Fiscal Imbalances and Asset Returns: Cross-Sector Fluctuations under the Aggregate Budget Constraint

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Abstract

We formulate the aggregate budget constraint as nesting the private, public, and external sector constraints. This structure implies cross-sector restrictions linking valuation ratios to future real returns and cash-flow growth. Using these restrictions, we show that the government surplus-to-debt ratio negatively predicts equity returns. The strength of this cross-sector predictability is robust and comparable to the own-sector predictability of the dividend–price ratio. We propose a model with distortionary taxation that accounts for these dynamics and use the cross-sector forecasts to quantify the implied tax distortions.

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Keywords: Equity predictability, public debt, fiscal imbalances, cross-sector predictability, output distortion

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1 Introduction

Present-value relations have become the empirical workhorse for extracting time-varying expectations of economic fundamentals from fluctuations in valuation ratios. A canonical example is the log-linearized present-value relation in Campbell and Shiller (1988b), a cornerstone of the equity return predictability literature,[1] which links fluctuations in the dividend–price ratio (hereafter, dp) to changing expectations of future equity returns and dividend growth. Similar present-value decompositions have also been applied to the public and foreign sectors of the U.S. economy. In the public sector, a high net surplus (taxes minus government spending) relative to government debt (nsb) predicts higher future primary surpluses and, to a lesser extent, lower future Treasury returns (Berndt et al., 2012; Campbell et al., 2023). In the foreign sector, current account imbalances, measured by net imports relative to the net value of foreign assets (nma), lead to higher future foreign portfolio returns (Gourinchas and Rey, 2007b).[2]

A common thread across this literature is that present-value relations are analyzed one sector at a time, with each sector treated in isolation. Consequently, fluctuations in a sector-specific valuation ratio are linked to that sector's own expected fundamentals, but not to time variation in fundamentals elsewhere in the economy. This sector-specific approach cannot address cross-sector questions such as whether fluctuations in nsb, reflecting changing fiscal conditions, lead to changes in expected equity returns or dividend growth, or conversely whether movements in dp, a state variable capturing conditions in the equity market, forecast Treasury returns or future net surpluses. More broadly, it remains unclear to what extent – and through which mechanisms – shocks originating in one sector are transmitted to others. Understanding these linkages is central to how fundamental economic shocks propagate across sectors.

Our paper offers three contributions toward addressing these issues. First, we develop a general framework that links the present-value relations of the private, public, and foreign sectors, allowing us to examine whether fluctuations in one sector lead time variations in expected returns or fundamentals in other sectors. The central identity linking the three sectors is the aggregate budget constraint (ABC), an accounting relation that we use to integrate the present-value conditions of the three sectors into a single intertemporal budget constraint. The ABC implies that the aggregate consumption-to-wealth ratio is a weighted average of three sector-specific valuation ratios: dp, nsb, and nma. Similarly, the ABC yields an expression for the return to total wealth as a weighted average of expected returns across the three sectors. Hence, fluctuations in any component of the consumption-to-wealth ratio can generate variation in expected returns or fundamental growth in its own sector or in either of the other two sectors.

Embedding the three present-value relations within the ABC makes clear that an imbalance

[1] A partial list includes Campbell (1993), Vuolteenaho (2002), Campbell and Vuolteenaho (2004), Larrain and Yogo (2008), Pástor and Stambaugh (2012), Menzly et al. (2004), Lettau and Ludvigson (2005), Lettau and Van Nieuwerburgh (2008), and Cochrane (2011).

[2] Present-value relations have also been applied to other asset classes. For corporate bonds, Nozawa (2017) shows that cash-flow variation is primarily driven by time-varying expected returns rather than default losses. In commercial real estate, Plazzi et al. (2010) demonstrate that rent-price ratios forecast future returns, not future rent growth.

in one sector must be absorbed either within that sector or elsewhere in the economy. While this general framework opens the possibility of cross-sector predictability, it does not by itself guarantee that such predictability will be present in the data. Two polar cases illustrate this point. If the expected returns on equities, Treasuries, and foreign assets are all driven by a single common shock, then each sectoral present-value relation collapses to the aggregate consumption-wealth relation. In this case, the other two sectoral constraints become essentially redundant, leaving no scope for cross-sector predictability. At the other extreme, if each sector's expected returns are driven by orthogonal shocks, then valuation ratios contain no cross-sector information. While those are admittedly unrealistic scenarios, they illustrate the point that only when fluctuations in valuation ratios propagate through the ABC, affecting consumption and discount rates, do cross-sector linkages emerge. As the consumption-wealth ratio aggregates all these forces, it cannot by itself disentangle their sources. A cross-sector approach that controls for each sector's valuation ratios is therefore essential for testing whether cross-sector linkages exist and for quantifying their magnitude.

Our second contribution is to document cross-sector predictability in U.S. data from 1952 to 2021. Using a unified present-value framework that accommodates negative surpluses, cash flows, and asset values across the private, public, and foreign sectors, we show that fiscal imbalances measured by the net surplus-to-debt ratio, *nsb*, predict future equity returns at quarterly and annual horizons. In unrestricted regressions, the marginal effect of *nsb* is negative and statistically significant, implying that higher surpluses relative to outstanding debt forecast lower subsequent equity returns. Including *nsb* increases explanatory power and reinforces, rather than weakens, the role of the dividend–price ratio. This result is robust to standard controls, including the tax-to-GDP ratio (Campbell et al., 2023), the Cochrane and Piazzesi (2005) forward-rate factor, and inflation as in Berndt et al. (2012). Among all cross-sector effects, *nsb* is the only predictor that is both economically meaningful and statistically robust.

We sharpen inference by exploiting the present-value restrictions implied by the sectoral budget constraints and the aggregate budget constraint. Casting the estimation in a constrained GMM framework, we treat these restrictions as linear constraints on predictive coefficients, which improves efficiency when the restrictions hold and mitigates bias when OLS estimates are distorted by persistent predictors. Under this approach, the predictive power of *nsb* strengthens further and remains significant at the 1% level across specifications. The framework also delivers a simple test of the restrictions, which we cannot reject in the data. Overall, the constrained approach yields sharper inference on cross-asset predictability, consistent with the role of cross-equation restrictions emphasized by Cochrane (2008) and related methodological work (Timmermann, 1993; Ang et al., 2007; Campbell and Thompson, 2008; Koijen and Van Nieuwerburgh, 2011). Our results are robust to econometric concerns with persistent predictors that include bias corrections, bootstrapped inference, and structural breaks. Simulations under a benchmark where *dp* fully explains expected returns show that the estimated *nsb* coefficient lies far outside the simulated range, indicating its predictive power is unlikely to be spurious.

Third, we examine why the net surplus-to-debt ratio forecasts expected equity returns within an asset-pricing framework. The key feature of the model is that fiscal policy shocks influence consumption growth by generating output distortions through future taxation or by stimulating activity through government spending, building on the frameworks of Barro (1979) and Jiang et al. (2024a). Under the aggregate budget constraint, a fiscal shock alters cash flows through two channels. First, taxation reallocates resources from the private to the public sector (abstracting from the foreign sector). Second, fiscal policy affects consumption dynamics and, through them, the pricing of risk across all sectors. It is this second channel that drives our main result: forward-looking fiscal adjustments that lower the surplus-to-debt ratio raise the equity risk premium and therefore predict higher future equity returns.

This mechanism also clarifies why cross-sector predictors are essential for identification. Fiscal shocks alter the common consumption-based risk component, so variation in the dividend-to-price ratio reflects not only within-sector dynamics but also fiscal spillovers transmitted through consumption. Consequently, predictive regressions that rely solely on within-sector valuation ratios confound equity-specific risk-premium variation with fiscal influences. Incorporating the surplus-to-debt ratio—anchored in the government's intertemporal budget constraint—helps isolate the fiscal component embedded in expected equity returns. In other words, the dividend-to-price ratio alone is insufficient to capture all channels of time-varying risk premia in equity returns. Consistent with the model, the surplus-to-debt ratio negatively predicts future equity returns even after controlling for the dividend-to-price ratio.

Our model builds on the government's optimal fiscal policy problem. The economy produces an exogenous endowment of raw output that accrues to the government. The government chooses tax and spending paths to maximize the representative household's lifetime utility, subject to the intertemporal public budget constraint and the value of outstanding debt. Taxation finances current spending and debt service but reduces the final output available to households by introducing distortions, while spending stimulates output and demand at the cost of higher debt. Households take these fiscal choices as given, hold equity and government debt, and price all claims using a consumption-based discount factor. Fiscal policy thus affects asset values through both cash-flow and discount-rate channels. When output distortions are absent, Ricardian Equivalence holds: consumption–wealth dynamics depend solely on the initial debt valuation, fiscal policy becomes neutral, and cross-sector predictability disappears.

Because government policy and asset prices are jointly determined in equilibrium, the model offers a tractable way to trace how fiscal expectations transmit through the economy. We study a marginal, anticipated deviation from the steady-state fiscal path—a small forward-looking fiscal shock—that captures how expectations of future tax or spending changes influence current asset prices through discount-rate effects. Such expectations need not correspond to observed policy events; valuation ratios often embed fiscal beliefs even in the absence of explicit actions, consistent with the empirical difficulty of identifying fiscal shocks. Anticipated fiscal tightening raises the public debt valuation, reduces the surplus-to-debt ratio and increases expected discount rates for

the private sector, giving rise to the cross-sector predictability documented in the data.

The model's tractability further allows us to link the estimated slope coefficients from predictive regressions directly to structural parameters of output distortions. Our calibration indicates that taxation reduces output by roughly 6%, while government spending raises it by about 2%. The negative predictability of equity returns thus stems from taxation's distortionary effect dominating the stimulative impact of spending, raising risk premia for the private sector.

Our approach of decomposing the economy into three sectors and imposing the aggregate budget constraint connects different strands of research on dynamic value adjustment. For equity, the seminal contribution of Campbell and Shiller (1988b), extended by Campbell (1991) to separate cash-flow and discount-rate shocks, underpins much of the work on return and dividend-growth predictability. This body of work highlights the limitations of the dividend-to-price ratio as a stand-alone predictor. We show that fiscal policy, captured by the net surplus-to-debt ratio, contains additional information beyond dp, and rationalize this finding within a theoretical framework.

Lettau and Ludvigson (2001) show that the aggregate consumption-to-wealth ratio contains information about future equity returns. Their analysis focuses on the private sector and distinguishes between labor and non-labor income in order to construct a state variable capturing expected return variation. We build on the same consumption-wealth dynamics but take a different perspective: exploiting the national accounts identity, we decompose consumption and wealth across the private, public, and foreign sectors, and ask whether returns and cash-flow growth in one sector can be predicted by valuation ratios from others. This cross-sector decomposition, complementary to the labor- versus non-labor split in Lettau and Ludvigson (2001), enables us to investigate cross-sector predictability and uncover new interactions between sectoral budget constraints and asset pricing.[3]

Recent work exploits the log debt-to-GDP identity to predict equity and Treasury returns (e.g., Cochrane, 2022; Liu, 2023; Jiang et al., 2024c). We extend this literature by showing that, through the aggregate budget constraint, the consumption-to-wealth ratio can be written as a weighted average of sector-specific valuation ratios, implying that the net surplus-to-debt ratio (nsb) is the key state variable governing the transmission of fiscal imbalances across sectors via consumption growth and discount rates. This central role reflects the fact that government debt is a claim on future fiscal surpluses, so fluctuations in nsb capture adjustments in the market value of public debt driven by anticipated surplus dynamics and provide a natural measure of the fiscal component of risk premia. While earlier studies using the government budget constraint focus on fiscal adjustment and inflation dynamics and therefore measure imbalances as linear combinations of tax-to-debt and spending-to-debt ratios under cointegration assumptions (Giannitsarou and Scott, 2006; Berndt et al., 2012; Berndt and Yeltekin, 2015), we build on Campbell et al. (2023) and directly log-linearize the surplus-to-debt ratio. This approach accommodates negative and highly volatile net surpluses and elevated government spending toward the end of the sample.

Our model builds on the optimal fiscal policy literature in which governments manage aggregate

[3]Conceptually, our framework could be extended by decomposing each sector's assets into human capital and investable components, in the spirit of Lettau and Ludvigson (2001). Since labor income is not directly observable, this would require constructing sector-specific analogues of "cay."

risk through tax smoothing (Barro, 1979; Lucas and Stokey, 1983; Bohn, 1990; Jiang et al., 2024a). A central question in that literature is the measurement of spending and tax multipliers using either DSGE models or narrative methods based on historical data (Blanchard and Perotti, 2002; Romer and Romer, 2010; Christiano et al., 2011; Auerbach and Gorodnichenko, 2012; Ilzetzki et al., 2013; Ramey and Zubairy, 2018). By explicitly incorporating equity pricing and exploiting cross-sector predictability, our approach implies the magnitude of tax distortions and government spending directly from asset-pricing dynamics, providing a novel and complementary alternative method to the existing approaches.

Our framework allows for risky government debt, consistent with the view that the path of net surpluses provides information about the degree of protection afforded to taxpayers relative to bondholders (Jiang et al., 2026). The presence of a debt return risk premium is closely connected to the literature on public debt sustainability.[4] Rather than focusing on expected returns on public debt, the analysis emphasizes the link between surplus dynamics and expected equity returns. This focus on the aggregate budget constraint and the cross-sector transmission of fiscal shocks complements the macro-finance literature studying how taxes and policy uncertainty affect firm-level investment and, ultimately, long-run growth (Pástor and Stambaugh, 2012; Croce et al., 2012a; Croce et al., 2012b; Croce et al., 2019; Corhay et al., 2023).

Our work is also related to Jiang et al. (2024b) who document deviations from the no-arbitrage relation between the market value of U.S. government debt and its fundamental value implied by future surpluses. This valuation puzzle relates to how the government can finance public debt at a lower cost than what its surpluses suggest.[5] While these studies highlight the safe-asset supply and convenience-yield channels in government debt markets, our analysis deliberately focuses on the fundamental channel linking fiscal policy, surplus expectations, and asset prices. Rather than attempting to reconcile the debt valuation puzzle itself, we use the surplus-to-debt ratio to study how fiscal distortions transmit across sectors and shape expected returns.

In international finance, the "valuation channel" captures how capital gains and losses affect the external balance sheet. Lane and Milesi-Ferretti (2001) pioneered this line of inquiry, and Gourinchas and Rey (2007b) formalized the approach using a foreign-sector budget constraint to test dynamic value adjustment. Subsequent work documents the rising importance of valuation effects across countries (Gourinchas, 2008; Gourinchas and Rey, 2014). We extend this research by developing a decomposition for foreign markets that accommodates the possibility of negative flows (net exports) and negative positions (net foreign assets), allowing us to apply a unified framework across sectors and examine whether foreign imbalances spill over into domestic asset markets.

[4] See inter alia: Blanchard (1985); Hamilton and Flavin (1986); Bohn (1998); Hall and Sargent (2011); Blanchard (2019); Mehrotra and Sergeyev (2021); Ball and Mankiw (2023); Mian et al. (2025); Elenev et al. (2025); and Liu et al. (2025)

[5] See inter alia: Krishnamurthy and Vissing-Jorgensen (2012); Gorton and Ordonez (2014); Du et al. (2018); He et al. (2019); Krishnamurthy and Lustig (2019); Gorton and Ordonez (2022); Atkeson et al. (2025); and Chen et al. (2025)

2 Present-Value Decompositions and Sector Constraints

Our starting point is the aggregate accounting identity of GDP:

$$Y_t = C_t + I_t + G_t + X_t - M_t \tag{1}$$

where Y_t denotes GDP, C_t denotes aggregate consumption, I_t is aggregate investment, G_t is government spending, X_t is exports, and M_t is imports. We can rewrite this identity as:

$$C_t = D_t + NS_t + NM_t\,, \tag{2}$$

where $D_t = Y_t - I_t - T_t$ is aggregate after-tax dividend to the private sector, $NS_t = T_t - G_t$ is the net government tax surplus computed as taxes T_t minus spending, and $NM_t = M_t - X_t$ is net imports. Conceptually, aggregate consumption is driven by flows from the private, the public, and the external sectors. The definition of dividend equal to output minus investment and taxes has been previously used in the literature to capture the net payout to the firm owners (see, e.g., Larrain and Yogo, 2008). Most asset pricing studies assume that aggregate consumption equals dividends, or $C_t = D_t$ (Lucas, 1978; Campbell, 1996). This implicitly side-steps the role of fiscal policy and net imports on asset returns, which is the focus of our paper.

We can formulate an intertemporal budget constraint for each sector. For the private sector, this is:

$$A_{t+1} + D_{t+1} = (1 + R^A_{t+1})A_t\,, \tag{3}$$

where A_t is the value of the private sector's wealth in the economy, and R^A_{t+1} its return.

The public sector's budget constraint is:

$$B_{t+1} + NS_{t+1} = (1 + R^B_{t+1})B_t\,, \tag{4}$$

where B_t is the total value of government debt outstanding, and R^B_{t+1} its return.

Finally, for the external sector, the budget constraint is:

$$F_{t+1} + NM_{t+1} = (1 + R^X_{t+1})F_t\,, \tag{5}$$

where F_t is the net value of foreign assets, i.e., assets minus liabilities, and R^X_{t+1} is the return on the foreign assets.

These budget constraints have previously been studied in isolation in the literature – see, among others, Campbell and Shiller (1988a) for the equity sector; Berndt et al. (2012) for the public sector; and Gourinchas and Rey (2007b) for the external sector.

We can define the total domestic wealth of the economy W_t as

$$W_t = A_t + B_t + F_t\,, \tag{6}$$

and using total outflows, C_t, we can write the aggregate budget constraint as:

$$W_{t+1} + C_{t+1} = (1 + R^W_{t+1})W_t \,. \tag{7}$$

The gross return to the wealth portfolio, $1 + R^W_{t+1}$, equals the weighted average of the returns to each sector, or

$$1 + R^W_{t+1} = (1 + R^A_{t+1})\frac{A_t}{W_t} + (1 + R^B_{t+1})\frac{B_t}{W_t} + (1 + R^X_{t+1})\frac{F_t}{W_t} \,. \tag{8}$$

2.1 Budget Constraints (BC) and Within-Sector Predictability

We obtain an expression relating the valuation ratio in each market to future asset returns and cash-flow growth by log-linearizing the intertemporal budget constraint and iterating forward the resulting expression. This approach has been pioneered in the equity context by Campbell and Shiller (1988b), which obtain an approximated relation with respect to the log dividend-to-price ratio. Their dynamic accounting identity forms the basis for the analysis of equity predictability that investigates the relative importance of time-variation in discount rates and expected dividend growth. The same approach cannot, however, be straightforwardly applied to the other two sectors, as cash flows to the public sector (i.e., net surplus) and to the foreign sector (i.e., exports minus imports) can turn negative, thereby invalidating the logarithmic transformation.

To overcome this issue, we adopt the approach in Campbell et al. (2023) for the public sector, and further extend it to the other two sectors. We provide a compact description of the approach to highlight the main results of interest for our empirical analysis, and then present the general framework with full derivations in Online Appendix Section A.

Campbell et al. (2023) linearize the surplus-debt ratio by exploiting the cointegrated relation between the log tax-debt and the log spending-debt ratio. Their method separately handles inflows (spending) and outflows (taxes) of the public sector and is based on an expansion of the log of one plus the surplus-to-debt ratio. Formally, the government's valuation ratio (net surplus-to-debt), denoted nsb, is approximated as:

$$nsb_{t+1} = \log\left(1 + \frac{T_{t+1} - G_{t+1}}{B_{t+1}}\right) \approx k + (1 - \rho^B)\left(\frac{1}{1-\beta}\tau b_{t+1} - \frac{\beta}{1-\beta}gb_{t+1}\right), \tag{9}$$

where τb and gb represent the log tax-to-debt and spending-to-debt ratios, respectively. The coefficient ρ^B is related to the steady-state level of nsb, while β is chosen to minimize the approximation error.

Compared to the standard Campbell and Shiller (1988b) approximation, this framework accommodates net government surpluses that turn negative, as in our sample period. Moreover, US net surplus has been historically volatile over extended periods, which proves to be a challenge for some of the alternative log-linearization approaches (Larrain and Yogo, 2008; Berndt et al., 2012). Furthermore, this approach allows a direct interpretation of nsb as a measure of fiscal space: high nsb periods are those when the government is in a strong fiscal position, while low nsb reflects

periods when the government is in a weak fiscal position.

Using this method, an approximation of the net government surplus growth Δns is

$$\Delta ns_{t+1} = \frac{1}{1-\beta}\Delta\tau_{t+1} - \frac{\beta}{1-\beta}\Delta s_{t+1}, \tag{10}$$

which is also a linear combination of tax and spending growth rate $\Delta\tau$, Δs. This allows a separate identification of the impact from tax and government spending, especially related to the recent non-stationary path of fiscal imbalance, as pointed out by Campbell et al. (2023).

To maintain consistency throughout our expressions, we apply the same framework to the private sector, which implies a log-linearization for the dividend-to-asset ratio $da_t = \log\left(1 + \frac{D_t}{A_t}\right)$.

The foreign sector necessitates further modifications to the framework, as not only net imports are negative in some periods, but also net foreign assets can be close to zero or negative when net foreign assets and liabilities are close to each other. These features present a challenge for the typical log-linearization for the external sector Gourinchas and Rey, 2007a. To address this issue, we separately approximate the dynamics of foreign assets and liabilities and regard net foreign asset values as a portfolio of the two accounts. This enables us to obtain a consistent log-linearization for the valuation ratio in the foreign sector, $nma = \log\left(1 + \frac{M_t - X_t}{F_t}\right)$, which captures external sector imbalances in a similar way as nsb captures fiscal imbalances in the public sector.

The linearization framework implies the following budget constraints (BCs) relating the valuation ratio to return and cash-flow growth:

$$da_t = (1-\rho^A)r^A_{t+1} - (1-\rho^A)\Delta d_{t+1} + \rho^A da_{t+1} \tag{11}$$

$$nsb_t = (1-\rho^B)r^B_{t+1} - (1-\rho^B)\Delta ns_{t+1} + \rho^B nsb_{t+1} \tag{12}$$

$$nma_t = (1-\rho^X)r^X_{t+1} - (1-\rho^X)\Delta nm_{t+1} + \rho^X nma_{t+1}. \tag{13}$$

These expressions closely resemble those from the standard Campbell and Shiller (1988b) expansion, although they are obtained using a modified approximation that accommodates negative and volatile cash flows. These intertemporal relations have been the foundation of separate strands of literature that analyze the predictability of equity returns, Treasury returns, and foreign asset returns. The conditioning variables used are either the within-sector conditioning variables or macroeconomic quantities.

2.2 The Aggregate Budget Constraint (ABC) and Cross-Sector Predictability

We can apply the log-linearization to the aggregate wealth portfolio, defined in Eq.(7). This yields the following expression for the consumption-to-wealth ratio, cw_t:

$$cw_t = \left(1-\rho^W\right)\sum_{j=0}^{\infty}\left(\rho^W\right)^j\left(r^W_{t+j+1} - \Delta c_{t+j+1}\right) \tag{14}$$

As we show in the appendix, the consumption-to-wealth ratio is a weighted average of the state

variables of the individual sectors, or

$$cw_t = \alpha_0 + w_A da_t + w_B nsb_t + w_X nma_t, \tag{15}$$

where w_A, w_B, w_X are the steady-state share of total wealth.

Eq. (15) represents an aggregate budget constraint (ABC) that ties the predictability of each sector together by enforcing the weighted average of the three valuation ratios in equations Eqs. (11)–(13). We provide detailed derivations in the appendix and show that

$$\begin{aligned} cw_t = w_A da_t + w_B nsb_t + w_X nma_t = \text{constant} &+ w_A \left(1-\rho^A\right) \sum_{j=0}^{\infty} \left(\rho^A\right)^j \left(r^A_{t+j+1} - \Delta d_{t+j+1}\right) \\ &+ w_B \left(1-\rho^B\right) \sum_{j=0}^{\infty} \left(\rho^B\right)^j \left(r^B_{t+j+1} - \Delta ns_{t+j+1}\right) \\ &+ w_X \left(1-\rho^X\right) \sum_{j=0}^{\infty} \left(\rho^X\right)^j \left(r^X_{t+j+1} - \Delta nm_{t+j+1}\right). \end{aligned} \tag{16}$$

In other words, if there is an imbalance in one sector, captured by a change in the state variable in that sector, this change will affect the consumption-to-wealth ratio and could generate a correspondence to future asset returns and cash flow growth rates in all the sectors.

If there is a single common factor driving the valuation dynamics across all three sectors, then each sector-specific budget constraint – Eq. (11) through Eq. (13) – is simply a re-statement of the aggregate consumption-wealth dynamic in Eq. (15). In this case, return predictability observed in the literature for each individual sector would merely reflect a shared systematic risk factor. However, if risks originate in one specific sector and propagate to others, the consumption-wealth identity aggregates heterogeneous sources of variation, and cross-sector predictability becomes empirically necessary to capture the full transmission mechanism.

To illustrate, consider a fiscal shock such as an unanticipated increase in taxes. This alters the net government surplus and hence the valuation ratio nsb_t. Through the aggregate budget constraint in Eq. (15), this shock shifts the consumption-to-wealth ratio and, consequently, affects equilibrium valuation across all sectors. A rise in expected taxes can depress private consumption, which in turn affects firm cash flows, dividend expectations, and the equity premium. Similarly, shifts in fiscal policy may influence the trade balance through changes in domestic demand, thereby affecting the valuation of foreign assets. In other words, fluctuations in fiscal policy not only impact future Treasury returns and surplus growth but may also be transmitted across asset markets.

This transmission is particularly relevant in the context of Ricardian Equivalence (Barro, 1979), which posits that current fiscal deficits imply future tax liabilities. A lower surplus today signals future tax increases, which investors may perceive as raising the cost of capital for firms and thus demand higher risk premiums. These fiscal-induced risks may also bear on the pricing of foreign assets, especially when fiscal imbalances lead to changes in exchange rate expectations or capital flows. Given the rising prominence of debt sustainability debates in the U.S., identifying how fiscal

policy affects other sectors' asset returns becomes a first-order question for macro-finance.

The extent to which state variables in one sector forecast returns in other sectors – above and beyond their own – is the key empirical question we address. If cross-sector transmission exists, then we should observe cross-predictability: variation in a sector-specific valuation ratio (such as nsb_t) should forecast returns outside its own sector, even after controlling for all sectoral state variables. Identifying this cross-sector effect requires a joint empirical framework that imposes both the aggregate constraint in Eq. (15) and the within-sector constraints in Eqs. (11)–(13). Specifically, the coefficients predicting consumption growth must align with the weighted average of the sector-specific cash-flow predictors, while the return and cash-flow predictors must satisfy each sector's intertemporal budget constraint.

Our empirical strategy is designed to disentangle these cross-sector interactions in the presence of heterogeneous risks. We adopt a system approach that jointly estimates return and cash-flow predictability equations across sectors, subject to the full set of intertemporal constraints. This enables us to trace how shocks propagate across the system and to attribute observed predictability to sector-specific or shared risk sources. Ultimately, our goal is to develop a more comprehensive and accurate framework for asset pricing that takes into account the complex interdependencies between different sectors of the economy.

Our approach differs from that of Lettau and Ludvigson (2001) in several respects. Lettau and Ludvigson (2001) use the deviation of log consumption from its long-run co-integrated relationship with labor income and wealth to predict equity returns. Their analysis focuses on the private sector and distinguishes between labor and non-labor income. Their primary goal was not to study cross-sector predictability, but rather to construct a state variable capturing expected equity return variation while controlling for labor income effects. In contrast, we build on the same consumption-wealth dynamics to investigate interactions across multiple sectors and examine predictability across them. This allows us to gain insights into the drivers of asset pricing and the interplay between sector-specific returns and cash flow growth rates.

3 Empirical Analysis

We apply the framework outlined above to US data. Section 3.1 describes the data source and presents summary statistics. Section 3.2 discusses our estimation strategy. Sections 3.3, 3.4, and 3.5 contain our main empirical results. In Section 3.6, we present additional econometric analyses and diagnostics.

3.1 Data Description and Stationarity Tests

We use quarterly U.S. data from 1952 to 2021. All returns and cash-flow series are expressed in real terms by deflating nominal values with the CPI.

For the private sector, we use equity returns and dividend growth from the CRSP value-weighted index covering NYSE, NASDAQ, and AMEX. Dividends are constructed as the trailing four-quarter

sum of payments and reinvested at the market rate, but our results are not sensitive to the reinvestment assumption. Quarterly dividend growth is defined as the log difference of annual dividends. To facilitate comparison with the equity predictability literature, we denote equity returns by r^E and the valuation ratio by dp, constructed as $\log(1 + D/P)$ for consistency with the linearization in Section 2.

Public-sector data are obtained from the NIPA tables. Following Berndt et al. (2012), taxes correspond to total receipts, and government spending equals current expenditures plus gross investment and capital transfers, net of depreciation and interest payments. The market value of public debt is taken from Hall and Sargent (2011) and Hall et al. (2018).

For the external sector, we follow Gourinchas and Rey (2007b) and Gourinchas et al. (2010) to construct gross positions, flows, and returns using data from the Flow of Funds Accounts, the Federal Reserve, and the BEA. The sample includes 13 advanced economies.

Table 1 reports summary statistics. Panel A shows that equity returns have the highest mean and volatility, while Treasury returns are substantially lower; foreign asset returns lie in between. Cross-asset return correlations are modest at the quarterly frequency. Panel B reports real cash-flow growth rates for dividends, taxes, government spending, imports, and exports. Average growth rates range from 2% to 4%, with dividends exhibiting the highest volatility. Cross-sector correlations are generally low, except between foreign inflows and outflows.

Panel C reports statistics for the valuation ratios that serve as predictors: the dividend–price ratio dp, the net surplus-to-debt ratio nsb, and the net foreign asset ratio nma. We also report the consumption–wealth ratio cw, which enters the aggregate budget constraint. All ratios are highly persistent, with univariate autoregressive roots ranging from 0.891 for nsb to 0.987 for dp, while cross-sector correlations remain small.

To assess stationarity, we report Augmented Dickey–Fuller (ADF) and KPSS tests. For dp, both tests indicate behavior close to a unit root. By contrast, nma clearly rejects the ADF null and does not reject KPSS. The net surplus-to-debt ratio also appears stationary, with an ADF statistic of −2.669 and a KPSS statistic of 0.679. Additional stationarity tests yield similar conclusions and suggest that nonstationarity concerns are primarily confined to dp. We discuss the implications for inference in Section 3.6.2.

We further examine the joint dynamics of the valuation ratios $x_t = [dp_t\, nsb_t\, nma_t]'$ using a VAR(1). Panel D of Table 1 reports the estimated coefficient matrix and residual correlations. The diagonal elements closely match the univariate autoregressive estimates, while off-diagonal coefficients and residual correlations are small and statistically insignificant.

Figure 1 plots the valuation ratios at annual frequency. The dividend–price ratio declines sharply after the mid-1990s; the net surplus-to-debt ratio turns persistently negative after the Global Financial Crisis; and the net foreign asset ratio exhibits larger fluctuations over the sample. While temporary co-movements across sectors are evident, they are not persistent, consistent with the weak unconditional correlations. In Section 3.6.2, we examine whether low-frequency shifts in these ratios affect our predictive results.

TABLE 1: **Summary Statistics of Returns, Cashflows, and Ratios**: This table presents summary statistics for real returns, cash-flow growth rates, and financial ratios. Panel A reports returns, where r^E is the equity return, r^D is the debt return, r^{XA} is the return on foreign assets, and r^{XL} is the return on foreign liabilities. Panel B presents growth rates in cash flows, including dividends (Δd), taxes (Δt), government expenditure (Δg), total imports (Δm), and total exports (Δx). Panel C shows valuation ratios: dp (dividend-to-price ratio), nsb (surplus-to-public debt ratio), nma (net import to net foreign asset ratio), and cw (consumption-to-wealth ratio). For all variables, we report the mean, standard deviation (Std), first-order autoregressive root (AR1), and test statistics from the Augmented Dickey-Fuller (ADF) and Kwiatkowski–Phillips–Schmidt–Shin (KPSS) tests. Panel D reports the autoregressive matrix Φ from a VAR(1) estimated on the three valuation ratios. The right-hand side of each panel reports correlations among the corresponding variables, except in Panel D where correlations refer to the residuals from the VAR(1) model. Data are quarterly, 1952-2021.

	Mean, Std, AR1				Correlations		
Panel A: Returns							
	r^E	r^D	r^{XA}	r^{XL}		r^D	r^X
Mean	0.071	0.018	0.048	0.042	r^E	0.037	-0.117
Std	0.169	0.039	0.098	0.096	r^D		-0.009
AR1	0.041	0.070	0.129	0.081			
Panel B: Cash-flows							
	Δd	Δt	Δg	Δm	Δx	Δns	Δnm
Mean	0.021	0.027	0.029	0.044	0.040	Δd -0.091	-0.175
Std	0.160	0.074	0.110	0.079	0.080	Δns	0.040
AR1	-0.008	-0.183	-0.250	0.004	-0.080		
Panel C: Valuation Ratios							
	dp	nsb	nma	cw		nsb	nma	cw
Mean	0.031	-0.013	0.062	0.150	dp	0.015	0.110	0.656
Std	0.006	0.027	0.069	0.007	nsb		0.092	0.313
AR1	0.987	0.891	0.936	1.004	nma			0.261
ADF	-1.416	-2.669	-4.448	0.234				
KPSS	2.352	0.679	0.217	1.330				
Panel D: Valuation Ratios, VAR(1)							
	dp	nsb	nma			nsb	nma
dp	0.986	-0.012	-0.085		dp	-0.123	-0.037
nsb	0.003	0.889	-0.033		nsb		0.074
nma	0.001	0.006	0.938				

3.2 Estimation Procedure

We begin by estimating standard predictive regressions that replicate own-sector return and cash-flow predictability using a sector's valuation ratio, following the framework in Section 2. We then study cross-sector effects by estimating multivariate OLS regressions that include valuation ratios from all sectors, augmented with standard controls for equity and bond risk premia. Finally, we implement a constrained estimation that jointly imposes the sector budget constraints (BC)

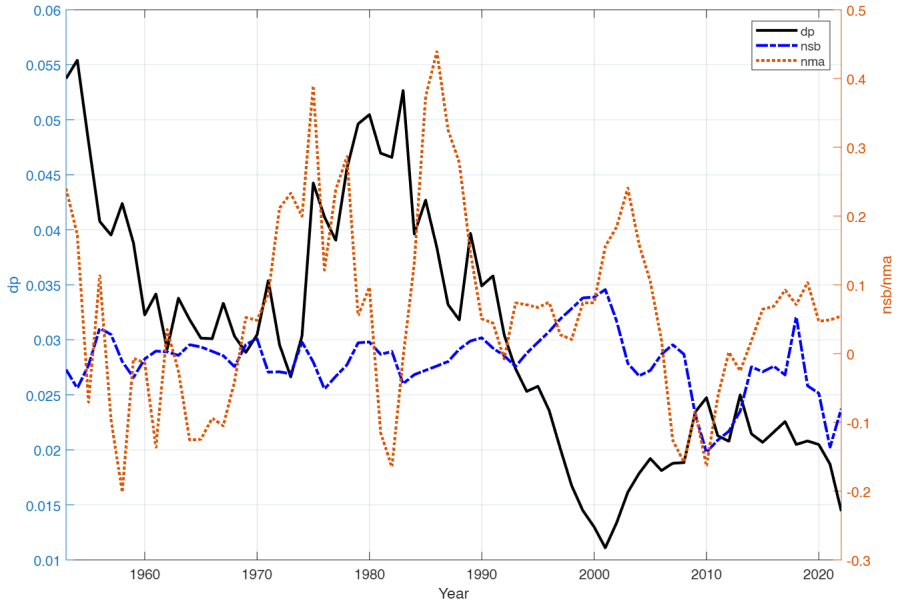


FIGURE 1: **Valuation ratios.** This figure plots the end-of-year time series of the dividend-to-price ratio (*dp*, left Y-axis), the net surplus-to-public-debt ratio (*nsb*, right Y-axis), and the net import-to-net-foreign-asset ratio (*nma*, right Y-axis) over the sample period 1952–2021.

and the aggregate budget constraint (ABC). Since our focus is on the gains from cross-sector restrictions, we summarize the constrained approach here and provide derivations and additional details in Online Appendix Section B.

The present-value identities imply predictive regressions relating sectoral returns, cash-flow growth, and valuation ratios to lagged valuation ratios from all sectors and controls. Let $\boldsymbol{b_r}$, $\boldsymbol{b_{cf}}$, and $\boldsymbol{\phi}$ denote the vectors of slope coefficients in these regressions. The sector budget constraints in Eqs. (11)–(13) imply the linear restrictions

$$\text{BC:} \qquad \boldsymbol{C} = \text{diag}(1-\rho^E, 1-\rho^B, 1-\rho^X)(\boldsymbol{b_r} - \boldsymbol{b_{cf}}) + \text{diag}(\rho^E, \rho^B, \rho^X)\boldsymbol{\phi}, \tag{17}$$

where $\boldsymbol{C} = [\boldsymbol{I}(N),\ \boldsymbol{0}_{N\times(M-N)}]$, with N denoting the number of sectors and $M-N$ the number of controls. In addition, the aggregate budget constraint implies

$$\text{ABC:} \qquad \boldsymbol{b_{cw}} = \boldsymbol{w}'\boldsymbol{\phi}, \tag{18}$$

where $\boldsymbol{b_{cw}}$ is the vector of slope coefficients in the regression of the future consumption-to-wealth ratio on the same predictors and $\boldsymbol{w}$ collects the aggregation weights from Eq. (15).

We estimate the system using constrained GMM, stacking the OLS moment conditions across sectors and imposing the BC and ABC as linear restrictions on the slope coefficients. Because the constraints can be written as $\boldsymbol{R}'\widehat{\boldsymbol{\beta}}_{\text{cgmm}} - \boldsymbol{c} = 0$, the constrained GMM estimator admits a closed-form expression as the unconstrained OLS estimator plus a correction term that reflects the imposed

restrictions (Online Appendix Eq. (B14)). This representation clarifies that the estimator balances the unrestricted estimates with the cross-equation restrictions and provides a natural chi-square test of the constraints based on the difference in objective functions.

Imposing the BC and ABC can improve inference in two ways. When the restrictions hold, the constrained estimator is more efficient and yields smaller standard errors. When the restrictions are misspecified but deviations are transitory, the constraints can also reduce bias in finite samples. Online Appendix Section B4 quantifies these variance and bias reductions in Monte Carlo simulations calibrated to our data. Because estimating the weight matrix may attenuate these gains, the net benefits of imposing the constraints remain an empirical question, which we address in the results below.

3.3 Univariate Estimates

Table 2 reports OLS estimates for univariate predictive regressions of returns and cash-flow growth in the private, public, and foreign sectors using each sector's valuation ratio. Slope coefficients and R^2 are reported at one-quarter ($h = 1$) and one-year ($h = 4$) horizons, with t-statistics based on Newey and West (1987) standard errors in parentheses.

For the private sector, the dividend-to-price ratio, dp, significantly predicts equity returns, r^E, consistent with the Campbell and Shiller (1988a) decomposition. Statistical significance increases at the one-year horizon ($t = 1.86$, $R^2 = 0.04$). dp does not meaningfully predict dividend growth at either horizon, consistent with Cochrane (2008) and Lettau and Van Nieuwerburgh (2008), confirming that using $da = \log(1 + D/P)$ does not alter the extent of predictability.

In the public sector, the log net surplus-to-debt ratio, nsb, is a strong and significant predictor of both government debt returns and cash-flow (i.e., net surplus or deficit) growth. Compared to equity, cash-flow predictability in the public sector is stronger, as the R^2s in the Δns regression are much larger at 0.05 (for $h = 1$) and 0.09 (for $h = 4$) compared to the corresponding return regression figures. This result aligns with evidence from Berndt et al. (2012), which shows that the bulk of fiscal shocks are absorbed by the surplus channel and to a much lesser extent through the debt valuation channel. The one-year estimate also aligns with that reported by Campbell et al. (2023) from an augmented VAR setting.

For the foreign sector, the log net import-to-foreign asset ratio, nma, significantly forecasts foreign investment returns: the slope is 0.04 ($t = 2.72$) at one quarter and 1.31 ($t = 2.96$, $R^2 = 0.11$) at one year. Cash-flow predictability is weak, with coefficients correctly signed but imprecisely estimated, consistent with Gourinchas and Rey (2007b).

3.4 Multivariate OLS Estimates

Table 3 presents the estimates from multivariate regressions that augment the unrestricted model with the valuation ratios of the other two sectors. These results provide prima facie evidence of cross-sector predictability and inform us about how shocks in one sector propagate to others.

TABLE 2: **Univariate predictive regressions**: The table presents the OLS estimates from regressing future h-period log returns and log cash-flow growth rates of the private sector (equity), public sector, and foreign sector on their current respective valuation ratios: the log dividend-to-price ratio dp, the log net surplus-to-debt ratio nsb, and the log net import-to-foreign asset ratio nma, as defined in section 2.1. From the quarterly log returns r_t series, we construct future h-period log returns as $\sum_{k=1}^{h} \rho^{k-1} r_{t+k}$. Future h-period log cash-flow growth is constructed analogously. In parentheses underneath the estimates, we report t-statistics based on Newey and West (1987) standard errors with 8 lags. Boldface coefficients are significant at the 10% confidence level. The coefficients for Δns are divided by 100. Data are quarterly, 1952-2021.

	1-quarter horizon ($h = 1$)						4-quarter horizon ($h = 4$)					
	Private		Public		Foreign		Private		Public		Foreign	
	r^E	Δd	r^D	Δns	r^X	Δnm	r^E	Δd	r^D	Δns	r^X	Δnm
dp	**0.801**	0.347					**2.986**	-0.031				
	(1.800)	(1.044)					(1.859)	(-0.027)				
nsb			**0.040**	**-1.096**					**0.185**	**-2.391**		
			(2.550)	(-2.559)					(3.401)	(-3.226)		
nma					**0.343**	-0.094					**1.312**	-0.432
					(2.715)	(-0.690)					(2.958)	(-0.989)
R^2	0.011	0.002	0.008	0.054	0.037	0.003	0.039	0.000	0.047	0.091	0.114	0.013

In Panel A of the table, the returns and cash flows of a given sector are regressed on all three lagged valuation ratios. Comparing these estimates with those in Panel A of Table 2, several conclusions emerge. The net surplus-to-debt ratio negatively predicts equity returns at both short and medium horizons, with coefficients of -0.21 and -0.76, both statistically significant at the 1% level or better. Including nsb has an economically substantial impact on the regression R^2 at both horizons, increasing it to 0.03 and 0.09, respectively. This result suggests that a higher than average surplus relative to outstanding debt today predicts lower equity returns in the future. Interestingly, the significance of nsb does not diminish the role of the dividend-to-price ratio dp, whose t-statistics are somewhat higher than in the univariate setting. Moreover, nsb negatively predicts dividend growth, with a coefficient of -0.122, though this result is marginally insignificant. Taken together, these findings suggest that nsb captures a component of equity expected returns that positively correlates with expected dividend growth and dilutes the signal in the dividend-to-price ratio. At the one-quarter horizon, the predictability of the surplus-to-debt ratio for equity returns is the only significant cross-sector effect, while at the annual horizon, the ratio also forecasts a decrease in risk premia in the foreign sector.

In Panel B of the table, we report analogous estimates after adding three control variables to all regressions: the log tax-to-GDP ratio (ty), which Campbell et al. (2023) argue ensures that the system accounts for the stationary relationship between tax and output; the Cochrane and Piazzesi (2005) tent-shaped factor for bond risk premia (CP), which is re-estimated using Fama-Bliss data over our sample period; and the inflation rate ($infl$), which is included in the benchmark model of Berndt et al. (2012). Overall, the controls enter with significant loadings across all equations. In particular, ty negatively predicts foreign asset returns at both horizons and positively forecasts annual returns to the public sector, consistent with evidence from Campbell et al. (2023). The

CP factor captures an important component of expected returns to government debt, confirmed by the positive and highly significant coefficients at both the one-quarter and one-year horizons, where its inclusion raises the R^2 to 0.30, as also found by Cochrane and Piazzesi (2005). Finally, inflation negatively forecasts short-horizon cash-flow growth in the foreign sector and annual equity returns. Since our cash-flow and return series are expressed in real terms, these findings are not mechanically driven by persistence in inflation.

Despite the addition of these controls, the negative relationship between the current net surplus-to-debt ratio and future equity returns remains statistically and economically intact. The net surplus-to-debt ratio also contains relevant information on bond risk premia r^D on top of the yield factor CP. Furthermore, the coefficients and t-statistics for the dividend-to-price ratio increase further in this augmented setting, suggesting that the controls help mitigate endogeneity concerns arising from correlated movements in expected returns and dividend growth. However, in contrast to Panel A, nsb no longer predicts foreign asset returns, indicating that this effect is not robust to controlling for the information contained in other variables.

3.5 Multivariate Estimates Under the BC and ABC

The analysis in the previous section overlooks the restrictions imposed by budget constraints. Therefore, we carry a constrained estimation that ensures consistency across sectors by first applying the budget constraint (BC) for each sector, as outlined in Eqs. (11)–(13), and then additionally enforcing the cross-sector constraint (ABC) in Eq.(15).

In Panel A of Table 4, we present GMM estimates for the return and cash-flow regressions of the three sectors that adhere to the BC. Compared to Table 3, we observe that enforcing this constraint impacts both the coefficients and, more importantly, their statistical significance, which generally increases. This is particularly evident for the dividend-to-price ratio, whose t-statistics in the equity return regression rise to 2.40 at the one-quarter horizon and 2.80 at the annual horizon, as well as for inflation. Notably, the enhanced significance of these variables does not come at the expense of nsb, which continues to robustly negatively predict equity returns, with t-statistics that are larger in absolute value. Its role in predicting government debt returns and, in particular, growth in deficit or surplus is also strengthened. Among the valuation ratios of the sectors, the net surplus-to-debt ratio stands out as the only significant cross-sector predictor, indicating that the impact of fiscal shocks on future private-sector returns is consistent with the budget constraint.

In the last row of the panel, we report the p-value for the chi-square test of the budget constraint (BC) across all sectors. At both horizons, these p-values are well beyond standard critical values, so we cannot reject the null hypothesis that the BCs hold in the data.

In Panel B of the table, we impose both the sectors' budget constraints and the aggregate budget constraint (ABC). Enforcing the constraints again leads to a general increase in the statistical significance of the coefficients at both horizons. For example, the coefficient of nsb in predicting Δns is now −2.119 (up from −2.076) and is much more precisely estimated, with a t-statistic of −4.340 (up from −3.630). Its relationship to future equity returns is slightly weaker at the

TABLE 3: **Multivariate predictive regressions**: The table presents the OLS estimates from regressing future h-period log returns and log cash-flow growth rates of the private sector (equity), public sector, and foreign sector on all current valuation ratios. Variable definitions follow Table 2. In Panel A the regressors are the valuation ratios only, while Panel B adds the following controls: the log tax-to-GDP ratio (ty), the Cochrane and Piazzesi (2005) factor (CP), and the inflation rate ($infl$). In parentheses underneath the estimates, we report t-statistics based on Newey and West (1987) standard errors with 8 lags. Boldface coefficients are significant at the 10% confidence level. The coefficients for Δns are divided by 100. Data are quarterly, 1952–2021.

	1-quarter horizon (h = 1)						4-quarter horizon (h = 4)					
Panel A: Multivariate regressions, valuation ratios only												
	Private		Public		Foreign		Private		Public		Foreign	
	r^E	Δd	r^D	Δns	r^X	Δnm	r^E	Δd	r^D	Δns	r^X	Δnm
dp	**0.822**	0.312	0.037	-0.119	0.228	1.027	**2.868**	-0.255	0.192	3.193	0.439	3.654
	(1.883)	(0.963)	(0.234)	(-0.094)	(0.181)	(0.571)	(1.948)	(-0.245)	(0.345)	(0.874)	(0.092)	(0.641)
nsb	**-0.209**	-0.122	**0.040**	**-1.111**	-0.407	-0.169	**-0.758**	-0.380	**0.192**	**-2.421**	**-2.031**	-0.406
	(-2.396)	(-1.619)	(2.478)	(-2.679)	(-1.432)	(-0.532)	(-2.524)	(-1.442)	(3.283)	(-3.246)	(-2.249)	(-0.362)
nma	-0.009	0.030	0.000	0.062	**0.355**	-0.098	-0.018	0.113	-0.020	0.239	**1.376**	-0.451
	(-0.233)	(0.996)	(-0.005)	(0.716)	(2.927)	(-0.745)	(-0.143)	(1.274)	(-0.405)	(0.842)	(3.285)	(-1.065)
R^2	0.029	0.011	0.013	0.055	0.045	0.006	0.090	0.030	0.053	0.108	0.149	0.021
Panel B: Multivariate regressions, valuation ratios plus controls												
	Private		Public		Foreign		Private		Public		Foreign	
	r^E	Δd	r^D	Δns	r^X	Δnm	r^E	Δd	r^D	Δns	r^X	Δnm
dp	**1.015**	0.270	0.023	-0.652	-0.915	2.095	**3.799**	-0.427	0.338	3.349	-0.654	4.864
	(2.206)	(0.686)	(0.163)	(-0.482)	(-0.763)	(1.099)	(2.653)	(-0.372)	(0.809)	(0.902)	(-0.140)	(0.921)
nsb	**-0.203**	-0.063	**0.037**	**-1.089**	0.064	-0.284	**-0.678**	-0.145	**0.112**	**-1.916**	-0.023	-0.536
	(-2.496)	(-0.948)	(1.920)	(-1.874)	(0.277)	(-0.758)	(-2.500)	(-0.616)	(1.997)	(-1.940)	(-0.029)	(-0.375)
nma	-0.001	0.042	-0.006	0.046	**0.393**	-0.095	-0.002	0.130	-0.049	0.262	**1.570**	-0.574
	(-0.035)	(1.312)	(-0.450)	(0.481)	(3.371)	(-0.743)	(-0.016)	(1.341)	(-1.408)	(0.911)	(4.256)	(-1.446)
ty	0.009	-0.089	0.004	-0.081	**-0.786**	0.265	0.038	-0.305	**0.125**	-0.573	**-2.564**	0.371
	(0.094)	(-1.000)	(0.176)	(-0.163)	(-2.734)	(0.757)	(0.144)	(-1.324)	(1.756)	(-0.769)	(-2.984)	(0.332)
CP	-0.019	-0.287	**0.310**	0.076	-0.059	1.073	1.076	0.344	**1.534**	2.478	0.228	**9.117**
	(-0.054)	(-1.033)	(2.752)	(0.079)	(-0.050)	(1.045)	(0.954)	(0.462)	(4.564)	(0.896)	(0.067)	(3.005)
$infl$	-0.997	-0.338	0.096	2.391	1.179	**-4.141**	**-4.350**	-0.640	-0.062	-3.499	-7.195	-3.854
	(-1.322)	(-0.766)	(0.757)	(1.375)	(0.718)	(-2.102)	(-3.045)	(-0.592)	(-0.174)	(-1.269)	(-1.603)	(-0.903)
R^2	0.039	0.017	0.058	0.061	0.068	0.028	0.147	0.045	0.300	0.128	0.219	0.081

one-quarter horizon but stronger at the one-year horizon. Interestingly, we observe that the net import-to-foreign asset ratio (nma) now significantly predicts annual cash-flow growth rates in the foreign sector with the expected negative sign. This result expands on the evidence presented in Gourinchas and Rey (2007b) and confirms that enforcing consistency across sectors improves inference. Overall, the other conclusions we draw from this setting are quantitatively similar to those from Panel A, and the p-values again cannot reject the null hypothesis.

3.6 Additional Econometric Analyses and Diagnostics

We conduct several analyses to verify that our findings are robust to econometric issues that typically arise in forecasting regressions of asset returns with persistent regressors. We focus our tests on the multivariate OLS specification in Panel A of Table 3 at the non-overlapping 1-quarter horizon, which actually delivers the most conservative figures (in terms of statistical significance)

TABLE 4: **Multivariate predictive regressions with budget constraints**: The table presents the constrained GMM estimates from regressing future h-period log returns and log cash-flow growth rates of the private sector (equity), public sector, and foreign sector on all current valuation ratios. Variables definition follows from Table 2. In Panel A, we impose the budget constraints of Eq. (11)–(13) within each sector, while in Panel B we impose the aggregate budget constraint of Eq. (15) across the three sectors. In parentheses underneath the estimates, we report t-statistics based on Newey and West (1987) standard errors with 8 lags for the 1-quarter horizon and 12 lags for the 4-quarter horizon. Boldface coefficients are significant at the 10% confidence level. The coefficients for Δns are divided by 100. Data are quarterly, 1952–2021.

	1-quarter horizon (h = 1)						4-quarter horizon (h = 4)					
Panel A: Imposing BC												
	Private		Public		Foreign		Private		Public		Foreign	
	r^E	Δd	r^D	Δns	r^X	Δnm	r^E	Δd	r^D	Δns	r^X	Δnm
dp	**0.973**	0.229	0.010	-0.798	-0.393	2.567	**3.730**	-0.420	0.340	3.303	-0.650	5.600
	(2.396)	(0.644)	(0.071)	(-0.940)	(-0.388)	(1.513)	(2.800)	(-0.420)	(0.810)	(1.460)	(-0.140)	(1.300)
nsb	**-0.224**	-0.059	**0.036**	**-1.229**	0.084	-0.222	**-0.670**	-0.130	**0.110**	**-2.076**	-0.030	-0.580
	(-2.975)	(-0.988)	(2.065)	(-3.630)	(0.399)	(-0.671)	(-2.550)	(-0.650)	(2.040)	(-3.630)	(-0.040)	(-0.480)
nma	0.009	**0.053**	-0.002	0.070	**0.420**	-0.055	0.000	0.130	-0.050	**0.291**	**1.580**	-0.550
	(0.281)	(1.791)	(-0.170)	(1.176)	(3.692)	(-0.463)	(-0.010)	(1.460)	(-1.490)	(1.890)	(4.510)	(-1.530)
ty	0.018	-0.123	0.004	-0.001	**-0.721**	0.228	0.040	-0.310	**0.120**	-0.485	**-2.580**	0.400
	(0.199)	(-1.639)	(0.161)	(-0.005)	(-2.896)	(0.710)	(0.150)	(-1.570)	(1.720)	(-1.090)	(-3.150)	(0.390)
CP	-0.010	-0.257	**0.264**	-0.138	-0.922	0.953	1.050	0.390	**1.530**	2.296	0.220	**8.820**
	(-0.030)	(-1.023)	(2.523)	(-0.209)	(-0.918)	(1.023)	(0.960)	(0.610)	(4.650)	(1.340)	(0.070)	(3.330)
$infl$	-1.065	-0.260	0.084	**3.589**	-0.453	**-4.652**	**-4.330**	-0.610	-0.060	**-3.304**	-7.020	-4.380
	(-1.440)	(-0.632)	(0.671)	(3.269)	(-0.293)	(-2.578)	(-3.390)	(-0.690)	(-0.170)	(-1.960)	(-1.550)	(-1.230)
	p-value constraint:		0.699				p-value constraint:		0.479			
Panel B: Imposing BC+ABC												
	Private		Public		Foreign		Private		Public		Foreign	
	r^E	Δd	r^D	Δns	r^X	Δnm	r^E	Δd	r^D	Δns	r^X	Δnm
dp	**1.012**	0.269	0.017	-0.841	-0.936	2.283	**3.790**	-0.490	0.330	2.995	-0.760	5.170
	(2.629)	(0.761)	(0.121)	(-1.026)	(-0.964)	(1.345)	(2.960)	(-0.490)	(0.780)	(1.490)	(-0.170)	(1.210)
nsb	**-0.205**	-0.057	**0.034**	**-1.185**	0.056	-0.284	**-0.680**	-0.120	**0.110**	**-2.119**	-0.020	-0.630
	(-2.735)	(-0.969)	(1.950)	(-3.615)	(0.267)	(-0.887)	(-2.670)	(-0.610)	(2.050)	(-4.340)	(-0.020)	(-0.540)
nma	-0.001	0.041	-0.005	0.049	**0.391**	-0.085	0.010	0.120	-0.050	**0.257**	**1.570**	**-0.590**
	(-0.036)	(1.423)	(-0.429)	(0.853)	(3.592)	(-0.734)	(0.050)	(1.470)	(-1.520)	(1.910)	(4.570)	(-1.700)
ty	0.012	-0.093	0.006	-0.029	**-0.800**	0.264	0.030	**-0.320**	**0.120**	-0.383	**-2.560**	0.450
	(0.128)	(-1.273)	(0.256)	(-0.094)	(-3.221)	(0.831)	(0.100)	(-1.690)	(1.710)	(-0.910)	(-3.170)	(0.440)
CP	-0.019	-0.269	**0.302**	0.080	0.038	0.966	0.970	0.430	**1.530**	2.479	0.200	**8.630**
	(-0.060)	(-1.079)	(2.972)	(0.135)	(0.040)	(1.031)	(1.000)	(0.700)	(4.640)	(1.600)	(0.060)	(3.360)
$infl$	-1.001	-0.329	0.096	**2.769**	1.274	**-4.312**	**-4.290**	-0.530	-0.060	**-3.480**	-7.000	-4.440
	(-1.355)	(-0.809)	(0.786)	(2.777)	(0.874)	(-2.386)	(-3.720)	(-0.600)	(-0.160)	(-2.540)	(-1.610)	(-1.300)
	p-value constraint:		0.362				p-value constraint:		0.834			

for the role of nsb in equity predictability. Consequently, we concentrate the discussion around this evidence.

3.6.1 Small-sample bias adjustment

Predictive regressions often involve highly persistent regressors whose innovations correlate with return shocks, leading to small-sample bias in OLS estimates (Stambaugh, 1999). In univariate return regressions, this bias is typically positive, as shocks to expected returns raise valuation ratios while depressing contemporaneous returns (Boudoukh et al., 2022).

We assess the importance of small-sample bias in our multivariate system using the procedure of Amihud and Hurvich (2004), which extends bias correction to multiple predictors by augmenting the predictive regression with bias-corrected residuals from the autoregressive processes of the predictors. We assume that the vector $x_t = [dp_t\, nsb_t\, nma_t]'$ follows a stationary Gaussian VAR(1). Tests fail to reject that the off-diagonal elements of the autoregressive matrix are jointly zero (Wald statistic 10.73, p-value 9.72%), so we implement the diagonal-VAR correction described in Section IV.B of Amihud and Hurvich (2004).

Bias correction increases the autoregressive root of nsb from 0.891 to 0.904, still well below unity. We then estimate multivariate predictive regressions including the current valuation ratios and the bias-corrected residuals of all ratios. Standard errors follow Amihud and Hurvich (2004), adjusted for residual autocorrelation as in Amihud et al. (2008).

Panel A of Table 5 reports the bias-adjusted estimates. Correcting for small-sample bias substantially weakens the predictive power of dp for equity returns, reducing its coefficient to 0.451 and its t-statistic to 1.03. By contrast, the coefficient on nsb remains negative and statistically significant at the 5% level (−0.162, $t = -1.96$), while nma continues to display no predictive content. Across other equations, the bias adjustment has relatively modest quantitative effects.

3.6.2 Nonstationarities

The high persistence of valuation ratios raises concerns that the distribution of the conventional t-test for return predictability may deviate in small samples from its asymptotic approximation. We conduct three set of analyses to assess the robustness of our findings to nonstationarity concerns.

First, we bootstrap the small-sample distribution of the OLS estimator by simulating 10,000 samples calibrated to the estimated system in Table 3, with predictor residuals resampled jointly. Panel B of Table 5 reports bootstrapped t-statistics. While standard errors increase for some coefficients, nsb continues to negatively predict equity returns ($t = -2.291$, $p = 2.2\%$).

Second, following Campbell and Yogo (2006), we apply the DF–GLS test to nsb, obtaining a 95% confidence interval for its autoregressive root of $[0.831, 0.947]$, indicating stationarity. A univariate regression of equity returns on nsb yields a coefficient of −0.21 ($t = -2.21$), and the resulting 90% Bonferroni confidence interval $[-0.370, -0.047]$ confirms statistical significance.[6]

Third, we address the possibility that valuation ratios exhibit low-frequency shifts in their means. Lettau and Van Nieuwerburgh (2008) document a break in the mean of the dividend-to-price ratio in 1991. They propose adjusting the series by demeaning it separately before and after 1991, which restores its ability to consistently forecast equity returns. For public debt valuation, Jiang et al. (2024c) identify a break in the mean of the government debt-to-output ratio in 2007, following the sharp run-up in debt during the financial crisis. Structural break tests detect a break in nsb in 2007 and in nma in 1968Q1. Panel C of Table 5 reports estimates using break-adjusted series. While adjusting dp substantially strengthens its predictive power, the coefficient on nsb remains negative and statistically significant, with only modest quantitative changes. Online

[6] A similar interval obtains when approximating the multivariate system using residualized variables.

TABLE 5: **Multivariate predictive regressions, econometric extensions**: The table presents the OLS estimates from regressing future 1-quarter log returns and log cash-flow growth rates of the private sector (equity), public sector, and foreign sector on all current valuation ratios. Variables definition follows from Table 2. In Panel A, we account for the bias in predictive regressions using the multivariate procedure in Amihud and Hurvich (2004) and Amihud et al. (2008) for a diagonal Φ, as detailed in Section 3.6.1. In Panel B, t-statistics (in square brackets) are computed from the bootstrap procedure described in Section 3.6.2. In Panel C, we break-adjust dp in 1991Q4, nsb in 2006Q4, and nma in 1968Q1, as detailed in Section 3.6.2, and compute t-statistics based on Newey and West (1987) standard errors with 8 lags. Boldface coefficients are significant at the 10% confidence level. The coefficients for Δns are divided by 100. Data are quarterly, 1952–2021.

Panel A: Adjusting for small-sample bias

	Private		Public		Foreign	
	r^E	Δd	r^D	Δns	r^X	Δnm
dp	0.451	0.422	0.043	0.002	0.167	0.977
	(1.030)	(1.314)	(0.273)	(1.352)	(0.133)	(0.726)
nsb	**-0.162**	**-0.159**	**0.041**	**-0.963**	-0.436	0.005
	(-1.959)	(-2.152)	(2.466)	(-2.315)	(-1.540)	(0.022)
nma	0.016	0.017	0.000	-0.000	**0.341**	-0.026
	(0.413)	(0.578)	(0.028)	(-0.656)	(2.768)	(-0.196)
R^2	0.014	0.015	0.013	0.041	0.043	0.002

Panel B: Bootstrapped standard errors

	Private		Public		Foreign	
	r^E	Δd	r^B	Δns	r^X	Δnm
dp	**0.822**	0.312	0.037	-0.119	0.228	1.027
	[1.665]	[0.691]	[0.305]	[-0.082]	[0.158]	[0.677]
nsb	**-0.209**	-0.122	**0.040**	**-1.111**	-0.407	-0.169
	[-2.291]	[-1.386]	[1.919]	[-4.110]	[-1.582]	[-0.614]
nma	-0.009	0.030	0.000	0.062	**0.355**	-0.098
	[-0.227]	[0.761]	[-0.007]	[0.518]	[3.089]	[-0.791]
R^2	0.029	0.011	0.013	0.055	0.045	0.006

Panel C: Break-adjusting valuation ratios

	Private		Public		Foreign	
	r^E	Δd	r^D	Δns	r^X	Δnm
dp	**2.313**	**1.473**	0.096	**-3.446**	-2.360	0.981
	(3.071)	(2.376)	(0.359)	(-1.767)	(-1.283)	(0.311)
nsb	**-0.224**	-0.072	0.029	**-1.989**	**-0.930**	-0.517
	(-2.042)	(-0.795)	(0.981)	(-2.902)	(-2.580)	(-1.277)
nma	-0.003	0.033	-0.000	-0.014	**0.379**	-0.100
	(-0.055)	(1.018)	(-0.000)	(-0.139)	(2.918)	(-0.619)
R^2	0.053	0.021	0.004	0.103	0.069	0.010

Appendix Figure C.1 shows that allowing for breaks at alternative dates does not materially affect the predictive content of nsb.

3.6.3 Simulation exercise

We conduct a simulation exercise, following Cochrane (2008), to assess how likely it is to observe negative predictability from nsb when expected returns are fully captured by the dividend–price

ratio. We generate 10,000 samples matching the length of our data, consisting of equity returns, dividend growth, dp, and nsb. Both valuation ratios follow AR(1) processes, and equity returns are generated under the null that dp is the sole predictor, with its coefficient implied by the budget constraint and the empirical value of ρ^E. Innovations are drawn jointly from a multivariate normal distribution calibrated to the residual covariance matrix.

We consider four scenarios varying the persistence of dp and nsb, including cases with a unit root in dp and bias-adjusted persistence for nsb. Across all scenarios, negative predictability from nsb arises with low probability under the null (see Online Appendix Figure C.2). Increasing persistence or correcting for small-sample bias further reduces this likelihood. These results indicate that nsb contains predictive information beyond that implied by the within-sector budget constraint.

4 Theoretical Framework and Implications

Our results underscore the central role of (sectoral and aggregate) budget constraints in asset-return dynamics: a higher surplus-to-debt ratio predicts a lower equity risk premium. This pattern reflects joint movements in fiscal policy, government spending, and private-sector profits that are hard to infer from public-sector data alone. Bringing in private- and external-sector information, together with the sector constraints and the ABC, sharpens identification of how fiscal adjustments transmit across the economy. Empirically, the strongest cross-sector effect is that the net surplus-to-debt ratio (nsb_t) negatively predicts future equity returns. To interpret this evidence, we develop a framework that links fiscal policy to private profitability through accounting identities and output distortions from taxes and spending. We present only the elements needed for the empirical connection and defer derivations to Appendix A.

4.1 Output Distortion and Cash Flows

To match our empirical focus on fiscal spillovers to the equity risk premium, we consider a closed economy and abstract from the external sector and investment. The two-sector accounting identity is

$$C_t = Y_t - G_t = (Y_t - T_t) + (T_t - G_t) = D_t + NS_t,$$

where $D_t \equiv Y_t - T_t$ denotes after-tax profits and $NS_t \equiv T_t - G_t$ is the net surplus.

Following Barro (1979) and related work (e.g., Jiang et al. (2024a)), we model taxes as distortional and government spending as output-stimulating. Let the economy produce raw output Y_t^r each period, with fiscal policy affecting finalized output via a deadweight-loss wedge:

$$Y_t = Y_t^r \left(1 - \theta(TY_t^r, GY_t^r)\right). \tag{19}$$

The wedge depends on tax and spending shares of raw output, $TY_t^r = \frac{T_t}{Y_t^r}$, $GY_t^r = \frac{G_t}{Y_t^r}$. To separate

tax and spending channels in pricing outcomes, assume

$$\theta(TY^r, GY^r) = \theta_\tau(TY^r) - \theta_g(GY^r). \tag{20}$$

Consistent with Barro (1979), both θ_τ and θ_g are smooth, convex, and increasing, avoiding corner solutions. For tractability, we adopt quadratic forms:

$$\theta_\tau(TY^r) = \frac{1}{2}c^\tau\left(\frac{T}{Y^r}\right)^2, \quad \theta_g(GY^r) = \frac{1}{2}c^g\left(\frac{G}{Y^r}\right)^2, \quad \theta(TY^r, GY^r) = \frac{c^\tau}{2}(TY^r)^2 - \frac{c^g}{2}(GY^r)^2, \tag{21}$$

with $c^\tau, c^g > 0$, so taxes reduce output while spending raises it.

Households consume all output allocated by the government (goods-market clearing, as in a Lucas-tree economy):

$$C_t = Y_t - G_t = Y_t^r\left(1 - \theta(TY_t^r, GY_t^r)\right) - G_t = Y_t^r - Y_t^r\theta(TY_t^r, GY_t^r) - Y_t^r GY_t^r. \tag{22}$$

In asset markets, households hold public debt (a claim on net surplus) and private equity (a claim on dividends). Dividends are post-tax output:

$$D_t = Y_t - T_t = Y_t^r\left(1 - \theta(TY_t^r, GY_t^r)\right) - T_t = Y_t^r - Y_t^r\theta(TY_t^r, GY_t^r) - Y_t^r TY_t^r. \tag{23}$$

Eqs. (22)–(23) show that TY^r and GY^r govern bond and equity cash flows and aggregate consumption.

To isolate the role of distortions, note that if $\theta \equiv 0$ (so $Y_t = Y_t^r$), then Eq. (23) reduces to

$$D_t = Y_t - T_t = Y_t^r\left(1 - TY_t^r\right). \tag{24}$$

In that case, taxes affect dividends only mechanically through the accounting identity. With distortions, taxes (and spending) additionally shift both D_t and C_t through Eq. (21), amplifying the cash-flow effects. Below we use a consumption-based pricing-kernel approach and our GMM estimates to quantify these distortions in U.S. data.

4.2 Fiscal Policy Changes, Asset Valuation, and Cross-Sector Predictability

Motivated by the amplification channel above, we use a comparative-static exercise (as in Barro (1979)) to reconcile the magnitude of equity-return predictability from fiscal imbalances observed in the data.

The first ingredient is a steady-state equilibrium pricing benchmark linking (TY^r, GY^r) to asset values. The government takes the dynamics of raw output Y_t^r as given and chooses tax and spending shares (TY_t^r, GY_t^r) to maximize the present value of household consumption, subject to the public-sector budget constraint (as in Eq. (4)). Fiscal policy determines the allocation of output across sectors; households take consumption in Eq. (22) as given and price public debt and private equity using the CRRA consumption-based kernel. Appendix derives the government's optimal

policy and the implied risk premia for equity and Treasuries.

In steady state, asset valuation prices the exogenous growth in Y_t^r, generating systematic co-movement across sectors. This standard channel aligns with our GMM framework, which imposes the ABC in Eq. (18) to control for economy-wide variation in risk premia.

The second ingredient is an anticipated deviation from the steady-state fiscal path:

$$\widetilde{TY}_{t+1}^r = TY_{t+1}^r e^{u_{t+1}^\tau}, \quad \widetilde{GY}_{t+1}^r = GY_{t+1}^r e^{u_{t+1}^g}, \quad u_{t+1}^g = \alpha u_{t+1}^\tau. \tag{25}$$

The shocks (u^τ, u^g) capture changes in the expected future fiscal path (e.g., shifts in the budget constraint, political uncertainty, or related forces). We treat them as small deviations from steady state that shift equilibrium pricing; α governs the co-movement of tax and spending shocks.

Anticipated fiscal adjustments revise expectations of cash-flow growth and expected returns through output distortions and consumption-based pricing, generating cross-sector predictability via current price adjustments. In particular,

$$nsb_t \to \widetilde{nsb}_t, \qquad \mathbb{E}_t r_{t+1}^E \to \mathbb{E}_t \widetilde{r}_{t+1}^E,$$

and the corresponding expected-return shift can be summarized by

$$\mathbb{E}_t \widetilde{r}_{t+1}^E - \mathbb{E}_t r_{t+1}^E = b(r_{t+1}^E, nsb_t)\, \widetilde{nsb}_t.$$

Empirically, the regressions are run on the post-shock variables $(\widetilde{r}_{t+1}^E, \widetilde{nsb}_t)$; the predictive coefficient $b(r_{t+1}^E, nsb_t)$ maps anticipated fiscal adjustments into risk-premium changes, as captured by our GMM setup.

A key implication is that $b(r_{t+1}^E, nsb_t)$ depends on how fiscal policy affects output, governed by (c^τ, c^g). To match the persistence of fiscal shifts and business-cycle durations, assume investors expect the shocks in Eq. (25) to be i.i.d. and to persist for K periods, with cumulative shock $u_{t+1\ldots t+K}$.[7] Details are in Appendix. The resulting link between (c^τ, c^g) and the predictive slope is summarized in the following proposition.

Proposition 1. *If a fiscal shock persists for K periods, denoted as $u_{t+1\ldots t+K}$, and shocks at each period are i.i.d., then:*

1. *The single-period predictive coefficient $b(nsb_t, r_{t+1}^E)$ is determined by the output distortion parameters c^τ and c^g:*

$$b(nsb_t, r_{t+1}^E) = \frac{1-\beta}{(1-\rho^K)(1-\alpha\beta)} \left(-\frac{T}{D} - c^\tau \frac{T}{D} + \alpha c^g \frac{G}{D} \right), \tag{26}$$

where $\frac{T}{D}$, $\frac{G}{D}$, $\frac{T}{Y}$, and $\frac{G}{Y}$ are steady-state constants, and $\beta = \frac{G}{T} < 1$ is the stable spending-to-tax ratio used in constructing nsb_t.

[7] Alternatively, one can assume an autoregressive fiscal-shock process. Relative to the i.i.d. case, the cumulative pricing impact differs only by a scalar determined by the autoregressive coefficient.

2. *The time-series response of nsb to fiscal shocks is persistent:*

$$\mathbb{E}_t\left(\widetilde{nsb}_{t+1} - nsb_{t+1}\right) = \frac{1-\rho^{K-1}}{1-\rho^K}\left(\widetilde{nsb}_t - nsb_t\right). \tag{27}$$

3. *The cumulative expected return over the next K periods is predicted by nsb_t with the coefficient:*

$$b\left(nsb_t, \sum_{k=1}^{K} \rho^{k-1} r^E_{t+k}\right) = \frac{1-\rho^K}{1-\rho} b(nsb_t, r^E_{t+1}). \tag{28}$$

Eq. (26) implies that higher tax distortion (c^τ) makes $b(nsb_t, r^E_{t+1})$ more negative, while stronger spending stimulus (c^g) makes it more positive. Eq. (27) shows that persistent fiscal shocks generate persistent movements in *nsb*, and Eq. (28) implies stronger long-horizon predictability the longer the expected fiscal episode.

We calibrate K using the significance pattern of nsb_t in predicting cumulative returns $\sum_{k=1}^{K} \rho^{k-1} r^E_{t+k}$. In Figure 2, we plot $b\big(nsb_t, \sum_{k=1}^{K} \rho^{k-1} r^E_{t+k}\big)$ as K increases and set $K = 20$, where the long-horizon coefficient begins to level off. This suggests investors expect fiscal shifts to persist for about five years.

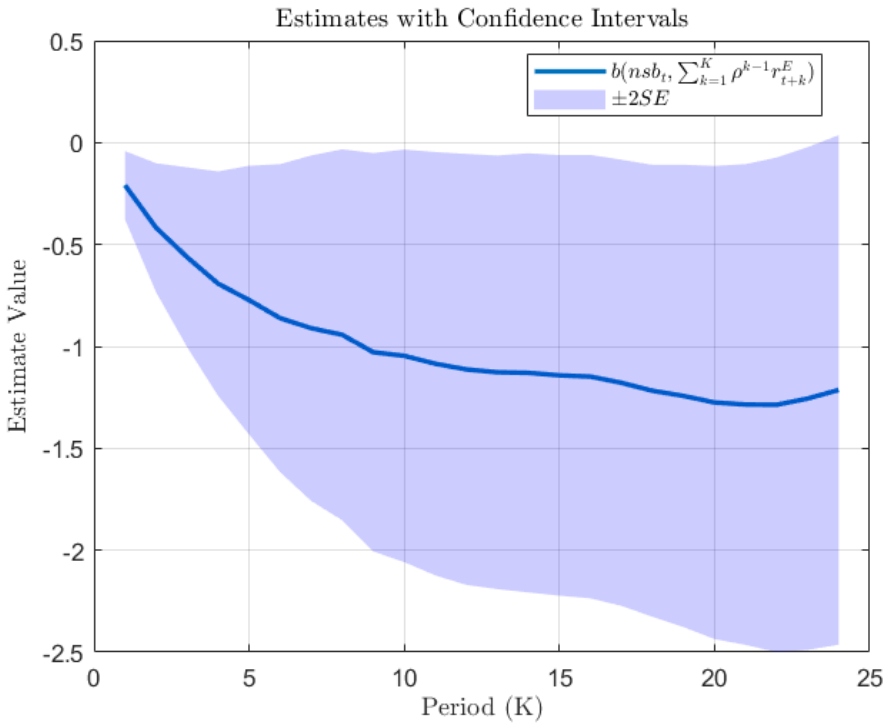


FIGURE 2: **Coefficient of long-horizon regression.** The figure plots the slope from regressing cumulative future equity returns, $\sum_{k=1}^{K} (\rho^E)^{k-1} r^E_{t+k}$, on nsb_t as K varies. Shaded areas show ±2 standard-deviation bands.

Output distortion is necessary to match the magnitude of $b(nsb_t, r^E_{t+1})$ estimated from the data.

Without distortion (i.e., $c^\tau = c^g = 0$ and $\alpha = 0$),

$$b(nsb_t, r^E_{t+1}) = -\frac{1-\beta}{1-\rho^K}\frac{T}{D}.$$

In that case, fiscal shocks affect dividends only through the ABC channel. Using our steady-state calibration,

$$\beta \approx \rho \approx 0.999, \qquad \frac{T}{D} \approx 0.2 \quad \Rightarrow \quad b(nsb_t, r^E_{t+1}) \approx -\frac{1}{K} \times 0.2 \approx -0.01,$$

whereas our empirical estimate $b(nsb_t, r^E_{t+1}) = -0.205$ implies a substantial role for output distortion beyond the ABC mechanism.

This corner case also echoes Ricardian Equivalence. If fiscal policy does not distort output ($c^\tau = c^g = 0$, $\alpha = 0$), the government would optimally forgo spending in Eq. (22) (since it does not raise output here), be indifferent among tax paths that satisfy the budget constraint, and consumption–wealth dynamics would depend only on the exogenous evolution of raw output. Eliminating fiscal policy's impact on output therefore restores Ricardian Equivalence: the government can "tax now" or "tax later" without changing equilibrium outcomes.

Overall, our model accounts for the strongly negative $b(nsb_t, r^E_{t+1})$ through fiscal policy's output effects governed by (c^τ, c^g). Since $b(nsb_t, r^E_{t+1})$ is significantly negative in the data, we can rule out the no-distortion corner solution and study the explicit mapping between $b(nsb_t, r^E_{t+1})$ and (c^τ, c^g).

The comparative-static approach is general and accommodates a wide class of steady-state benchmarks while isolating the fiscal channel through output distortion. This matches our empirical strategy: using state variables $\boldsymbol{x_t}$ (including the within-sector predictor dp) to capture the pre-shock equity risk premium,

$$r^E_{t+1} = \boldsymbol{b_r^E}\, \boldsymbol{x_t}, \quad \Rightarrow \quad \mathbb{E}_t \widetilde{r}^E_{t+1} = \boldsymbol{b_r^E}\, \boldsymbol{x_t} + b(r^E_{t+1}, nsb_t)\, \widetilde{nsb}_t.$$

As in the empirical tests, controlling for benchmark predictors isolates the incremental role of nsb; in the model, the steady-state benchmark isolates the fiscal channel. This alignment allows us to interpret $b(nsb_t, r^E_{t+1})$ structurally and infer (c^τ, c^g) from financial-market data.

We detail how to back out c^τ and c^g from $b(nsb_t, r^E_{t+1})$ in the next subsection. Finally, while we focus on a two-sector setting, the framework can accommodate the three-sector ABC in Eq. (18) by introducing an exogenous import shock u^X_{t+1}:

$$\frac{NM_{t+1}}{Y^r_{t+1}} = \frac{NM}{Y^r} e^{u^X_{t+1}}.$$

Consumption under the ABC becomes

$$C_t = Y^r_t \left(1 - \theta(TY^r_t, GY^r_t)\right) - G_t + NM_t.$$

Import shocks can then affect nma_t and sectoral risk premia. Since our empirical evidence finds no predictive power of nma_t for returns in other sectors, we omit this extension.

4.3 Distortion Implied from Return Predictability

The significantly negative estimate of $b(nsb_t, r^E_{t+1})$ implies output distortion and a violation of Ricardian Equivalence. This raises a central question: how strongly does fiscal policy affect finalized output? Our parametric specification,

$$\theta(TY^r, GY^r) = \frac{c^\tau}{2}(TY^r)^2 - \frac{c^g}{2}(GY^r)^2,$$

links $b(nsb_t, r^E_{t+1})$ directly to (c^τ, c^g), allowing financial markets' forward-looking information to discipline fiscal distortion magnitudes.

We use the GMM estimate $b(nsb_t, r^E_{t+1}) = -0.205$ from Panel B of Table 4 to calibrate (c^τ, c^g) via Eq. (26). This slope alone does not identify both parameters because raw output Y^r_t is unobservable (only finalized GDP Y_t is observed). We therefore write

$$TY^r = \frac{T}{Y}\frac{Y}{Y^r}, \quad GY^r = \frac{G}{Y}\frac{Y}{Y^r}, \tag{29}$$

and use Eq. (21) together with additional steady-state restrictions from the equilibrium model to jointly solve for c^τ, c^g, and Y/Y^r. Details are in Appendix; we summarize the solution below.

Proposition 2. *Given the predictive coefficient $b(nsb_t, r^E_{t+1})$, the output distortion parameters are solved as follows:*

- *The net distortion in output is*

$$\frac{Y}{Y^r} = \frac{-Br + \sqrt{Br^2 + 2Ar}}{Ar}, \tag{30}$$

  where Ar and Br are functions of $b(nsb_t, r^E_{t+1})$, the tax–spending comovement coefficient α, and steady-state values.

- *The distortion coefficients in Eq. (21) are*

$$c^\tau = -\frac{1}{1+\alpha} - b(nsb_t, r^E_{t+1})\frac{(1-\rho^K)(1-\alpha\beta)}{(1-\beta)(1+\alpha)}\frac{D}{T} + \frac{\alpha}{1+\alpha}\left(\frac{T}{Y}\frac{Y}{Y^r}\right)^{-1}, \tag{31}$$

$$c^g = \left[\frac{1}{1+\alpha}\left(\frac{T}{Y}\frac{Y}{Y^r}\right)^{-1} + \frac{1}{1+\alpha} + b(nsb_t, r^E_{t+1})\frac{(1-\rho^K)(1-\alpha\beta)}{(1-\beta)(1+\alpha)}\frac{D}{T}\right]/\beta. \tag{32}$$

Here, the steady-state values $\frac{T}{D}$, $\frac{G}{D}$, $\frac{T}{Y}$, and $\frac{G}{Y}$ are constants, and $\beta = \frac{G}{T} < 1$ is the stable spending-to-tax ratio used in Eq. (9).

We implement this solution using $b(nsb_t, r^E_{t+1}) = -0.205$. Along with $K = 20$ from Figure 2, we calibrate α from the joint dynamics of spending and taxation by regressing $gb = \log(G/B)$ on $tb = \log(T/B)$, yielding $\alpha = 0.827$. Other steady-state ratios (e.g., T/Y and D/T) are set to long-run historical averages, while ρ and β are inferred from the public-sector identity Eq. (9).

This procedure maps an empirically observed $b(nsb_t, r^E_{t+1})$ into an implied pair (c^τ, c^g), making $b(nsb_t, r^E_{t+1})$ an indicator of how fiscal-policy expectations affect output and asset prices. We further interpret this mapping below.

4.3.1 Dividend Distortion

From Eq. (23) and the fiscal-shock specification Eq. (25), dividend growth responds as

$$\Delta\tilde{d}_{t+1} - \Delta d_{t+1} = \left(-\frac{T}{D} - c^\tau \frac{T}{D}\right) u^\tau_{t+1} + \left(c^g \frac{G}{D}\right) u^g_{t+1}. \tag{33}$$

Given $b(nsb_t, r^E_{t+1})$, equations Eqs. (31)–(32) pin down (c^τ, c^g) and thus the loadings of dividend-growth changes on $(u^\tau_{t+1}, u^g_{t+1})$, i.e., the implied dividend distortion from taxes and spending.

In Figure 3, we plot the tax and spending contributions in Eq. (33), normalizing $u^\tau = 1$ and setting $u^g = \alpha$ as in Eq. (25). When the predictive coefficient is zero, the two effects offset. By contrast, our negative GMM benchmark (dashed line) indicates that tax distortions dominate spending stimulus, implying a sharp drag on dividend growth.

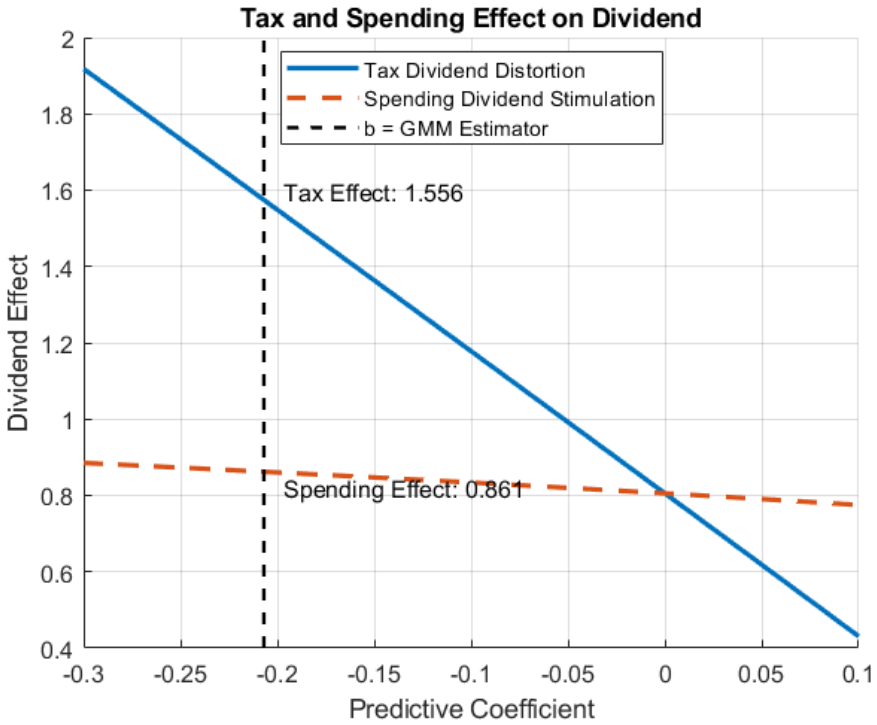


FIGURE 3: **Dividend distortion and predictive coefficient.** The figure shows how the tax and spending effects on dividend growth in Eq. (33) vary with $b(nsb_t, r^E_{t+1})$ in Eq. (26). The tax effect is plotted with the opposite sign for comparability. The y-axis reports the change in dividend growth for $u^\tau = 1$ and $u^g = \alpha$. Calibration uses $K = 20$ and $\alpha = 0.827$; the benchmark $b(nsb_t, r^E_{t+1}) = -0.205$ (Panel B, Table 4) is marked by the dashed line.

4.3.2 Output Distortion

The same mapping yields the implied tax distortion and spending stimulus in output:

$$\theta_\tau(TY^r) = \frac{1}{2}c^\tau\left(\frac{T}{Y}\frac{Y}{Y^r}\right)^2, \quad \theta_g(GY^r) = \frac{1}{2}c^g\left(\frac{G}{Y}\frac{Y}{Y^r}\right)^2.$$

Under our baseline calibration ($K = 20$, $\alpha = 0.827$),

$$\theta_\tau(TY^r) = 0.063, \quad \theta_g(GY^r) = 0.024 \quad \Rightarrow \quad \theta(TY^r, GY^r) = 0.039, \quad \frac{Y}{Y^r} = 0.961.$$

Thus taxation reduces output by about 6.3%, partially offset by a spending-induced gain of about 2.4%; the negative $b(nsb_t, r^E_{t+1})$ reflects tax distortions dominating spending stimulus.

In Figure 4, we plot $\theta_\tau(TY^r)$ and $\theta_g(GY^r)$ across values of $b(nsb_t, r^E_{t+1})$. Consistent with Eq. (26), a more negative slope corresponds to stronger tax distortion and weaker spending stimulus (and vice versa).

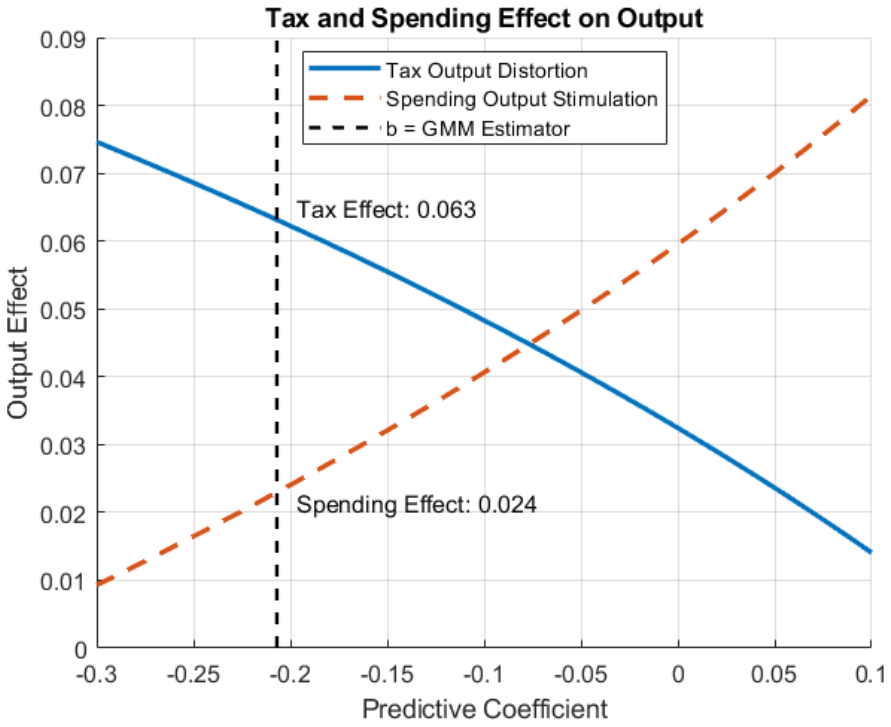


FIGURE 4: **Output distortion and predictive coefficient.** The figure plots the implied $\theta_\tau(TY^r)$ and $\theta_g(GY^r)$ from Eq. (21) as functions of $b(nsb_t, r^E_{t+1})$ in Eq. (26). The tax effect is plotted with the opposite sign for comparability. Calibration uses $K = 20$ and $\alpha = 0.827$; the benchmark $b(nsb_t, r^E_{t+1}) = -0.205$ (Panel B, Table 4) is marked by the dashed line.

4.3.3 Remarks

The coefficients (c^τ, c^g) are closely related to the literature on tax and spending multipliers. Analogous to Eq. (33), output growth responds to Eq. (25) as

$$\Delta\tilde{y}_{t+1} - \Delta y_{t+1} = \left(-c^\tau \frac{T}{Y}\right) u^\tau_{t+1} + \left(c^g \frac{G}{Y}\right) u^g_{t+1}. \tag{34}$$

Since

$$\Delta\tilde{y}_{t+1} - \Delta y_{t+1} \approx \left(\tilde{Y}_{t+1} - Y_{t+1}\right)/Y_t,$$

the implied responses in Eq. (34) translate into fiscal multipliers—how one dollar of taxes/spending changes finalized output. We defer additional discussion of multiplier measurement to Appendix, as it parallels a broad literature beyond this paper and is a natural topic for follow-up work.

Finally, the framework can accommodate channels beyond output distortion through which fiscal policy may affect equity markets (e.g., default risk or inflation dynamics, allowing fiscal decisions in real terms). While we leave such extensions to future research, our empirical analysis partially addresses them by using real returns and growth rates and controlling for debt sustainability via the tax-to-GDP ratio in estimation.

5 Conclusion

In this study, we investigate the joint dynamics and predictability of asset returns for the equity, treasury, and foreign asset investment sectors, as each represents a distinct asset class with different dynamics and potential investors. We find that the predictability of each sector's own returns remains largely intact even after accounting for state variables in other sectors.

To better understand the joint dynamics of the three sectors, we impose the aggregate budget constraint and estimate all predictive regressions jointly. The results of this estimation highlight the importance of considering the aggregate cyclical movement of the economy as a whole when examining the joint dynamics of asset returns. Based on our multivariate analysis, we find that the net surplus-to-debt ratio negatively predicts the risk premium in the equity and foreign asset investment sectors, as a lower government surplus could signal a contractionary fiscal policy in the future and higher risk premiums for investing in equity or foreign assets. However, this mechanism may be difficult to observe in the short term based solely on public sector data. Our results suggest that considering data from all three sectors and imposing aggregate budget constraints can help to better identify how this fiscal policy adjustment channel propagates throughout the economy.

Overall, our study contributes to the literature on asset return predictability by examining the joint dynamics of asset returns across different markets, considering their respective valuation ratios and cash flows, and incorporating budget constraints to better understand their interactions. Further research could explore the conditional correlations and regime-switching of the joint dynamics of the three sectors, as well as the implications of our findings for portfolio diversification and asset allocation.

We propose a theoretical framework to explain our finding that fiscal surplus negatively affects the equity risk premium. This framework enables us to leverage asset market data as a forward-looking estimator of the impact of tax distortions and government spending on output. Additionally, our approach is flexible enough to incorporate alternative channels beyond output distortion through which fiscal policy can influence equity markets, such as inflation risk and public debt default risk. We leave these extensions for future research.

Appendix A Derivations in the Theoretical Model

For tractability, assume i.i.d. growth in raw output as in Jiang et al. (2024a):

$$\frac{Y^r_{t+1}}{Y^r_t} = \exp\left(g - \frac{\sigma^2}{2} + \sigma\varepsilon_{t+1}\right). \tag{A1}$$

A1 Government's Optimal Tax and Spending

The government takes $\{Y^r_t\}$ and the pricing kernel as given and chooses tax and spending shares to maximize the representative household's indirect utility $F_0 = \max_{\{C_t\}} \mathbb{E}_0\Big[\sum_{t=0}^{\infty} e^{-\beta t} U(C_t)\Big]$. In this setup, maximizing F_0 is equivalent to maximizing household wealth,

$$W_0 = \mathbb{E}_0\Big[\sum_{s=1}^{\infty} M_s C_s\Big] = \mathbb{E}_0\Big[\sum_{s=1}^{\infty} M_s\big(Y^r_s - Y^r_s\theta(TY^r_s, GY^r_s) - Y^r_s GY^r_s\big)\Big], \tag{A2}$$

where $T_t = TY^r_t \times Y^r_t$ and $G_t = GY^r_t \times Y^r_t$.

The quadratic distortion function in Eq. eq:quadratic is $\theta_\tau(TY^r) = \frac{1}{2}c^\tau\left(\frac{T}{Y^r}\right)^2$, $\theta_g(GY^r) = \frac{1}{2}c^g\left(\frac{G}{Y^r}\right)^2$, so that $\theta(TY^r, GY^r) = \frac{c^\tau}{2}(TY^r)^2 - \frac{c^g}{2}(GY^r)^2$, and

$$\frac{\partial\theta}{\partial TY^r} = c^\tau TY^r, \qquad \frac{\partial\theta}{\partial GY^r} = -c^g GY^r. \tag{A3}$$

As shown below, $-c^\tau$ and c^g map into the output effects (multipliers) of taxes and spending.

The government chooses $\{TY^r_{t+1}, GY^r_{t+1}\}$ subject to the intertemporal budget constraint

$$B_0 \le \mathbb{E}_0\Big[\sum_{s=1}^{\infty} \frac{M_s}{M_0}(T_s - G_s)\Big], \tag{A4}$$

with debt dynamics

$$B_{t+1} = B_t(1 + R^B_{t+1}) - T_{t+1} + G_{t+1}.$$

The government treats debt valuation as exogenous (hence R^B_{t+1} adjusts with the value of debt).

We assume the government maximizes Eq. (A2) augmented by a quadratic "private benefit" term, $\frac{s^\tau}{2}Y^r_{t+1}(TY^r_{t+1})^2$, which captures non-benevolent motives (e.g., private benefits/credibility; Barro, 1973). Relative to a standard tax-smoothing setup with exogenous spending, this term helps match empirical magnitudes of tax/spending effects on output and induces a joint determination of optimal taxes and spending.

The Bellman equation is

$$P(B_t, Y^r_t) = \max_{TY^r_{t+1}, GY^r_{t+1}} \mathbb{E}_t\left[m_{t+1}\left(Y^r_{t+1} - Y^r_{t+1}\theta(TY^r_{t+1}, GY^r_{t+1}) - Y^r_{t+1}GY^r_{t+1} + \frac{s^\tau}{2}Y^r_{t+1}(TY^r_{t+1})^2 + P(B_{t+1}, Y^r_{t+1})\right)\right]. \tag{A5}$$

With i.i.d. Y^r we can use guess-and-verify to obtain a steady state with constant TY^r and GY^r and a stable debt-to-output ratio. The optimized value function is homogeneous: $P^*(B, Y^r) =$

$p^*(B/Y^r)\,Y^r$, implying a constant marginal value of debt, $P_B^* = \frac{\partial P^*}{\partial B} = p'(B/Y^r)$. Optimal fiscal shares satisfy the (steady-state) first-order conditions

$$-1 - \frac{\partial \theta}{\partial GY^r} = -p'(B/Y^r), \qquad -\frac{\partial \theta}{\partial TY^r} + s^\tau TY^r = p'(B/Y^r).$$

As a parsimonious baseline we set $s^\tau = 2c^\tau$ and focus on calibrating c^τ and c^g (the fiscal multipliers). The main prediction—negative predictability of the surplus-to-debt ratio for equity returns—is not sensitive to s^τ, and this baseline yields empirically reasonable multipliers (Blanchard and Perotti, 2002; Romer and Romer, 2010; Christiano et al., 2011; Auerbach and Gorodnichenko, 2012; Ilzetzki et al., 2013; Ramey and Zubairy, 2018). Using Eq. (A3), the FOCs imply

$$c^\tau TY^r + c^g GY^r = 1. \tag{A6}$$

In addition, under i.i.d. Y^r, the steady-state surplus-to-output ratio is pinned down by initial debt: $B_0/Y_0^r \propto TY^r - GY^r$. Together, these conditions determine the optimal fiscal policy.

When output distortions and private benefits are shut down ($s^\tau = c^\tau = c^g = 0$), the government optimally sets $G = 0$ (since spending does not raise output here) and fiscal policy becomes neutral for consumption-wealth dynamics: $W_0 = \mathbb{E}_0\Big[\sum_{s=1}^\infty M_s C_s\Big] = \mathbb{E}_0\Big[\sum_{s=1}^\infty M_s Y_s^r\Big]$, $P^*(B, Y^r) \propto Y^r$. The government is then indifferent among all tax/spending paths satisfying Eq. eq:Debtcon, restoring Ricardian equivalence.

A2 Asset Valuation

Fiscal policy determines household consumption and after-tax dividends:

$$C_t = Y_t - G_t = Y_t^r\big(1 - \theta(TY_t^r, GY_t^r)\big) - G_t = Y_t^r - Y_t^r\theta(TY_t^r, GY_t^r) - Y_t^r GY_t^r, \tag{A7}$$

$$D_t = Y_t - T_t = Y_t^r\big(1 - \theta(TY_t^r, GY_t^r)\big) - T_t = Y_t^r - Y_t^r\theta(TY_t^r, GY_t^r) - Y_t^r TY_t^r. \tag{A8}$$

With constant TY^r and GY^r, all cash flows are proportional to Y_t^r and share the same growth rate: $\frac{NS_{t+1}}{NS_t} = \frac{D_{t+1}}{D_t} = \frac{C_{t+1}}{C_t} = \frac{Y_{t+1}^r}{Y_t^r} = \exp\Big(g - \frac{\sigma^2}{2} + \sigma\varepsilon_{t+1}\Big)$. The risk-free rate under the CRRA pricing kernel is

$$r = s - \log \mathbb{E}_t\left[\left(\frac{C_{t+1}}{C_t}\right)^{-\gamma}\right] = s + \gamma g - \frac{\gamma(\gamma+1)}{2}\sigma^2. \tag{A9}$$

Dividends can be written as $D_t = Y_t^r(1 - \theta - TY^r)$. Define $\lambda = \gamma\sigma^2$ and $\delta = r + \lambda - g$. Then

$$E_t\left[\frac{M_{t+1}}{M_t}Y_{t+1}^r\right] = Y_t^r e^{-\delta}, \qquad E_t\left[\frac{M_{t+k}}{M_t}Y_{t+k}^r\right] = Y_t^r e^{-k\delta}, \ \forall k. \tag{A10}$$

Equity value and return are therefore

$$A_t = \mathbb{E}_t\left[\sum_{u=t+1}^{\infty}\frac{M_u}{M_t}(1 - \theta - TY^r)\,Y_u^r\right] = \frac{e^{-\delta}}{1 - e^{-\delta}}\,D_t, \tag{A11}$$

$$R^E_{t+1} = \frac{A_{t+1} + D_{t+1}}{A_t} = \exp\left(r + \lambda - \frac{\sigma^2}{2} + \sigma\varepsilon_{t+1}\right), \qquad \mathbb{E}_t\left[R^E_{t+1}\right] = e^{r+\lambda}, \tag{A12}$$

with risk premium $\lambda = \gamma\sigma^2$. Under i.i.d. growth, $dp_t = \delta = r+\lambda-g$; $\rho^E = e^{-\delta}$; $\mathbb{E}_t r^E_{t+1} = r+\lambda$; $\mathbb{E}_t \Delta d_{t+1} = g$. The log-linearized budget constraint (as used in the main text) reduces to

$$(1-\rho^E)\mathbb{E}_t r^E_{t+1} - (1-\rho^E)\mathbb{E}_t \Delta d_{t+1} = (1-e^{-\delta})\delta = dp_t - \rho^E \mathbb{E}_t dp_{t+1}.$$

In this simplified setting, the equity and public debt shares a common discount factor, $\rho = \rho^E = \rho^B = e^{-\delta}$. Similarly, $\frac{NS_t}{B_t} = \frac{1-e^{-\delta}}{e^{-\delta}} = \frac{1-\rho}{\rho}$. The public sector also satisfies $(1-\rho^B)\mathbb{E}_t r^B_{t+1} - (1-\rho^B)\mathbb{E}_t \Delta ns_{t+1} = (1-e^{-\delta})\delta = nsb_t - \rho^B \mathbb{E}_t nsb_{t+1}$.

A3 Response of Cash Flows to Fiscal Shocks

Given static within-sector risk-return relations, we adopt a comparative-static exercise (as in Barro, 1979). Consider a small, temporary deviation in expected fiscal shares:

$$T\tilde{Y}^r_{t+1} = TY^r_{t+1}e^{u_{t+1}}, \qquad G\tilde{Y}^r_{t+1} = GY^r_{t+1}e^{u^g_{t+1}}, \qquad u^g_{t+1} = \alpha u_{t+1}. \tag{A13}$$

Then $\Delta\tilde{t}_{t+1} - \Delta t_{t+1} = u_{t+1}$, $\Delta\tilde{g}_{t+1} - \Delta g_{t+1} = \alpha u_{t+1}$. For any variable x, define its response coefficient

$$U_x = \frac{\tilde{x}_{t+1} - x_{t+1}}{u_{t+1}}. \tag{A14}$$

Using the public-sector log-linearization, $\frac{TY^r}{TY^r - GY^r} = \frac{T}{T-G} = \frac{1}{1-\beta}$, $\Delta ns_{t+1} = \frac{1}{1-\beta}\Delta t_{t+1} - \frac{\beta}{1-\beta}\Delta g_{t+1}$, we obtain $\Delta\tilde{ns}_{t+1} - \Delta ns_{t+1} = U_{\Delta ns}u_{t+1} = \frac{1-\alpha\beta}{1-\beta}\,u_{t+1}$. For dividends, a first-order expansion of Eq. (A8) implies $\tilde{D}_{t+1} = D_{t+1} - \frac{\partial\theta}{\partial TY^r}(e^{u_{t+1}}-1)Y^r_{t+1} - \frac{\partial\theta}{\partial GY^r}(e^{\alpha u_{t+1}}-1)Y^r_{t+1} - (e^{u_{t+1}}-1)TY^rY^r_{t+1}$. Using $(e^\epsilon - 1) \approx \epsilon$ and $1-\epsilon \approx e^{-\epsilon}$ for small ϵ gives

$$\tilde{D}_{t+1} \approx D_{t+1}\exp\left[\left(-\frac{T}{D} - \frac{\partial\theta}{\partial TY^r}\frac{Y^r}{D} - \alpha\frac{\partial\theta}{\partial GY^r}\frac{Y^r}{D}\right)u_{t+1}\right].$$

With Eq. (A3), the dividend-growth response is

$$\Delta\tilde{d}_{t+1} - \Delta d_{t+1} = U_{\Delta d}u_{t+1} = \left(-\frac{T}{D} - c^\tau\frac{T}{D} + \alpha c^g\frac{G}{D}\right)u_{t+1}. \tag{A15}$$

Analogously, from Eq. (A7) the consumption-growth response is

$$\Delta\tilde{c}_{t+1} - \Delta c_{t+1} = U_{\Delta c}u_{t+1} = \left(-\alpha\frac{G}{C} - c^\tau\frac{T}{C} + \alpha c^g\frac{G}{C}\right)u_{t+1}. \tag{A16}$$

A4 Response of Prices to Fiscal Shocks and Return Predictability

Let u_{t+1} have mean μ_u and variance σ_u^2. Since $r = s - \log \mathbb{E}_t\left[\left(\frac{C_{t+1}}{C_t}\right)^{-\gamma}\right]$, the shock-induced change in the risk-free rate is

$$\tilde{r} - r = \gamma U_{\Delta c}\mu_u + \frac{1}{2}(\gamma U_{\Delta c})^2\sigma_u^2.$$

Now suppose the fiscal shock persists for K periods, with $\{u_{t+k}\}_{k=1}^K$ i.i.d. The shock affects the bond value through the first K cash flows:

$$\tilde{B}_t = B_t + \mathbb{E}_t \sum_{k=1}^{K}\left[\frac{M_{t+k}}{M_t} NS_{t+k}\left(e^{(-\gamma U_{\Delta c}+U_{\Delta ns})u_{t+k}} - 1\right)\right].$$

Using Eq. (A10) and the argument that NS/B is a constatn at steady state, this simplifies to

$$\tilde{B}_t = B_t\left(1 + (1-\rho^K)\,\mathbb{E}_t\left[e^{(-\gamma U_{\Delta c}+U_{\Delta ns})u_{t+1}} - 1\right]\right) \approx B_t\left(1 + (1-\rho^K)\log \mathbb{E}_t\left[e^{(-\gamma U_{\Delta c}+U_{\Delta ns})u_{t+1}}\right]\right),$$

hence

$$\tilde{b}_t = b_t + (1-\rho^K)\log \mathbb{E}_t\left[e^{(-\gamma U_{\Delta c}+U_{\Delta ns})u_{t+1}}\right].$$

To obtain return responses, use the (log-linearized) budget constraints:

$$\mathbb{E}_t r_{t+1}^B = \mathbb{E}_t \Delta ns_{t+1} + \frac{1}{1-\rho} nsb_t - \frac{\rho}{1-\rho}\mathbb{E}_t nsb_{t+1}, \qquad \mathbb{E}_t r_{t+1}^E = \mathbb{E}_t \Delta d_{t+1} + \frac{1}{1-\rho} da_t - \frac{\rho}{1-\rho}\mathbb{E}_t da_{t+1}.$$

First, $\mathbb{E}_t\left(\Delta\tilde{ns}_{t+1} - \Delta ns_{t+1}\right) = \log \mathbb{E}_t\left[e^{U_{\Delta ns}u_{t+1}}\right] = U_{\Delta ns}\mu_u + \frac{1}{2}(U_{\Delta ns})^2\sigma_u^2$. With $nsb_t \approx \kappa + (1-\rho)(\log ns_t - \log b_t)$, we have

$$\frac{1}{1-\rho}(\tilde{nsb}_t - nsb_t) = -(\tilde{b}_t - b_t) = -(1-\rho^K)\log \mathbb{E}_t\left[e^{(-\gamma U_{\Delta c}+U_{\Delta ns})u_{t+1}}\right],$$

and similarly,

$$\frac{\rho}{1-\rho}\mathbb{E}_t(\tilde{nsb}_{t+1} - nsb_{t+1}) = \rho\,\mathbb{E}_t(\Delta\tilde{ns}_{t+1} - \Delta ns_{t+1}) - (\rho - \rho^K)\log \mathbb{E}_t\left[e^{(-\gamma U_{\Delta c}+U_{\Delta ns})u_{t+1}}\right].$$

Accounting for the remaining $K-1$ periods of shocks implies

$$\mathbb{E}_t(\tilde{nsb}_{t+1} - nsb_{t+1}) = (1-\rho)\mathbb{E}_t(\Delta\tilde{ns}_{t+1} - \Delta ns_{t+1}) + \frac{1-\rho^{K-1}}{1-\rho^K}(\tilde{nsb}_t - nsb_t).$$

Reorganizing yields the expected return responses:

$$\mathbb{E}_t(\tilde{r}_{t+1}^B - r_{t+1}^B) = (1-\rho)(\tilde{r}-r) + (1-\rho)\gamma U_{\Delta c}U_{\Delta ns}\sigma_u^2, \qquad \mathbb{E}_t(\tilde{r}_{t+1}^E - r_{t+1}^E) = (1-\rho)(\tilde{r}-r) + (1-\rho)\gamma U_{\Delta c}U_{\Delta d}\sigma_u^2.$$

To focus on the risk-premium channel (consistent with evidence that expected returns are more

predictable than cash-flow growth), we normalize the cash-flow mean effect to zero:

$$\mathbb{E}_t(\Delta\tilde{ns}_{t+1} - \Delta ns_{t+1}) = U_{\Delta ns}\mu_u + \frac{1}{2}(U_{\Delta ns})^2\sigma_u^2 = 0 \quad \Rightarrow \quad \mu_u = -\frac{1}{2}U_{\Delta ns}\sigma_u^2.$$

Under this normalization, the induced variation in $\tilde{r}-r$ is small (and equals zero when $\gamma U_{\Delta c} = U_{\Delta ns}$). Retaining only the risk-premium channel gives

$$\mathbb{E}_t(\tilde{r}^E_{t+1} - r^E_{t+1}) = (1-\rho)\gamma U_{\Delta c}U_{\Delta d}\sigma_u^2, \qquad \tilde{nsb}_t - nsb_t = (1-\rho)(1-\rho^K)\gamma U_{\Delta c}U_{\Delta ns}\sigma_u^2.$$

Hence the single-period predictive coefficient is

$$b(nsb_t, r^E_{t+1}) = \frac{\mathbb{E}_t(\tilde{r}^E_{t+1} - r^E_{t+1})}{\tilde{nsb}_t - nsb_t} = \frac{1}{1-\rho^K}\cdot\frac{U_{\Delta d}}{U_{\Delta ns}}.$$

Since shocks are i.i.d. over the K-period episode, the long-horizon change in expected returns is

$$\mathbb{E}_t\left(\sum_{k=1}^{K}\rho^{k-1}\tilde{r}^E_{t+k} - \sum_{k=1}^{K}\rho^{k-1}r^E_{t+k}\right) = (1-\rho^K)\gamma U_{\Delta c}U_{\Delta d}\sigma_u^2,$$

so the long-horizon predictive coefficient equals $\frac{1-\rho^K}{1-\rho}\cdot b(nsb_t, r^E_{t+1})$. Finally, under a $(K-1)$-period shock episode starting at $t+1$,

$$\tilde{nsb}_{t+1} - nsb_{t+1} = (1-\rho)(1-\rho^{K-1})\gamma U_{\Delta c}U_{\Delta ns}\sigma_u^2 \quad \Rightarrow \quad \mathbb{E}_t(\tilde{nsb}_{t+1} - nsb_{t+1}) = \frac{1-\rho^{K-1}}{1-\rho^K}(\tilde{nsb}_t - nsb_t).$$

These expressions complete the proof of the return-predictability proposition in the main text. As a useful benchmark, when $c^\tau = c^g = 0$ and $\alpha = 0$, we have $U_{\Delta c} = 0$ and fiscal shocks do not affect risk pricing (Ricardian equivalence).

A5 Calibration of Output Distortion

We compute steady-state ratios using sample averages and set $\beta = \rho = 0.999$ (per the public-sector present-value identity). Since $c^\tau TY^r + c^g GY^r = (c^\tau + c^g\beta)TY^r = 1$, combining this restriction with Eq. eq:bcalibrate yields

$$c^\tau = -\frac{1}{1+\alpha} - b(nsb_t, r^E_{t+1})\frac{(1-\rho^K)(1-\alpha\beta)}{(1-\beta)(1+\alpha)}\frac{D}{T} + \frac{\alpha}{1+\alpha}(TY^r)^{-1}, \tag{A17}$$

$$c^g = \left(\frac{1}{1+\alpha}(TY^r)^{-1} + \frac{1}{1+\alpha} + b(nsb_t, r^E_{t+1})\frac{(1-\rho^K)(1-\alpha\beta)}{(1-\beta)(1+\alpha)}\frac{D}{T}\right)\Big/\beta. \tag{A18}$$

Since T/Y is observable but Y^r is not, write $TY^r = \frac{T}{Y}\cdot\frac{Y}{Y^r}$, and solve for Y/Y^r from the net-distortion definition:

$$\frac{Y}{Y^r} = 1 - \frac{1}{2}c^\tau\left(\frac{T}{Y^r}\right)^2 + \frac{1}{2}c^g\left(\frac{G}{Y^r}\right)^2 = 1 - \frac{1}{2}c^\tau\left(\frac{T}{Y}\cdot\frac{Y}{Y^r}\right)^2 + \frac{1}{2}c^g\left(\frac{G}{Y}\cdot\frac{Y}{Y^r}\right)^2,$$

which implies the quadratic equation

$$\frac{1}{2}\left(c^{\tau}\left(\frac{T}{Y}\right)^{2}-c^{g}\left(\frac{G}{Y}\right)^{2}\right)\left(\frac{Y}{Y^{r}}\right)^{2}+\frac{Y}{Y^{r}}-1=0. \tag{A19}$$

Substituting Eqs. (A17)–(A18) into Eq. (A19) yields a quadratic in Y/Y^r. Define

$$Ar=(1+\beta)\left[-\frac{1}{1+\alpha}-b(nsb_t,r^{E}_{t+1})\frac{(1-\rho^{K})(1-\alpha\beta)}{(1-\beta)(1+\alpha)}\frac{D}{T}\right]\left(\frac{T}{Y}\right)^{2},\qquad Br=\frac{1}{2}\cdot\frac{\alpha-\beta}{1+\alpha}\cdot\frac{T}{Y}+1.$$

Then

$$\frac{Y}{Y^{r}}=\frac{-Br+\sqrt{Br^{2}+2Ar}}{Ar}. \tag{A20}$$

A6 Fiscal Multipliers

Output growth responds to fiscal policy shocks through the same distortion channel as dividends and consumption:

$$\Delta\tilde{y}_{t+1}-\Delta y_{t+1}=\left(-c^{\tau}\frac{T}{Y}\right)u^{\tau}_{t+1}+\left(c^{g}\frac{G}{Y}\right)u^{g}_{t+1}, \tag{A21}$$

where tildes denote outcomes under the distorted fiscal path and $u^{\tau}_{t+1}, u^{g}_{t+1}$ are tax and spending shocks. In steady state at time t, a one-period tax shock implies

$$\frac{\mathrm{E}_t(\Delta\tilde{y}_{t+1}-\Delta y_{t+1})}{\mathrm{E}_t\left(\Delta\tilde{t}_{t+1}-\Delta t_{t+1}\right)}=-\frac{T}{Y}c^{\tau}.$$

Using the first-order approximations $\Delta\tilde{y}_{t+1}-\Delta y_{t+1}\approx\frac{\tilde{Y}_{t+1}-Y_{t+1}}{Y_t}$, $\Delta\tilde{t}_{t+1}-\Delta t_{t+1}\approx\frac{\tilde{T}_{t+1}-T_{t+1}}{T_t}$, and T_t/Y_t at its steady-state value, the tax multiplier is

$$\frac{\mathrm{E}_t\left(\tilde{Y}_{t+1}-Y_{t+1}\right)}{\mathrm{E}_t\left(\tilde{T}_{t+1}-T_{t+1}\right)}=-c^{\tau}. \tag{A22}$$

Analogously, for a one-period government spending shock,

$$\frac{\mathrm{E}_t\left(\tilde{Y}_{t+1}-Y_{t+1}\right)}{\mathrm{E}_t\left(\tilde{G}_{t+1}-G_{t+1}\right)}=c^{g}. \tag{A23}$$

Hence c^{τ} and c^{g} are forward-looking tax and spending multipliers in our framework: the expected dollar change in output induced by a one-dollar change in taxes or government spending. Combined with the mapping in Eqs. (A17)–(A18), they can be inferred from the predictive coefficient $b(nsb_t, r^{E}_{t+1})$.

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