 1}{1-\sigma_g} - \sum_{j=\{n,x\}} \frac{\left(h_t^j\right)^{1+\varphi}}{1+\varphi} \right] + \\
& \mathbb{E}_0 \sum_{t=0}^{\infty} \beta^t \mu_{1t} \left\{ q_t d_t - q_t \left(1+r_{t-1}\right) d_{t-1} + \right. \\
& + q_t d_t^g - q_t \left(1+r_{t-1}\right) d_{t-1}^g + \\
& \left. + \left[1-\left(1-\alpha_x\right)\xi\right] q_t p_t^{x,\$} y_t^x - \left(1-\alpha_n\right) \xi p_t^n y_t^n - q_t p_t^{x,\$} \left(c_t^x + g_t^x\right) - q_t p_t^{m,\$} \left(c_t^m + g_t^m\right) \right\} + \\
& E_0 \sum_{t=0}^{\infty} \beta^t \mu_{2t} \left[q_t d_t^g - q_t \left(1+r_{t-1}\right) d_{t-1}^g + \left(\theta_t - 1\right) c_t - g_t \right] +
\end{aligned}
$$

$$\mathbb{E}_0 \sum_{t=0}^{\infty} \beta^t \mu_{3t} \left[\phi_t \left(c_t - bc_{t-1}\right)^{-\sigma_c} - \theta_t \lambda_t\right] +$$

$$E_0 \sum_{t=0}^{\infty} \beta^t \mu_{4t} \left[q_t \lambda_t - \beta \left(1 + r_t\right) q_{t+1} \lambda_{t+1}\right] +$$

$$\mathbb{E}_0 \sum_{t=0}^{\infty} \beta^t \mu_{5t} \left[c_t^n - (1-\chi) \left(p_t^n\right)^{-\epsilon} c_t\right] + \mathbb{E}_0 \sum_{t=0}^{\infty} \beta^t \mu_{6t} \left[c_t^\tau - \chi \left(p_t^\tau\right)^{-\epsilon} c_t\right] +$$

$$\mathbb{E}_0 \sum_{t=0}^{\infty} \beta^t \mu_{7t} \left[c_t^m - \nu \left(\frac{q_t p_t^{m,\$}}{p_t^\tau}\right)^{-\eta} c_t^\tau\right] + \mathbb{E}_0 \sum_{t=0}^{\infty} \beta^t \mu_{8t} \left[c_t^x - (1-\nu) \left(\frac{q_t p_t^{x,\$}}{p_t^\tau}\right)^{-\eta} c_t^\tau\right] +$$

$$\mathbb{E}_0 \sum_{t=0}^{\infty} \beta^t \mu_{9t} \left\{ y_t^n - \left[a_t^n \left(h_t^n\right)^{\alpha_n \xi} \left((1-\alpha_n) \xi \frac{p_t^n}{q_t p_t^{m,\$}}\right)^{(1-\alpha_n)\xi}\right]^{\frac{1}{1-(1-\alpha_n)\xi}} \right\} +$$

$$\mathbb{E}_0 \sum_{t=0}^{\infty} \beta^t \mu_{10t} \left\{ y_t^x - \left[a_t^x \left(h_t^x\right)^{\alpha_x \xi} \left((1-\alpha_x) \xi \frac{p_t^{x,\$}}{p_t^{m,\$}}\right)^{(1-\alpha_x)\xi}\right]^{\frac{1}{1-(1-\alpha_x)\xi}} \right\} +$$

$$\mathbb{E}_0 \sum_{t=0}^{\infty} \beta^t \mu_{11t} \left[\left(h_t^n\right)^{1+\varphi} - \lambda_t \xi \alpha_n p_t^n y_t^n\right] + \mathbb{E}_0 \sum_{t=0}^{\infty} \beta^t \mu_{12t} \left[\left(h_t^x\right)^{1+\varphi} - \lambda_t \xi \alpha_x q_t p_t^{x,\$} y_t^x\right] +$$

$$\mathbb{E}_0 \sum_{t=0}^{\infty} \beta^t \mu_{13t} \left[\left(\chi \left(p_t^\tau\right)^{1-\epsilon} + (1-\chi) \left(p_t^n\right)^{1-\epsilon}\right)^{\frac{1}{1-\epsilon}} - 1\right] +$$

$$\mathbb{E}_0 \sum_{t=0}^{\infty} \beta^t \mu_{14t} \left[\left(\nu \left(p_t^{m,\$}\right)^{1-\eta} + (1-\nu) \left(p_t^{x,\$}\right)^{1-\eta}\right)^{\frac{1}{1-\eta}} q_t - p_t^\tau\right] +$$

$$\mathbb{E}_0 \sum_{t=0}^{\infty} \beta^t \mu_{15t} \left[g_t^n - (1-\chi) \left(p_t^n\right)^{-\epsilon} g_t\right] + \mathbb{E}_0 \sum_{t=0}^{\infty} \beta^t \mu_{16t} \left[g_t^\tau - \chi \left(p_t^\tau\right)^{-\epsilon} g_t\right] +$$

$$\mathbb{E}_0 \sum_{t=0}^{\infty} \beta^t \mu_{17t} \left[g_t^m - \nu \left(\frac{q_t p_t^{m,\$}}{p_t^\tau}\right)^{-\eta} g_t^\tau\right] + \mathbb{E}_0 \sum_{t=0}^{\infty} \beta^t \mu_{18t} \left[g_t^x - (1-\nu) \left(\frac{q_t p_t^{x,\$}}{p_t^\tau}\right)^{-\eta} g_t^\tau\right] +$$

$$\mathbb{E}_0 \sum_{t=0}^{\infty} \beta^t \mu_{19t} \left[y_t^n - c_t^n - g_t^n\right],$$

where

$$r_t = r^* + s + \psi \left[\exp\left(d_t - d\right) - 1\right] - \rho_x \ln\left(p_t^{x,\$}\right), \tag{I.3}$$

$$r_t^g = r^* + s + \psi_g \left[\exp\left(d_t^g - d^g\right) - 1\right] - \rho_{x,g} \ln\left(p_t^{x,\$}\right). \tag{I.4}$$

The first order conditions are:

$$\lambda_t : \mu_{4t} q_t = (1 + r_{t-1}) q_t \mu_{4t-1} + \mu_{11t} \xi \alpha_n p_t^n y_t^n + \mu_{12t} \xi \alpha_x q_t p_t^{x,\$} y_t^x + \mu_{3t} \theta_t, \tag{I.5}$$

$$c_t : \quad \mu_{5t} c_t^n + \mu_{6t} c_t^\tau + \sigma_c c_t \left[\mu_{3t} \phi_t \left(c_t - bc_{t-1}\right)^{-\sigma_c - 1} - \mathbb{E}_t \mu_{3t+1} \phi_{t+1} \left(c_{t+1} - bc_t\right)^{-\sigma_c - 1}\right] =$$

$$= c_t \left[\phi_t \left(c_t - bc_{t-1}\right)^{-\sigma_c} - \beta b \mathbb{E}_t \phi_{t+1} \left(c_{t+1} - bc_t\right)^{-\sigma_c}\right] + \mu_{2t} \left(\theta_t - 1\right) c_t, \tag{I.6}$$

$$c_t^n : \mu_{19t} = \mu_{5t}, \tag{I.7}$$

$$
\begin{aligned}
c_t^{\tau} &: \mu_{7t} c_t^m + \mu_{8t} c_t^x = \mu_{6t} c_t^{\tau}, \\
c_t^m &: \mu_{1t} q_t p_t^{m,\$} = \mu_{7t}, && \text{(I.8)} \\
c_t^x &: \mu_{1t} q_t p_t^{x,\$} = \mu_{8t}, && \text{(I.9)} \\
h_t^n &: \mu_{11t} \left(1+\varphi\right) \left(h_t^x\right)^{1+\varphi} = \left(h_t^n\right)^{1+\varphi} + \mu_{9t} \frac{\xi \alpha_n}{1-\left(1-\alpha_n\right)\xi} y_t^n, && \text{(I.10)} \\
h_t^x &: \mu_{12t} \left(1+\varphi\right) \left(h_t^x\right)^{1+\varphi} = \left(h_t^x\right)^{1+\varphi} + \frac{\xi \alpha_x}{1-\left(1-\alpha_x\right)\xi} \mu_{10t} y_t^x, && \text{(I.11)} \\
y_t^n &: \mu_{11t} \lambda_t \alpha_n \xi p_t^n + \mu_{1t}\left(1-\alpha_n\right) \xi p_t^n = \mu_{9t} + \mu_{19t}, && \text{(I.12)} \\
y_t^x &: \mu_{12t} \lambda_t \alpha_x \xi q_t p_t^{x,\$} = \mu_{1t} \left[1-\left(1-\alpha_x\right)\xi\right] q_t p_t^{x,\$} + \mu_{10t}, && \text{(I.13)} \\
p_t^n &: \epsilon \left(\mu_{5t} c_t^n + \mu_{15t} g_t^n\right) + \mu_{13t}(1-\chi)\left(p_t^n\right)^{1-\epsilon} = \mu_{1t}\left(1-\alpha_n\right)\xi p_t^n y_t^n + \\
&\quad + \mu_{11t} \alpha_n \xi p_t^n y_t^n + \mu_{9t} \frac{\left(1-\alpha_n\right)\xi}{1-\left(1-\alpha_n\right)\xi} y_t^n, && \text{(I.14)} \\
p_t^{\tau} &: \epsilon \left[\mu_{6t} c_t^{\tau} + \mu_{16t} g_t^{\tau}\right] + \mu_{13t} \chi \left(p_t^{\tau}\right)^{1-\epsilon} = \eta \left[\mu_{7t} c_t^m + \mu_{8t} c_t^x + \mu_{17t} g_t^m + \mu_{18t} g_t^x\right] + \\
&\quad + \mu_{14t} p_t^{\tau}, && \text{(I.15)} \\
q_t &: \mu_{12t} \lambda_t \xi \alpha_x q_t p_t^{x,\$} y_t^x + \mu_{4t-1} q_t \left(1+r_{t-1}\right) \lambda_{t-1} = q_t \mu_{4t} \lambda_t + \\
&\quad + \eta \left(\mu_{7t} c_t^m + \mu_{8t} c_t^x + \mu_{17t} g_t^m + \mu_{18t} g_t^x\right) + \\
&\quad + q_t \mu_{1t} \Big[d_t - \left(1+r_{t-1}\right) d_{t-1} + d_t^g - \left(1+r_{t-1}^g\right) d_{t-1}^g + \\
&\quad + \left[1-\left(1-\alpha_x\right)\xi\right] p_t^{x,\$} y_t^x - p_t^{x,\$}\left(c_t^x + g_t^x\right) - p_t^{m,\$}\left(c_t^m + g_t^m\right)\Big] + \\
&\quad + q_t \mu_{2t} \left[d_t^g - \left(1+r_{t-1}\right) d_{t-1}^g\right] + \mu_{14t} p_t^{\tau} + \mu_{9t} \frac{\left(1-\alpha_n\right)\xi}{1-\left(1-\alpha_n\right)\xi} y_t^n, && \text{(I.16)} \\
d_t^g &: \left(\mu_{1t} + \mu_{2t}\right) q_t = \beta \mathbb{E}_t \left(\mu_{1t+1} + \mu_{2t+1}\right) q_{t+1} \left[1 + r_t^g + \psi_g \left(e^{d_t^g - d^g}\right) d_t^g\right], && \text{(I.17)} \\
g_t &: \quad \left[\left(1-\phi_t\right)\left(g_t - b g_{t-1}\right)^{-\sigma_g} - \beta b \mathbb{E}_t \left(1-\phi_{t+1}\right)\left(g_{t+1} - b g_t\right)^{-\sigma_g}\right] g_t = \mu_{2t} g_t + \\
&\quad + \mu_{15t} g_t^n + \mu_{16t} g_t^{\tau}, && \text{(I.18)} \\
g_t^n &: \mu_{15t} = \mu_{19t}, && \text{(I.19)} \\
g_t^{\tau} &: \mu_{16t} g_t^{\tau} = \mu_{17t} g_t^m + \mu_{18t} g_t^x, && \text{(I.20)} \\
g_t^m &: q_t p_t^{m,\$} \mu_{1t} = \mu_{17t}, && \text{(I.21)} \\
g_t^x &: q_t p_t^{x,\$} \mu_{1t} = \mu_{18t}, && \text{(I.22)} \\
\theta_t &: \mu_{2t} c_t = \mu_{3t} \lambda_t. && \text{(I.23)}
\end{aligned}
$$

And 19 Lagrange multipliers with associated constraints.

I.1 Steady State

The steady state consists of the following set of equations:

$$1 = \beta\left(1+r^*\right), \tag{I.24}$$

$$c^n = (1-\chi)\left(p^n\right)^{-\epsilon} c, \tag{I.25}$$

$$c^{\tau} = \chi\left(p^{\tau}\right)^{-\epsilon} c, \tag{I.26}$$

$$c^m = \nu c^\tau, \tag{I.27}$$
$$c^x = (1-\nu)\, c^\tau, \tag{I.28}$$
$$g^n = (1-\chi)\,(p^n)^{-\epsilon}\, g, \tag{I.29}$$
$$g^\tau = \chi\,(p^\tau)^{-\epsilon}\, g, \tag{I.30}$$
$$g^m = \nu g^\tau, \tag{I.31}$$
$$g^x = (1-\nu)\, g^\tau, \tag{I.32}$$
$$\lambda = \phi \frac{[(1-b)\, c]^{-\sigma_c}}{\theta}, \tag{I.33}$$
$$q = p^\tau, \tag{I.34}$$
$$\mu_1 p^\tau = \mu_6, \tag{I.35}$$
$$\mu_1 q = \mu_8, \tag{I.36}$$
$$\mu_1 q = \mu_7, \tag{I.37}$$
$$\mu_{19} = \mu_{15}, \tag{I.38}$$
$$\mu_{15} = \mu_5, \tag{I.39}$$
$$\mu_{17} = p^\tau \mu_1, \tag{I.40}$$
$$\mu_{18} = p^\tau \mu_1, \tag{I.41}$$
$$\mu_{16} = p^\tau \mu_1, \tag{I.42}$$

the conditions related to implementability:

$$\begin{aligned} r\frac{d}{y}\left\{[1-(1-\alpha_n)\,\xi]\, p^n y^n + [1-(1-\alpha_x)\,\xi]\, p^\tau y^x\right\} = \\ = [1-(1-\alpha_x)\,\xi]\, p^\tau y^x - (1-\alpha_n)\,\xi p^n y^n - \chi\,(p^\tau)^{1-\epsilon}\,(c+g)\,, \end{aligned} \tag{I.43}$$
$$(\theta - 1)\, c = g, \tag{I.44}$$
$$y^n = \left[(h^n)^{\alpha_n \xi}\left((1-\alpha_n)\,\xi\frac{p^n}{p^\tau}\right)^{(1-\alpha_n)\xi}\right]^{\frac{1}{1-(1-\alpha_n)\xi}}, \tag{I.45}$$
$$y^x = \left[(h^x)^{\alpha_x \xi}\left((1-\alpha_x)\,\xi\right)^{(1-\alpha_x)\xi}\right]^{\frac{1}{1-(1-\alpha_x)\xi}}, \tag{I.46}$$
$$\theta\,(h^n)^{1+\varphi} = [(1-b)\, c]^{-\sigma_c}\,\alpha_n p^n y^n, \tag{I.47}$$
$$\theta\,(h^x)^{1+\varphi} = [(1-b)\, c]^{-\sigma_c}\,\alpha_x p^\tau y^x, \tag{I.48}$$
$$1 = \chi\,(p^\tau)^{1-\epsilon} + (1-\chi)\,(p^n)^{1-\epsilon}, \tag{I.49}$$
$$y^n = (1-\chi)\,(p^n)^{-\epsilon}\,(c+g)\,, \tag{I.50}$$

and the conditions associated with Lagrange multipliers:

$$-rp^\tau \mu_4 = \mu_{11}\alpha_n \xi p^n y^n + \mu_{12}\alpha_x \xi p^\tau y^x + \mu_3 \theta, \tag{I.51}$$
$$\begin{aligned} \phi\,[(1-b)\, c]^{-\sigma_c}\,(1-b\beta) + \mu_2\,(\theta-1) = \mu_5\,(1-\chi)\,(p^n)^{-\epsilon} + \mu_6 \chi\,(p^\tau)^{-\epsilon} + \\ + \mu_{13} + \mu_3 \phi \sigma_c\,[(1-b)\, c]^{-\sigma_c - 1}\,(1-b\beta)\,, \end{aligned} \tag{I.52}$$

$$\mu_{11}\left(1+\varphi\right)\left(h^{x}\right)^{1+\varphi}=\left(h^{n}\right)^{1+\varphi}+\frac{\alpha_{n}\xi}{1-\left(1-\alpha_{n}\right)\xi}\mu_{9}y^{n}, \tag{I.53}$$

$$\mu_{12}\left(1+\varphi\right)\left(h^{x}\right)^{1+\varphi}=\left(h^{x}\right)^{1+\varphi}+\frac{\alpha_{x}\xi}{1-\left(1-\alpha_{x}\right)\xi}\mu_{10}y^{x}, \tag{I.54}$$

$$\theta\left(\mu_{9}+\mu_{5}\right)=\mu_{11}\left[\left(1-b\right)c\right]^{-\sigma_{c}}\alpha_{n}p^{n}+\mu_{1}\left(1-\alpha_{n}\right)\xi p^{n}, \tag{I.55}$$

$$\mu_{12}\left[\left(1-b\right)c\right]^{-\sigma_{c}}\alpha_{x}p^{\tau}=\theta\left\{\left[1-\left(1-\alpha_{n}\right)\xi\right]\mu_{1}p^{\tau}+\mu_{10}\right\}, \tag{I.56}$$

$$\begin{aligned}\epsilon\mu_{5}(1-\chi)\left(p^{n}\right)^{-\epsilon}\left(c+g\right)+\mu_{13}(1-\chi)\left(p^{n}\right)^{1-\epsilon}&=\mu_{1}\left(1-\alpha_{n}\right)\xi p^{n}y^{n}+\\&+\mu_{11}\xi p^{n}y^{n}+\frac{\alpha_{n}\xi}{1-\left(1-\alpha_{n}\right)\xi}\mu_{9}y^{n},\end{aligned} \tag{I.57}$$

$$\left(\epsilon-\eta\right)\mu_{1}\left(c+g\right)+\mu_{13}=\mu_{14}\frac{\left(p^{\tau}\right)^{\epsilon}}{\chi}, \tag{I.58}$$

$$\begin{aligned}\left(\mu_{12}\xi\alpha_{x}p^{\tau}y^{x}+r^{*}p^{\tau}\mu_{4}\right)\left[\left(1-b\right)c\right]^{-\sigma_{c}}&=\theta\eta\mu_{1}\chi\left(p^{\tau}\right)^{1-\epsilon}\left(c+g\right)+\\&+\theta\mu_{1}\left\{-r\frac{d}{y}\left\{\left[1-\left(1-\alpha_{n}\right)\xi\right]p^{n}y^{n}+\right.\right.\\&+\left[1-\left(1-\alpha_{x}\right)\xi\right]p^{\tau}y^{x}\right\}+\\&\left.+\left[1-\left(1-\alpha_{x}\right)\xi\right]p^{\tau}y^{x}-\chi\left(p^{\tau}\right)^{1-\epsilon}\left(c+g\right)\right\}+\\&+\theta\left\{\mu_{14}p^{\tau}+\mu_{9}\frac{\left(1-\alpha_{n}\right)\xi}{\left[1-\left(1-\alpha_{n}\right]\xi\right.}y^{n}\right\},\end{aligned} \tag{I.59}$$

$$\mu_{1}\frac{d}{y}\theta=-\frac{\mu_{4}p^{\tau}\left[\left(1-b\right)c\right]^{-\sigma_{c}}}{\left\{\left[1-\left(1-\alpha_{n}\right)\xi\right]p^{n}y^{n}+\left[1-\left(1-\alpha_{x}\right)\xi\right]p^{\tau}y^{x}\right\}}, \tag{I.60}$$

$$\left(1-\phi\right)\left[\left(1-b\right)g\right]^{-\sigma_{g}}\left(1-b\beta\right)=\mu_{2}+\mu_{15}\left(1-\chi\right)\left(p^{n}\right)^{-\epsilon}+\mu_{1}\chi\left(p^{\tau}\right)^{1-\epsilon}, \tag{I.61}$$

$$\mu_{2}\theta c=\mu_{3}\phi\left[\left(1-b\right)c\right]^{-\sigma_{c}}, \tag{I.62}$$

$$\frac{g}{y}=\frac{g}{\left\{\left[1-\left(1-\alpha_{n}\right)\xi\right]p^{n}y^{n}+\left[1-\left(1-\alpha_{x}\right)\xi\right]p^{\tau}y^{x}\right\}}. \tag{I.63}$$

Appendix References

ABEL, A. B. (1990): "Asset Prices under Habit Formation and Catching Up with the Joneses," *American Economic Review*, 80, 38–42.

AGUIAR, M. AND G. GOPINATH (2007): "Emerging Market Business Cycles: The Cycle Is the Trend," *Journal of Political Economy*, 115, 69–102.

AREZKI, R. AND M. BRÜCKNER (2012): "Commodity Windfalls, Democracy and External Debt," *Economic Journal*, 122, 848–866.

ARSENEAU, D. M. AND S. LEDUC (2013): "Commodity Price Movements in a General Equilibrium Model of Storage," *IMF Economic Review*, 61, 199–224.

BARRO, R. J. (1981): "Output Effects of Government Purchases," *Journal of Political Economy*, 89, 1086–1121.

CHRISTIANO, L. J. AND M. EICHENBAUM (1992): "Current Real-Business-Cycle Theories and Aggregate Labor-Market Fluctuations," *The American Economic Review*, 82, 430–450.

CONSTANTINIDES, G. M. (1990): "Habit Formation: A Resolution of the Equity Premium Puzzle," *Journal of Political Economy*, 98, 519–543.

DE WINNE, J. AND G. PEERSMAN (2016): "Macroeconomic Effects of Disruptions in Global Food Commodity Markets: Evidence for the United States," *Brookings Papers on Economic Activity*, 47, 183–286.

DI PACE, F., L. JUVENAL, AND I. PETRELLA (2024): "Terms-of-Trade Shocks are Not all Alike," *American Economic Journal: Macroeconomics*, Forthcoming.

DRECHSEL, T. AND S. TENREYRO (2018): "Commodity Booms and Busts in Emerging Economies," *Journal of International Economics*, 112, 200–218.

FERNÁNDEZ, A., D. GUZMAN, R. E. LAMA, AND C. A. VÉGH (2021): "Procyclical Fiscal Policy and Asset Market Incompleteness," NBER Working Papers 29149, National Bureau of Economic Research, Inc.

HAMANN, F., J. C. MENDEZ-VIZCAINO, E. G. MENDOZA, AND P. RESTREPO-ECHAVARRIA (2023): "Natural Resources and Sovereign Risk in Emerging Economies: A Curse and a Blessing," Working Paper 31058, National Bureau of Economic Research.

JANZEN, J. P., A. SMITH, AND C. A. CARTER (2018): "Commodity Price Comovement and Financial Speculation: The Case of Cotton," *American Journal of Agricultural Economics*, 100, 264–285.

KÄNZIG, D. R. (2021): "The Macroeconomic Effects of Oil Supply News: Evidence from OPEC Announcements," *American Economic Review*, 111, 1092–1125.

LEITH, C., I. MOLDOVAN, AND R. ROSSI (2015): "Monetary and Fiscal Policy Under Deep Habits," *Journal of Economic Dynamics and Control*, 52, 55–74.

MACDONALD, S. AND L. MEYER (2018): "Long Run Trends and Fluctuations In Cotton Prices," MPRA Paper 84484, University Library of Munich, Germany.

MENDES, A. AND S. M. PENNINGS (2020): "One Rule Fits All Heterogeneous Fiscal Rules for Commodity Exporters When Price Shocks Can Be Persistent: Theory and Evidence," Policy Research working paper 9400, World Bank Group.

MENDOZA, E. G. (1995): "The Terms of Trade, the Real Exchange Rate, and Economic Fluctuations," *International Economic Review*, 36, 101–137.

RIASCOS, A. AND C. VÉGH (2003): "Procyclical Government Spending in Developing Countries: The Role of Capital Market Imperfections," Tech. rep., Working paper, University of Maryland.

SCHMITT-GROHÉ, S. AND M. URIBE (2003): "Closing Small Open Economy Models," *Journal of International Economics*, 61, 163–185.

——— (2007): "Optimal Simple and Implementable Monetary and Fiscal Rules," *Journal of Monetary Economics*, 54, 1702–1725.

——— (2018): "How Important are Terms-of-Trade Shocks?" *International Economic Review*, 59, 85–111.

SOHNGEN, B. AND R. HAYNE (1994): "The 'Great' Price Spike of '93: An Analysis of Lumber and Stumpage Prices in the Pacific Northwest," Forest Service Research Paper PNW-RP-476, United States Department of Agriculture.

URIBE, M. AND V. Z. YUE (2006): "Country Spreads and Emerging Countries: Who Drives Whom?" *Journal of International Economics*, 69, 6–36.

WORLD BANK, G. (2001): "Global Economic Prospects," Tech. rep., Washington, D.C.