Tension for the President: How Does Presidential Approval Rating Predict Growth and Asset Prices?

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Abstract

We study how presidential approval shapes U.S. growth and asset prices. Using Gallup presidential approval (PAR) data, we show that future output growth and the equity risk premium are highest when approval is in an intermediate range and lowest at the extremes—a hump-shaped relation in PAR. We summarize this pattern with approval tightness, which peaks when electoral competition is tight and approval is most sensitive to economic performance. High approval tightness forecasts stronger growth, higher expected returns, lower volatility, restrained federal spending, and improved governance. A political-economy model rationalizes these patterns: high approval tightness disciplines rent-seeking incumbents, reduces fiscal distortion and uncertainty, and boosts growth and asset valuations.

JEL classification: G12, E32, E44, E62, D72
Keywords: Presidential Approval, Political Support, Fiscal Policy, Political Uncertainty, Stock Return Predictability

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1 Introduction

A central theme in political economy is that elections discipline government behavior: incumbents internalize the economic consequences of policy because voters can reward or punish them at the ballot box. A large theoretical and empirical literature studies how electoral incentives constrain rent extraction and shape fiscal policy, public spending, and debt dynamics (e.g., Barro, 1973; Ferejohn, 1986; Yared, 2010; Nordhaus, 1975; Rogoff, 1990; Alesina and Tabellini, 1990; Acemoglu et al., 2008; Boyer et al., 2024), among many others. These ideas have a simple implication: political support is not just "politics"—it is a constraint that can shape real decisions with real economic consequences. We measure political support using the presidential approval rate (PAR) from the Gallup Poll—a long-running, nationally representative survey that produces arguably the longest continuous time series of public opinion in modern U.S. politics.[1]

We ask a simple but under-explored question: do movements in presidential approval show up in the economy and financial markets? The electoral-discipline logic suggests that the relation need not be monotonic. When approval is very high, the incumbent is less electorally constrained; when approval is very low, the marginal electoral return to better performance may be limited. By contrast, when approval is near a knife-edge, electoral constraints are tightest, sharpening incentives to pursue growth-supporting policies. Guided by this perspective, we ask two sets of questions. First, is PAR related to subsequent economic growth and asset returns—and does the relation differ between high-approval and low-approval periods? Second, do changes in approval coincide with changes in government behavior, especially fiscal choices, and with shifts in political uncertainty that are relevant for risk premia? We answer these questions and organize our main findings around three primary contributions.

Our first main finding is that presidential approval contains strong *nonlinear* forecasting information for both the macroeconomy and asset prices. Using data from 1969 to 2024, we show that future output growth and the equity risk premium are highest when approval is in an intermediate range and lowest at the extremes—a hump-shaped relation in PAR. For illustration, we simply

[1] Since the late 1930s, Gallup has repeatedly asked whether respondents approve or disapprove of the way the incumbent president is handling the job, yielding PAR. In political science, PAR is commonly interpreted as a proxy for presidential political capital and has been shown to co-move with economic performance and to track the president's ability to advance preferred policies (e.g., Neustadt, 1960; Mueller, 1970; Rivers and Rose, 1985; Canes-Wrone and De Marchi, 2002; Kriner and Schickler, 2014).

sum log excess returns on the CRSP value-weighted market portfolio over our sample—without compounding or assuming dividend reinvestment—and obtain a cumulative log excess return of 291.1%. Strikingly, 209.3%, about 72% of this total, are earned in quarters when lagged PAR lies between its 25th and 75th percentiles. A simple median-split regression exercise makes this hump shape more transparent. In full-sample linear regressions of next-quarter equity excess returns and real GDP (or PCE) growth on PAR, coefficients are small and statistically weak. Estimating the same one-variable regressions separately in low-approval and high-approval subsamples, however, yields economically large and sharply estimated slopes of opposite signs. For example, for next-quarter equity excess returns the slope is 0.433 when approval is below its sample median and −0.197 when approval is above, compared to an insignificant full-sample slope of 0.022. In other words, the economic effects of presidential support depend critically on the level of approval, and a simple linear specification in PAR is misspecified.

To capture this nonlinearity in a parsimonious way, we introduce a new political measure, presidential approval rate tightness,

$$\text{PART} \equiv \text{PAR}\big(1 - \text{PAR}\big).$$

By construction, PART is highest when approval is near 50%, where electoral competition is tightest, and declines toward zero as approval approaches either extreme, where political support is relatively secure or largely out of reach. Empirically, we find that the approval tightness enters predictive regressions with a sizable positive coefficient for real growth and equity excess returns; since the tightness measure is a concave function of approval rate, this positive coefficient implies a hump-shaped relation between political support and economic growth, as well as the equity premium.

Regressions of next-quarter real GDP and PCE growth on the level of approval alone yield small, imprecise coefficients and essentially no explanatory power, whereas replacing PAR with PART immediately reveals strong predictability: higher approval tightness forecasts faster real growth, and this pattern persists for cumulative four-quarter and one-year-ahead horizons. The information in approval tightness is not subsumed by approval levels or by standard macro expectations. In joint regressions that include both PAR and PART, the coefficient on PART remains large and

significant, and the joint specification increases explanatory power. Adding Survey of Professional Forecasters expectations leaves the coefficient on PART essentially unchanged, indicating that this political factor is not fully incorporated into these survey forecasts.

Turning to financial markets, the approval–tightness by itself forecasts next-period equity excess returns with adjusted R^2 of 1.3% at the monthly frequency and 3.14% at the quarterly frequency. The predictive power of approval tightnes is stable across monthly and quarterly data and persists at longer horiozns. Moreover, adding the dividend–price ratio of Campbell and Shiller (1988), the standard return predictors in Welch and Goyal (2008), and macroeconomic controls leaves the predictive coefficients largely unchanged, indicating that the approval–tightness embeds information about time variation in the equity risk premium that is not captured by valuation ratios or conventional macroeconomic factors.

Guided by the literature on political uncertainty and risk premia (e.g., Pastor and Veronesi (2012)), we also study relation between approval tightness and volatility. We find that PART negatively predicts multiple volatility and uncertainty measures, including the CBOE Volatility Index (VIX), constructed from S&P 500 index options; the Economic Policy Uncertainty (EPU) index of Baker et al. (2016); and the financial, macro, and real uncertainty measures of Jurado et al. (2015). Consistent with the hump-shaped patterns for growth and returns, volatility and uncertainty are lowest when approval tightness is highest and rise as presidential support moves toward either extreme.

Several further aspects of our empirical analysis are worth emphasizing. First, the predictive power of approval tightness survives conditioning on prominent proxies for public beliefs, including the economic-belief disagreement index of Huang et al. (2021) and the Economic Policy Uncertainty (EPU) index. Second, a battery of standard stationarity tests indicates that both PAR_t and PART_t can be treated as stationary time series over our sample. Third, the predictive relations we document remain when we correct for small-sample bias using the approach of Amihud and Hurvich (2004). Finally, and most importantly, our results survive standard out-of-sample tests: for example, a univariate PART_t regression delivers out-of-sample R^2 values of 2.02% and 5.06% when forecasting monthly and quarterly equity excess returns, respectively.

The fact that approval tightness predicts real growth, equity excess returns, and volatility indices points to an intriguing pattern: periods in which electoral competition is tight—when PAR is in

an intermediate range and PART is high—are followed by stronger growth, higher asset valuations, and lower contemporaneous volatility than either high- or low-approval regimes.

Our second contribution is to uncover the economic channel that makes intermediate-approval (high-PART) periods special. Central to our interpretation is that PART measures the *sensitivity* of presidential approval to economic performance.

As a first step, we estimate a logistic relation between the approval rate and a set of economic performance measures, controlling for the fixed effects of major political events. In this specification, real growth loads positively on approval, while inflation, the unemployment rate, and fiscal-condition variables such as tax revenues and expenditure shares of GDP enter with significantly negative coefficients. The logit link is both tractable and realistic: it explicitly respects that presidential approval is a probability-like object bounded between zero and one. These features help rationalize why linear models have struggled to fully capture approval dynamics in earlier work Berlemann and Enkelmann (2014); Gordon (2024).

Importantly, under the logit mapping the derivative of approval with respect to the economic fundamental is exactly $\mathrm{PAR}(1-\mathrm{PAR}) = \mathrm{PART}$, so the logit estimates imply that approval tightness coincides with the local responsiveness of political support to fundamentals. While this identity is specific to the logistic link, the underlying shape is much more general: whenever approval is generated by a unimodal probability mapping with moderate skewness, the marginal response of approval to fundamentals is single-peaked, with sensitivity highest in the intermediate range of PAR and attenuated near the extremes. We confirm this robustness with a probit specification, which delivers similar coefficients and an almost identical sensitivity profile.

Taken together, our results show that PART serves as an empirical measure of approval sensitivity. This perspective yields a structural interpretation of why tight-approval periods are special: when PART is high, electoral discipline is strongest and incumbents have the greatest incentive to choose policies that improve economic performance in order to raise their probability of remaining in office. Building on this interpretation, we show that approval tightness is closely tied to fiscal discipline and governance quality. Tighter approval conditions forecast significantly lower real growth of federal government spending, especially the non-defense purpose component, at both quarterly and annual horizons, with effects that are economically large, more pronounced at longer horizons, and robust to conditioning on inflation, lagged spending growth, and alternative measures

of fiscal conditions such as the surplus-to-debt ratio, tax revenues, and expenditure shares of GDP. At the same time, high PART is associated with better subsequent scores on perception-based measures of control of corruption and government capacity from the Worldwide Governance Indicators. In sum, periods of tight electoral competition are followed by restrained expenditure growth and stronger government capacity, consistent with political-economy models in which heightened electoral discipline curbs rent-seeking and improves public-sector performance.

Our third contribution is to develop a structural model that unifies our empirical findings, in which PART—as approval sensitivity—disciplines the government's rent-seeking behavior through inefficient fiscal spending. We build on a standard tax-smoothing framework à la Barro (1979), where a benevolent government collects taxes to finance exogenous public-good production, and extend it by allowing incumbents to allocate resources to unproductive expenditures that generate private rents. Rent extraction distorts final output and also raises political uncertainty: higher rents increase macroeconomic volatility and worsen the valuation of public debt, thereby raising financing costs. In equilibrium, rent seeking deteriorates economic fundamentals but is constrained by electoral incentives.

The key mechanism is that the responsiveness of approval to fundamentals tightens or relaxes electoral discipline across political states. When approval is intermediate (high PART), the marginal electoral return to improving fundamentals is high, so policymakers have strong incentives to restrain rents and tilt policy toward growth-enhancing choices. When approval is very high or very low (low PART), incentives weaken, rent-seeking expands, and political uncertainty rises. The solution of the model shows that optimal rent extraction is a convex, U-shaped function of the initial approval level, and political uncertainty inherits the same U-shape in PAR. Through the distortionary effects of rents on output and discount rates, this pattern translates into a hump-shaped, concave relation between approval and economic fundamentals, as well as private-sector profits and expected returns. The model thus provides a unified interpretation of why intermediate-approval periods are associated with stronger subsequent growth and higher expected returns, together with subdued spending growth and lower uncertainty.

Related Literature

Our work is closely related to the literature studying the relation between political approval and economic outcomes. Liu and Shaliastovich (2022) document that high presidential job approval ratings forecast a decline in the dollar risk premium and a reduction in future economic volatility. We focus on a different set of outcomes and a different state variable. We introduce approval tightness as a new predictor and show that it captures a state-dependent, hump-shaped relation between presidential approval and subsequent equity returns. Using an extended sample for real activity, we also recover the positive relation between growth and the level of approval, while showing that approval tightness is a more powerful predictor both statistically and in economic magnitude.

Our interpretation of presidential approval as a forward-looking signal of policy direction is related to event-study evidence on political news and markets. In our framework, when approval tightness is high, incumbents face stronger electoral discipline, reduce policy-induced uncertainty, and raise the near-term valuation of risky assets, so high approval tightness predicts higher equity returns. Liu and Shaliastovich (2021) provide complementary evidence by documenting positive abnormal stock market returns around major U.S. political announcement events, including presidential and midterm elections and State of the Union addresses.

Our paper also contributes to the broader literature on how elections and partisan politics affect financial markets and firm decisions. Santa-Clara and Valkanov (2003) document the "presidential puzzle": average excess market returns are higher under Democratic than under Republican presidents. Pástor and Veronesi (2020) provide a theoretical explanation, focusing on political cycles driven by the party controlling the government. Belo and Yu (2013) show that public infrastructure investment is an important amenity in industrial production and that higher government investment rates positively forecast production growth and stock returns. Belo et al. (2013) measure industry-level exposure to government spending and document that industries with high exposure earn higher risk premia during Democratic administrations.

We also complement work that links presidential approval to the cross-section of stock returns. Chen et al. (2023) construct a presidential economic approval rating (PEAR) index using multiple national polls on the president's handling of the economy and show that stocks with high exposure

to the PEAR factor earn lower average returns. While their focus is on cross-sectional pricing of approval risk, we emphasize the time-series, nonlinear relation between approval, growth, and aggregate risk premia.

Finally, our measure of approval tightness is related to, but distinct from, disagreement-based predictors. Approval tightness reflects dispersion of voter views when rating the incumbent, yet its information content is not subsumed by the investor-disagreement index of Huang et al. (2021), which is constructed as a principal component of 24 disagreement measures, including household forecasts and the Michigan Survey of Consumer Attitudes. Their disagreement index has strong predictive power for equity returns, with in- and out-of-sample R^2 of 2.52% and 1.56% at the one-month horizon. Our approval-tightness measure, rooted in an electoral-discipline channel, delivers comparable performance, forecasting aggregate stock returns robustly both in-sample and out-of-sample, with corresponding R^2 of 2.90% and 2.02% at the one-month horizon. Our results thus echo the vast literature on equity return predictability.[2]

Our paper is also related to the literature on how political and policy uncertainty affect asset valuation. Kelly et al. (2016) test the theory of Pástor and Veronesi (2013) and show that electoral uncertainty commands a risk premium, using option prices on equity market indices around election events in 20 countries. Brogaard and Detzel (2015) show that U.S. Economic Policy Uncertainty (EPU) positively forecasts excess equity returns. Bretscher et al. (2020) use a VAR with time-varying uncertainty to infer latent volatility in government spending and corporate income tax rates and find that higher government-spending uncertainty predicts lower output growth and capital investment and higher excess returns on Treasury bonds.

Our model builds on the optimal fiscal policy literature in which governments manage aggregate risk through tax smoothing (Barro, 1979; Lucas and Stokey, 1983; Bohn, 1990; Jiang et al., 2024a). We depart from the canonical setup by allowing the government to extract rents via inefficient spending. Related theoretical work on how elections discipline rent extraction includes Barro (1973), Ferejohn (1986), and Yared (2010). More broadly, our paper connects to research on how political power shapes fiscal and macroeconomic policy (e.g., Nordhaus, 1975; Rogoff, 1990; Alesina and Tabellini, 1990; Acemoglu et al., 2008; Boyer et al., 2024), among many others.

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

2 Data

2.1 Presidential Approval Rating (PAR) Index

We measure public support for the incumbent president using Gallup's Presidential Approval Rating (PAR). Gallup surveys a nationally representative panel of approximately 1,500 adults one or two times per month, although some months contain no poll. Respondents are asked: "Do you approve or disapprove of the way [current president] is handling his job as president?" Their answers are classified into three category: Approve, Disapprove, or Unsure/NoData. Because the third category conflates neutral views with non-response, we discard it and rescale the Approve and Disapprove shares so that they sum to 100%.

We make three data-processing adjustments to ensure that the presidential approval rating we construct uses the most recent public information, while importantly do not induce look-ahead bias in our empirical analysis. Firstly, when some polls span among months (start in month t but end in the month t+1), we assign their results in month t+1. Secondly, if Gallup conducts no poll in a given month, we forward-fill that observation with the most recent non-missing value. Thirdly, because Gallup often release their data in the next month after the polls are conducted, we consider one month publication lag for PAR. Then presidential approval ratings are aggregated into monthly, quarterly, or annual data. To reduce potential measurement errors and capture long-term trends, we smooth the monthly series with a six-month moving average (two quarters for the quarterly series; no smoothing is applied to annual data). The sample spans January 1969 through December 2024, covering the full presidential tenures from Richard M. Nixon to Joe Biden. We do not extend the sample further back because the early Gallup series contains long stretches of consecutive missing observations.

Figure 1 displays the evolution of Gallup presidential approval from January 1969 to December 2024. The series is smoothed using a six-month moving average and is segmented by administration. Warm-colored segments indicate Republican presidents, including Richard M. Nixon, Gerald R. Ford, Ronald Reagan, George H. W. Bush, George W. Bush, and Donald Trump. Cool-colored segments indicate Democratic presidents, including Jimmy Carter, Bill Clinton, Barack Obama, and Joe Biden. Approval exhibits substantial variation across administrations, with sharp movements around major historical episodes and persistent swings over multi-year horizons. Several

pronounced increases occur near the beginning of new presidencies, followed by gradual mean reversion within terms. Shaded recession intervals coincide with several notable shifts in the approval series, highlighting the close interaction between presidential approval dynamics and the business-cycle environment.

2.2 The Role of Presidential Approval Rating Tightness (PART)

Our objective is to study how political support affects asset prices and economic growth. A simple linear specification that uses PAR_t as a predictor delivers, at best, weak full-sample predictive relations for returns and growth. However, the data exhibit pronounced state dependence. Figure 2 illustrates this point by plotting one-quarter-ahead equity excess returns, economic uncertainty as measured by Jurado et al. (2015)[3], real GDP growth, and real PCE growth against PAR_t. The dashed gray line reports the full-sample linear fit. Splitting the sample at the median of PAR_t (vertical dashed line) reveals a clear threshold nonlinearity and a "slope-reversal" pattern: within the low-approval and high-approval subsamples, the fitted slopes are economically large and statistically significant, but they have opposite signs. As a result, the full-sample slope is attenuated toward zero even though predictability is strong within each regime.

As motivating evidence, Figure 2 points to a hump-shaped relation for growth and the equity risk premium, and a mirror-image U-shaped relation for uncertainty. Subsequent growth and equity excess returns are highest when presidential approval is intermediate, while uncertainty is lowest in that same region; both high-approval and low-approval regimes are associated with weaker growth and returns and elevated uncertainty. Intuitively, intermediate approval reflects a political environment in which electoral competition is tight and re-election outcomes are at a knife-edge. To capture this state, we define presidential approval tightness as

$$\text{PART}_t \equiv \text{PAR}_t\big(1 - \text{PAR}_t\big). \tag{1}$$

This statistic is maximized at 0.25 when PAR_t is near one half, i.e., when marginal votes are most valuable, and it declines as PAR_t moves away from one half toward either extreme. Thus, PART_t captures a distinct dimension of the political state: while PAR_t measures the *level* of approval,

[3]On the authors' website, this series is now labeled *financial uncertainty* and is accompanied by separate measures of real and macro uncertainty. We examine all three measures in the subsequent empirical analysis.

PART_t measures the *tightness* of political support and, therefore, the strength of marginal electoral incentives.

Because PART_t peaks at intermediate approval levels, including this tightness measure in a linear regression framework captures the threshold nonlinearity and slope-reversal pattern documented in Figure 2. In the remainder of the paper, we use PART_t as a primary predictor to study the relation between political approval and economic outcomes. We further develop a theoretical framework that links political incentives to macroeconomic outcomes and asset prices, explaining why knife-edge, high-tightness political states are associated with higher subsequent growth and returns and lower uncertainty. Central to our framework is the idea that tighter electoral conditions discipline inefficient government spending and rent extraction; fluctuations in PART_t therefore transmit to growth and asset returns through a fiscal channel.

In implementation, we compute PART_t from the smoothed monthly PAR_t series (using the same trailing moving average) and then aggregate it to quarterly and annual frequency as needed. Figure 3 plots PART_t from January 1969 to December 2024. Over the full sample, PART_t averages 0.233 and typically lies between 0.20 and 0.25, consistent with approval rates that remain away from the extremes. The series displays pronounced troughs during episodes in which approval moves far from 50%—either unusually high or unusually low—corresponding to periods of entrenched political support or pervasive disapproval.

For our baseline results, we use the smoothed series of PAR_t and PART_t, shown in Figure 1 and Figure 3, as the primary predictors in order to reduce measurement noise. All results remain statistically significant when we instead use the original, unsmoothed series; these robustness checks are reported in Appendix B. As an additional robustness exercise, we remove abrupt shifts ("jumps") in PAR_t and PART_t and repeat our empirical analysis. The results remain qualitatively unchanged across all specifications, as documented in Appendix B.

2.3 Asset Pricing and Macroeconomic Variables

The dependent variables and control variables are constructed as follows. We proxy the aggregate equity market return using the CRSP total value-weighted return index. The dividend–price ratio is constructed by combining the CRSP price index with dividends and the corresponding price index excluding dividends, which allows us to recover dividend payments and form the dividend–

price ratio. We also include standard return predictors compiled by Welch and Goyal (2008), including stock variance, the book-to-market ratio, net equity expansion, the Treasury bill rate, the long-term yield, the long-term return, the term spread, the default yield spread, the default return spread, and the earnings–price ratio. Equity excess returns are measured as continuously compounded log returns net of the risk-free rate (from the same source). Following Baker and Wurgler (2006), macroeconomic controls include industrial production growth, consumption growth for durable goods, nondurable goods, and services, employment growth, inflation, and an NBER recession dummy. All series are sampled at monthly frequency from 1969:01 to 2024:12 and are aggregated to quarterly or annual frequency as needed. In Appendix B, we also collect Treasury yield data from Liu and Wu (2021).

The volatility indices include the VIX, the Economic Policy Uncertainty index of Baker et al. (2016), and the financial, macro, and real uncertainty measures of Jurado et al. (2015). All volatility measures are standardized. VIX is available from 1990:01 to 2024:12, while the remaining measures span 1969:01 to 2024:12.

To capture economic fundamentals, we focus on real growth in output and consumption. Quarterly GDP and PCE from the National Income and Product Accounts (NIPA) are converted to real growth by subtracting CPI inflation (from FRED) from their log growth rates; cumulative real growth over horizon h is constructed analogously. These series are sampled at quarterly frequency over 1969Q1–2024Q4 and aggregated to annual frequency when required.

Real growth expectations are obtained from the Survey of Professional Forecasters (SPF). The SPF reports level forecasts for real consumption and real GDP, and we mirror our construction of realized real growth to obtain expected real GDP and PCE growth. We use one-quarter-ahead and four-quarter cumulative growth expectations at the quarterly frequency, and one-year-ahead growth expectations at the annual frequency. GDP expectations are available from 1969Q1 (one-quarter ahead) and 1969Q4 (cumulative four-quarter), whereas PCE expectations and annual expectations start in 1981 (PCE: 1981Q3; annual: 1981). Missing observations in the cumulative four-quarter GDP growth expectations are filled forward using the most recent available growth expectation.

We obtain annual Worldwide Governance Indicators (WGI) data from the World Bank. The WGI summarize perceptions of governance quality using information from more than 30 underlying sources, including think tanks, international organizations, non-governmental organizations, and

private firms. These inputs are aggregated into six dimensions: Voice and Accountability; Political Stability and Absence of Violence and Terrorism; Government Effectiveness; Regulatory Quality; Rule of Law; and Control of Corruption. The dataset covers more than 200 economies from 1996 to 2025, with annual country-level scores on each dimension reported on a 0–100 scale. We focus on the United States and use changes in U.S. scores to capture time variation in governance quality. We define a comprehensive governance-capacity measure as the average change across the six dimensions.

We also collect data summarizing the fiscal position of the government. Fiscal data are primarily drawn from NIPA tables available from FRED or the BEA. Following Berndt et al. (2012), we construct tax revenues and government spending using NIPA Table 3.2. Taxes are measured as total receipts (line 37). Government spending is defined as current expenditures (line 41) plus gross government investment (line 42) plus capital transfer payments (line 43), minus consumption of fixed capital (line 45) and net interest payments (line 29). For the market value of public debt, we use the series on marketable debt compiled by Hall and Sargent (2011) and Hall et al. (2018). Using these data, we construct measures of fiscal position including the log debt-to-GDP ratio (used by, e.g., Cochrane, 2022; Liu, 2023; Jiang et al., 2024b), the tax-to-GDP ratio, the log expenditure-to-GDP ratio, and the surplus-to-debt ratio following Campbell et al. (2023).

2.4 Summary Statistics and Stationarity Tests

Table 1 reports summary statistics for the main predictors, asset returns, and economic growth variables. Panel A shows that presidential approval averages about 0.55, indicating that voters express net support for the incumbent president on average. Approval tightness averages 0.233, close to its theoretical upper bound of 0.25, suggesting that approval often lies near the knife-edge region in which marginal votes are most valuable. The correlation between PAR and PART is strongly negative, at approximately −0.70, reflecting the right-skewness of PAR over the sample. In our regression analysis, the variance inflation factor for PAR and PART is around 2, indicating that multicollinearity is not a serious concern.

Neither PAR nor PART is strongly correlated with the equity valuation ratio dp (the dividend–price ratio), suggesting that political approval captures variation distinct from financial valuation risk. In contrast, political approval is correlated with measures of federal debt valuation. We

consider two such measures. The first is the net surplus-to-debt ratio *nsb*, which captures fiscal imbalance relative to current debt valuation and is used by Campbell et al. (2023) and Gao et al. (2025). The second is the log debt-to-GDP ratio, which measures public debt relative to output and is widely used in the literature (e.g., Cochrane, 2022; Liu, 2023; Jiang et al., 2024b; Jiang et al., 2024c). Panel A shows that PAR is positively correlated with *nsb* and negatively correlated with the debt-to-GDP ratio, whereas PART exhibits the opposite pattern, suggesting that approval levels and approval tightness are closely related to fiscal conditions and debt valuation.

Panel A also reports stationarity tests for PAR and PART to ensure the validity of our predictive analysis. We report results from the Augmented Dickey–Fuller test, the Phillips–Perron test, and the Dickey–Fuller GLS test, each of which has a unit root as the null hypothesis. Both PAR and PART reject the unit-root null at the 1% level across all three tests. The net surplus-to-debt ratio *nsb* also rejects the unit-root null and is therefore treated as stationary; consistent with Campbell et al. (2023), we use it as a control variable in predicting government expenditure growth in subsequent sections. By contrast, the dividend–price ratio *dp* fails to reject the unit-root null across all three tests, indicating a high degree of persistence consistent with near-nonstationarity.[4] The U.S. log debt-to-GDP ratio also appears close to a unit-root process. This finding replicates results in the literature showing that, in recent samples, U.S. public debt is valued above fundamentals due to its role as a safe asset.[5]

Panel B summarizes quarterly equity excess returns and real growth rates for GDP, PCE, and federal government expenditure. The annualized equity excess return has a mean of about 5.2% and a volatility of 17.6%. Average growth rates range from roughly 2.2% to 6.4% per year, with government expenditure growth being the most volatile, exhibiting a standard deviation of about 12%. GDP and PCE growth are strongly positively correlated, while both are negatively correlated with government expenditure growth, indicating that government spending growth is countercyclical over the sample. All return and growth series reject the unit-root null at the 1% level across the three tests, indicating that these outcome variables are well characterized as stationary.

[4]The literature applies structural or statistical restrictions to address the near unit-root behavior of the dividend–price ratio; see, for example, Timmermann, 1996, Lettau and Van Nieuwerburgh, 2008, and Koijen and Van Nieuwerburgh, 2011.

[5]See, among others, 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.

3 Empirical Analysis

3.1 PART and Predictability Results

3.1.1 PART Predicts Economic Growth

Table 2 reports predictive regressions of real economic growth on presidential approval, PAR_t, and presidential approval tightness, PART_t. The dependent variables are real GDP growth and real PCE growth, reported in the left and right blocks of the table, respectively. Panels differ by sampling frequency and forecast horizon. Panel A uses quarterly data and forecasts next-quarter growth. Panel B uses quarterly data and forecasts cumulative four-quarter growth. Panel C uses annual data and forecasts next-year growth.

In each panel and for each dependent variable, we estimate

$$Y_{t:t+h} = \alpha + \beta^{PAR} PAR_t + \beta^{PART} PART_t + \beta^{Controls} Controls_t + \varepsilon_{t:t+h},$$

where $Y_{t:t+h}$ denotes cumulative economic growth from t to $t+h$. Columns (1) and (2) report univariate specifications that include PAR_t or PART_t separately. Column (3) includes both approval measures jointly. Column (4) adds control variables that account for persistence in growth and for real-time expectations. The persistence control is the contemporaneous growth rate, $GDPG_t$ or $PCEG_t$, depending on the outcome.

The expectations control is constructed from the Survey of Professional Forecasters (SPF) and is matched to both the sampling frequency and the forecast horizon. When forecasting next-quarter growth in Panel A, we include the SPF expectation of next-quarter growth, $E_t[GDPG_{t+1}]$ or $E_t[PCEG_{t+1}]$. When forecasting cumulative four-quarter growth in Panel B, we include the SPF expectation of four-quarter growth, $E_t[GDPG_{t:t+4}]$ or $E_t[PCEG_{t:t+4}]$. When forecasting next-year growth in Panel C, we include the SPF expectation of next-year growth, $E_t[GDPG_{t+1}]$ or $E_t[PCEG_{t+1}]$, measured at the annual frequency. Newey–West standard errors are computed using four lags.

Panel A reports quarterly one-step-ahead forecasts for GDP and PCE growth. For GDP growth, the univariate regression with PAR_t has little explanatory power, whereas the univariate regression with PART_t is statistically significant at the 1% level. The joint specification that includes both

variables delivers substantially stronger predictability: both coefficients are statistically significant at the 1% level, and the adjusted R^2 rises to 5.46% from 0.09% and 1.01% in the respective univariate regressions. The same pattern holds for PCE growth. Both PAR_t and PART_t are statistically significant in the joint regression, and the adjusted R^2 increases to 4.37% from 0.50% and 0.19%. When we add survey expectations and lagged growth in column (4), the approval measures remain jointly informative and statistically significant, and the model fit increases sharply, with adjusted R^2 reaching 26.22% for GDP growth and 28.06% for PCE growth.

Panels B and C examine longer-horizon effects. Panel B uses cumulative four-quarter growth, which produces overlapping dependent variables, while Panel C uses annual one-year-ahead growth, which avoids overlapping outcome windows. The qualitative conclusions are consistent across the two approaches. For four-quarter GDP growth, the joint specification raises the adjusted R^2 to 17.67%, relative to 1.51% and 3.23% in the univariate regressions, and both PAR_t and PART_t are statistically significant. For four-quarter PCE growth, the joint regression similarly increases the adjusted R^2 to 16.00% from 2.97% and 1.28%. In the non-overlapping annual specifications, the joint regression again delivers sizable gains in explanatory power, with adjusted R^2 reaching 18.57% for GDP growth and 16.94% for PCE growth. These results remain robust after adding horizon-matched SPF expectations and lagged growth. Taken together, the overlapping four-quarter results and the non-overlapping one-year results corroborate the conclusion that PAR_t and PART_t contain complementary information for forecasting real economic growth at longer horizons.

3.1.2 PART Predicts Equity Returns and Volatility Indices

Table 3 evaluates the predictive power of presidential approval, PAR_t, and approval tightness, PART_t, for equity excess returns using the predictive regression

$$R_{t:t+h} = \alpha + \beta^{PAR} PAR_t + \beta^{PART} PART_t + \beta^{Controls} Controls_t + \varepsilon_{t:t+h},$$

where $R_{t:t+h}$ denotes the cumulative equity excess return from t to $t+h$. Columns (1) and (2) report univariate specifications in which PAR_t or PART_t is the sole predictor. Column (3) includes both approval measures jointly. Column (4) adds a comprehensive set of control variables, including the financial return predictors compiled by Welch and Goyal (2008) and macroeconomic controls

following Baker and Wurgler (2006).

To assess robustness across sampling frequencies, Panel A reports monthly regressions and Panel B reports quarterly regressions. Within each panel, the left block uses a one-period horizon, while the right block uses a longer cumulative horizon, with $h = 1$ and $h = 12$ months in Panel A and $h = 1$ and $h = 4$ quarters in Panel B. In Panel A, the univariate regression with PAR_t again exhibits little predictive content and yields a negative adjusted R^2 of -0.14%, whereas PART_t is positive and statistically significant at the 1% level, with an adjusted R^2 of 1.30%. The joint specification shows clear incremental explanatory power: both PAR_t and PART_t are statistically significant, and the adjusted R^2 increases to 2.90%. This result is robust to controlling for the full set of standard predictors used in the literature, including the dividend–price ratio; in these specifications, both PAR_t and PART_t retain economically large and statistically significant coefficients. The longer-horizon 12-month regressions confirm the same pattern, although the overlapping nature of the dependent variable mechanically raises the overall level of R^2.

Panel B reports quarterly forecasts and delivers qualitatively similar conclusions. At the one-quarter horizon, PAR_t alone has limited predictive content, whereas PART_t is positive and statistically significant, with an adjusted R^2 of 3.14%. When PAR_t and PART_t are included jointly, both coefficients are statistically significant and the model fit improves substantially, with the adjusted R^2 rising to 7.85%. After adding the standard control set, both predictors remain significant and the adjusted R^2 increases further to 15.24%. The right block reports cumulative four-quarter excess returns and yields a similar pattern: PART_t remains predictive, the joint specification improves fit relative to either univariate regression, and the control-augmented specification delivers the strongest performance, with adjusted R^2 reaching 21.60%.

Taken together, across horizons and sampling frequencies, the evidence indicates that PAR_t and PART_t contain complementary information for forecasting equity excess returns, and that their joint predictive power remains pronounced after conditioning on standard financial and macroeconomic predictors.

For robustness, Table 4 further evaluates whether this predictability survives after conditioning on prominent proxies for public beliefs and political-cycle conditions.

The public-belief category comprises two predictors. The first is the economic belief disagreement index of Huang et al. (2021). This measure aggregates dispersion in expectations across a

broad set of underlying disagreement series and is motivated by theories in which disagreement is linked to uncertainty and belief distortions that affect expected returns. The second predictor is the Economic Policy Uncertainty index of Baker et al. (2016). Both series are used in standardized form, so the corresponding coefficients quantify return predictability per standard deviation change in the belief proxy. Quarterly values of these series are computed as quarterly averages of the monthly data.

The political-cycle category comprises six indicators capturing partisan effects and election-timing dynamics. Since the inauguration of a new president is typically around January 20, we treat February as the first month of a new four-year presidential term, with an exception that August 1974 as the start of President Ford's term. The *Honeymoon* indicator equals one during the first twelve months of a new presidency and zero otherwise. The variable *Party* equals one when the incumbent president is a Democrat and zero otherwise. The indicators $Midterm^{Before}$ and $Midterm^{After}$ equal one during the five months immediately preceding and following a midterm election, and $Election^{Before}$ and $Election^{After}$ are defined analogously for presidential elections. These variables are constructed at the monthly frequency and then aggregated to quarterly frequency using the within-quarter rule described above: quarterly values of the dummy variables equal one if any of the three months in the quarter takes the value one and zero otherwise.

We estimate one-step-ahead predictive regressions at the monthly frequency in Panel A and at the quarterly frequency in Panel B. Each row augments the baseline specification that includes PAR_t and PART_t with the additional predictor listed in the *Controls* column, while the row labeled *All* includes all control predictors jointly. Panel A shows that the predictive content of PAR_t and PART_t is not subsumed by the competing predictors. Across all specifications, both coefficients remain positive and statistically significant, including when we condition on the belief proxies *Disagreement* and *EPU* and on the political-cycle indicators. Panel B delivers a similar conclusion at the quarterly horizon. The coefficients on PAR_t and PART_t remain positive and statistically significant in every row, indicating that their joint predictability is robust to conditioning on belief and political-cycle predictors.

Crucially, the *All* row shows that PAR_t and PART_t remain predictive when the full set of public-belief and political-cycle controls is included simultaneously, accompanied by a sizable increase in explanatory power. This evidence supports the interpretation that presidential approval and

approval tightness forecast equity excess returns through channels that are not spanned by existing belief-based proxies and standard political-cycle predictors. Instead, PAR_t and PART_t appear to contain complementary information about political conditions that helps to predict future equity returns.

Table 5 examines whether presidential approval and approval tightness predict subsequent volatility indices using the predictive regression

$$Y_{t:t+h} = \alpha + \beta^{PAR} PAR_t + \beta^{PART} PART_t + \beta^{Control} Control_t + \varepsilon_{t:t+h},$$

where $Y_{t:t+h}$ denotes the average volatility level over the window between month t and month $t+h$. The volatility measures include the CBOE Volatility Index (VIX), constructed from S&P 500 index options; the Economic Policy Uncertainty (EPU) index of Baker et al. (2016); and the financial, macro, and real uncertainty measures of Jurado et al. (2015). The control vector $Control_t$ includes lagged uncertainty Y_t to account for persistence in uncertainty dynamics. We consider forecast horizons $h = 1$, $h = 6$, and $h = 12$ to capture both short-run and longer-run effects. All uncertainty proxies are standardized to facilitate comparability across measures. Inference is based on Newey–West standard errors, with the HAC lag length set to 6 for $h = 1$ and $h = 6$, and to 12 for $h = 12$.

The table delivers two main findings. First, Table 5 shows a broadly negative association between presidential approval and subsequent uncertainty across horizons. With the exception of the VIX, the coefficients on PAR_t are negative and statistically significant for all uncertainty proxies at $h = 1$, $h = 6$, and $h = 12$. This pattern is consistent with the evidence in Liu and Shaliastovich (2022) that higher presidential approval is associated with lower uncertainty. Second, approval tightness exhibits an even more uniform pattern. The coefficients on PART_t are negative across all uncertainty measures and horizons, with the single exception of the EPU measure at $h = 1$, where PART_t is not statistically significant.

Taken together, the results indicate that both the level and the tightness of presidential approval contain systematic information about future volatility, with the tightness measure delivering more robust and pervasive predictive content across uncertainty proxies and forecast horizons. These findings highlight the special role of political states in which incumbent re-election is at a knife-edge—characterized by intermediate PAR and high PART—which are associated with high asset

returns and low volatility. Consistent with our theoretical framework, heightened approval tightness disciplines inefficient government spending and reduces the associated political uncertainty. As an associated implication, we also test the bond return predictability and put the results in Appendix B for the interest of space.

3.1.3 Robustness: Small-Sample Bias Adjustment

Inference in return and growth predictability regressions can be distorted in finite samples when the predictors are highly persistent. Table 1 shows that PAR and PART reject unit roots but remain substantial persistence. In such settings, Stambaugh (1999) shows that if the predictor innovations are correlated with the regression error, OLS can suffer from small-sample bias, so the estimated forecasting slope is systematically distorted even when the regressor is stationary. Amihud and Hurvich (2004) propose a practical bias-correction procedure for predictive regressions with persistent predictors. We follow their approach and re-estimate our main specifications.[6]

Table 6 reports Amihud-Hurvich bias-adjusted predictive regressions of equity excess returns on PAR_t and PART_t. Panel A presents monthly results and Panel B presents quarterly results. Within each panel, the left block reports one-period-ahead returns and the right block reports longer-horizon cumulative returns. In Panel A, the one-month horizon results mirror the baseline OLS findings. PAR_t alone has little predictive content, while PART_t is strongly significant. When the two variables are included jointly, both coefficients remain statistically significant and the model fit improves relative to either univariate specification. This complementarity persists after adding the full control set, with both coefficients remaining significant. The right block of Panel A shows that the same pattern extends to 12-month cumulative returns. PART_t remains significant, and the joint specification improves explanatory power and delivers statistically significant coefficients once controls are included. Panel B yields the same qualitative conclusions at the quarterly frequency. For one-quarter-ahead returns, both PAR_t and PART_t are significant in the joint specification, and the control-augmented regression strengthens the results further. For four-quarter cumulative returns, the joint specification again delivers statistically significant coefficients and higher explana-

[6] Amihud and Hurvich (2004) provide a straightforward implementation for univariate and multivariate settings when the VAR(1) coefficient matrix of predictors is diagonal. When this condition fails, accounting for estimation error becomes more involved. We follow their recommended procedure: we first estimate a VAR(1) for the predictor set and implement the off-diagonal correction only when the VAR exhibits sufficiently strong cross-predictive dynamics.

tory power once controls are included. Taken together, the persistence of joint predictability across panels, horizons, and control specifications after small-sample bias adjustment indicates that the complementary forecasting power of PAR_t and PART_t is not an artifact of small-sample bias.

Table 7 reports Amihud-Hurvich bias-adjusted predictive regressions of real GDP and real PCE growth on PAR_t and PART_t. Panel A presents one-quarter-ahead forecasts. The joint specification continues to deliver strong predictability for both GDP and PCE growth. After adding horizon-matched SPF expectations and lagged growth in column (4), the estimated coefficients on PAR_t and PART_t remain economically meaningful, but their statistical significance is somewhat weaker than in the baseline OLS results, especially for PART_t. Panel B uses cumulative four-quarter growth and continues to show large and statistically significant coefficients on both PAR_t and PART_t, including after controlling for SPF expectations and lagged growth. Panel C uses annual one-year-ahead growth, which avoids overlapping windows, and delivers the same qualitative conclusion. The joint specification remains strongly significant and explains a substantial share of variation in both GDP and PCE growth. Taken together, the four-quarter and non-overlapping one-year results indicate that the joint predictive power of PAR_t and PART_t is robust to small-sample bias adjustment and is not driven by finite-sample distortions.

3.1.4 Robustness: Out-of-sample Results

The in-sample evidence in Table 3 indicates that presidential approval (PAR) and approval tightness (PART) are jointly informative about future excess returns. We next ask whether this predictability survives an out-of-sample evaluation. Table 8 summarizes out-of-sample forecasting performance for equity excess returns and real growth rate of GDP, PCE, using PAR, PART, and their combination (PAR+$PART$) as predictors.

Forecasts are generated from recursively (expanding-window) estimated predictive regressions at the horizon and sampling frequency indicated in each panel. The initialization window is 1969-1996, and the out-of-sample evaluation period runs from 1997 through 2024. Forecast accuracy is assessed relative to the historical-mean benchmark using the Campbell and Thompson (2008) out-of-sample R^2 statistic,

$$R^2_{\text{OS}} = 1 - \frac{\sum_{t=M+1}^{T} \left(Y_t - \hat{Y}_t\right)^2}{\sum_{t=M+1}^{T} \left(Y_t - \bar{Y}_t\right)^2}, \tag{2}$$

where $\hat{Y}_t$ is the model-based forecast and $\overline{Y}_t$ is the recursive historical mean. Statistical significance is evaluated using the MSFE-adjusted Clark-West statistic (Clark and West, 2007). To avoid look-ahead bias and complications from overlapping multi-horizon prediction, we do not implement multi-horizon out-of-sample forecasts; instead, we report one-step-ahead predictions at different sampling frequencies (e.g., one-month versus one-quarter for returns, and one-quarter versus one-year for growth).

Table 8 highlights substantial complementarity between PAR and PART in out-of-sample forecasting. Panel A reports out-of-sample results for equity excess returns. Using PAR alone yields negative R^2_{OS} at both the monthly and quarterly frequencies. In contrast, PART delivers economically meaningful improvements, with R^2_{OS} equal to 2.02% at the one-month horizon and 5.06% at the one-quarter horizon, and both gains are statistically significant at the 1% level based on the Clark-West test. Importantly, combining the two predictors strengthens performance further, raising R^2_{OS} to 3.47% at the monthly frequency and 8.98% at the quarterly frequency; the corresponding Clark-West statistics reject the null of no improvement at the 1% and 5% levels, respectively.

For macroeconomic outcomes, the joint specification is especially powerful. Panel B reports out-of-sample results for GDP growth. Neither PAR nor PART alone delivers robust forecast gains, whereas their combination generates statistically significant improvements at both frequencies, with R^2_{OS} equal to 1.93% at the one-quarter horizon and 24.46% at the one-year horizon. Panel C reports out-of-sample results for PCE growth and shows a similar pattern. The combined predictor yields R^2_{OS} of 1.88% at the one-quarter horizon and 18.64% at the one-year horizon, and these gains are statistically significant at conventional levels, while the univariate specifications provide little out-of-sample improvement.

3.2 PART as a Sensitivity of Approval to Economic Fundamentals

We document a new stylized fact about the interaction between the level and the tightness of presidential approval. The responsiveness of approval to economic fundamentals is strongest when approval lies in an intermediate range and markedly weaker when approval is either very high or very low. This pattern accords with a knife-edge political environment in which marginal voters are most pivotal, magnifying the electoral consequences of economic performance. Accordingly, presidential approval rating tightness—defined as the hump-shaped transformation of PAR—captures variation

in sensitivity that is not reflected by the approval level alone.

We illustrate this finding in several steps. First, a large literature shows that presidential approval responds to economic fundamentals and political events (e.g., Mueller, 1970; Kriner and Schickler, 2014; Gordon, 2024). Second, we emphasize a feature that is often left implicit in this literature: by construction, the presidential approval rate PAR is bounded in $[0,1]$, akin to a probability. Any realistic model of approval dynamics must therefore respect this bounded support by mapping an underlying (potentially unbounded) index of economic and political conditions into an approval rate. We capture this idea with a generic CDF link,

$$\mathrm{PAR}_t = \Lambda(\ell_t),$$

where ℓ_t is an economic fundamentals index and $\Lambda(\cdot)$ is a cumulative distribution function. Under this mapping, the marginal sensitivity of approval to fundamentals is given by $\Lambda'(\ell_t)$, the associated probability density function. For any unimodal distribution—a property satisfied by a broad class of standard distributions—$\Lambda'(\ell)$ is single-peaked. When this mapping has moderate skewness, as a result, approval is most responsive in the intermediate range and progressively less responsive near the bounds, generating the hump-shaped sensitivity pattern we document.

We take a logit mapping as our baseline specification,

$$\mathrm{PAR}_t = \Lambda(\ell_t), \qquad \Lambda(\ell) = \frac{1}{1+\exp(-\ell)}, \tag{3}$$

where ℓ_t is an economic fundamentals index that depends linearly on economic performance and political events. This specification is straightforward to estimate using standard logistic regression. More importantly, it delivers a direct interpretation of approval tightness as sensitivity to fundamentals: under Eq. (3), $\Lambda'(\ell) = \Lambda(\ell)\big(1-\Lambda(\ell)\big)$, so that

$$\frac{\partial \mathrm{PAR}_t}{\partial \ell_t} = \Lambda(\ell_t)\big(1-\Lambda(\ell_t)\big) = \mathrm{PART}_t.$$

In practice, approval rates rarely approach 0 or 1. To better match the observed empirical

range, we also consider a scaled logistic specification,

$$\text{PAR}_t = a + b\,\Lambda(\ell_t),$$

where (a, b) determine the support $(a, a + b)$ and are calibrated to match the range of approval rates in our sample. We report results from the scaled specification as our main estimates; results from the standard logit specification are qualitatively similar.

To test the logit relation empirically, we invert the approval rate to recover a time series for the economic performance index ℓ_t and examine its relation to economic fundamentals,

$$\ell_t = \alpha + \beta\,\ell_{t-1} + \delta^{\top}\text{Politics}_{t-1} + \gamma\,X_t + \varepsilon_{t+1}, \tag{4}$$

where X_t collects economic variables including real GDP growth, real PCE growth, the unemployment rate, $\log(\text{Tax/GDP})_t$, and $\log(\text{Spending/GDP})_t$, and we include inflation as an additional control. Because tax and spending measures are subject to transitory shocks and are typically more volatile than real activity, we smooth these series using eight-quarter rolling windows.

We also include the lagged economic performance index ℓ_{t-1} and a vector of political-event indicators, Politics_{t-1}, to account for non-economic shocks to approval. Following Gordon (2024), these indicators are constructed using fixed event windows. Specifically, we include a Watergate indicator equal to one from 1973Q3 to 1974Q2, a Gulf War indicator from 1991Q1 to 1991Q2, a 9/11 indicator from 2001Q4 to 2003Q2, and a post-Katrina Bush indicator from 2005Q4 to 2008Q4. We also include administration-specific indicators for the Trump period (2017Q1–2020Q4) and for the Biden slump (2023Q3–2024Q4), as well as a post-millennium indicator equal to one from 2001Q1 to 2024Q4. Finally, we control for a honeymoon effect using an indicator equal to one during the first four quarters of each new presidency.

For robustness, we also estimate a scaled probit specification,

$$\text{PAR}_t = a + b\,\Phi(\ell_t),$$

where $\Phi(\cdot)$ denotes the standard normal CDF. Both logit and probit links are commonly used when modeling dependent variables bounded in $[0, 1]$. Because the hump-shaped sensitivity arises for

any unimodal distribution, our conclusions do not hinge on a particular link function. We therefore treat the logit specification as our baseline and use the probit results as a robustness check.

Table 9 reports the results linking approval to economic performance. Panel A presents estimates under the scaled logit mapping. In columns (1) and (2), real GDP growth and real PCE growth enter positively and are statistically significant, indicating that stronger real growth boosts presidential approval rate. In column (3), the unemployment rate enters negatively and is highly significant, and inflation is negative and significant across all specifications, consistent with weaker labor market conditions and higher inflation reducing approval. Columns (4) and (5) add fiscal variables; both $\log(\text{Tax/GDP})_t$ and $\log(\text{Spending/GDP})_t$ enter with negative coefficients, with spending strongly significant. Column (6) includes the full set of economic variables jointly. Unemployment, inflation, and the fiscal variables remain negative and statistically significant, while the marginal contribution of real GDP and PCE growth attenuates once the full set of controls is included.

Panel B reports the corresponding results under the scaled probit specification. The qualitative patterns closely mirror those in Panel A: real GDP and PCE growth enter positively, unemployment and inflation enter negatively, and the fiscal variables enter with negative coefficients and similar significance patterns.

Figure 4 visualizes the empirical link between presidential approval and economic fundamentals implied by the full-specification logit and probit regressions. Each panel plots the observed presidential approval rate on the vertical axis against the fitted value of the latent economic-performance index $\hat{\ell}_t$ on the horizontal axis. Blue markers represent the realized approval rate, while the red curve overlays the fitted mapping from the economic fundamental index to approval under the corresponding link function. The left panel uses a scaled logit specification, and the right panel uses a scaled probit specification.

Across both link functions, the mapping is strongly nonlinear, with the slope of approval with respect to economic performance being steepest in the intermediate range of approval and flattening toward the extremes. This pattern implies that changes in economic fundamentals translate into the largest movements in approval precisely when approval is near the knife-edge, while the same economic shocks have attenuated effects when approval is very high or very low. This evidence provides a direct visualization of the mechanism underlying our approval-tightness measure: approval

tightness $\text{PART}_t = \text{PAR}_t(1 - \text{PAR}_t)$ captures this variation in sensitivity in a parsimonious way. In this sense, PART_t summarizes the sensitivity of political approval to economic fundamentals, which in turn motivates its strong predictive power for growth, asset returns, and volatility documented in the preceding sections.

3.3 PART and Government Spending, Control of Corruption, and Government Capacity

Building on the finding that PART_t captures the sensitivity of presidential approval to economic fundamentals, we next test whether this sensitivity disciplines government behavior.

Table 10 examines whether approval tightness forecasts subsequent government expenditure growth using

$$dG_{t:t+h} = \alpha + \beta^{PAR} PAR_t + \beta^{PART} PART_t + \beta^{Controls} Controls_t + \varepsilon_{t:t+h},$$

where $dG_{t:t+h}$ denotes cumulative real federal government expenditure growth from t to $t+h$. To capture both short-run and longer-run effects, Panel A reports quarterly regressions for horizons $h = 1$ and $h = 4$. To corroborate the longer-horizon results without overlapping outcome windows, Panel B reports annual regressions for $h = 1$. We focus on the joint dynamics of PAR_t and $PART_t$ and progressively add controls across columns. Column (1) includes lagged real government expenditure growth dG_t. Column (2) additionally controls for inflation. Columns (3) through (5) further add one fiscal-conditions variable at a time.[7] Fiscal conditions are measured by the net surplus-to-debt ratio nsb_t, $\log(\text{Tax/GDP})_t$, and $\log(\text{Expenditure/GDP})_t$, where Tax denotes federal tax revenues and Expenditure denotes federal government expenditures.

The estimates indicate that tighter approval conditions—higher PART_t—are associated with lower subsequent government expenditure growth. Across horizons and sampling frequencies, PAR_t and $PART_t$ are jointly informative, and their coefficients are consistently negative. In the quarterly regressions, both variables are statistically significant for one-quarter-ahead growth and for cumulative four-quarter growth, with larger magnitudes at the longer horizon. These negative relations remain intact after controlling for inflation and conditioning on alternative measures of

[7] Because these fiscal-condition measures are highly collinear, each specification includes at most one of them.

fiscal conditions. The annual regressions deliver the same qualitative message for non-overlapping one-year growth, with both PAR_t and $PART_t$ remaining statistically significant.

Moreover, the negative effect of PART_t on subsequent expenditure growth complements standard fiscal-control channels implied by the government budget constraint. A stronger fiscal position, measured by a higher surplus-to-debt ratio nsb_t, predicts higher subsequent expenditure growth at both quarterly and annual horizons. By contrast, a higher lagged expenditure share, $\log(\text{Expenditure/GDP})_t$, predicts lower subsequent expenditure growth and is statistically significant across horizons. Importantly, controlling for fiscal conditions does not attenuate either the magnitude or the significance of the predictive effect of PART_t.

The negative relation between PART_t and subsequent government expenditure growth has a natural interpretation. A tighter electoral environment disciplines fiscal inefficiency by increasing the political cost of rent extraction and corruption when re-election is uncertain. A large literature documents how corruption and rent-seeking behavior affect economic performance and interact with political incentives (e.g., Shleifer and Vishny, 1993; Mauro, 1998; Duchin and Sosyura, 2012; Koijen et al., 2016; Kang, 2016; Gupta et al., 2025).

To further verify this channel, and consistent with our definition of federal government expenditure, we obtain annual national defense spending from NIPA Table 3.17 as the sum of items explicitly classified as national defense. We then define nondefense spending as total federal expenditure net of the national defense component. Table 11 reports the results in the left and right panels, using the same specification as in Table 10. As a necessary public good, national defense spending growth is highly persistent and is primarily predicted by its own lag, past inflation, and fiscal condition variables, but it exhibits no systematic relation with PART_t. In contrast, nondefense spending growth is less persistent and is not predicted by inflation, yet it is significantly and negatively predicted by PART_t. This contrast indicates that presidential approval tightness disciplines discretionary government spending to a much greater extent than mandated defense expenditures.

We further examine the electoral-discipline channel by testing whether tighter approval conditions are associated with lower corruption and higher government capacity, using perception-based measures from the Worldwide Governance Indicators. Table 12 evaluates this implication by relating one-year-ahead measures of governance quality to PAR_t and $PART_t$. The left block uses

the Control of Corruption index, and the right block uses the Governance Capacity index. Higher values of both measures indicate better governance outcomes. The data are annual and span 1997 to 2024. Specifications begin with univariate regressions that include either PAR_t or $PART_t$, then include both jointly, and finally add the lagged dependent variable Y_t to account for persistence in governance outcomes.

In the Control of Corruption regressions, the coefficient on PART_t is positive and statistically significant, and it remains significant when PAR_t is included jointly and when the lagged dependent variable is added. This implies that higher approval tightness is associated with improved control of corruption. In the Governance Capacity regressions, PART_t is again positive and highly significant across specifications, whereas PAR_t does not exhibit robust predictive content. Overall, the evidence indicates that tighter approval conditions are associated with stronger subsequent governance capacity and improved control of corruption.

Our theoretical framework incorporates this empirically validated channel as a central mechanism: higher PART_t strengthens electoral discipline, curbs rent-seeking behavior, and thereby affects economic performance and asset prices through fiscal policy adjustments.

4 Theoretical Framework

We develop a parsimonious political-economy framework that unifies our empirical findings. First, presidential approval tightness (PART) captures the sensitivity of approval to economic fundamentals and disciplines government behavior, including public spending and corruption. Second, PART forecasts equity returns, economic growth, and volatility. Central to our framework is the mechanism that a tighter electoral environment increases the responsiveness of approval to economic performance, thereby strengthening the government's incentive to curb rent-seeking behavior.

For notational convenience, throughout this section we use p to denote the presidential approval rate PAR, and define approval tightness as

$$d \;\equiv\; \mathrm{PART} \;\equiv\; p(1-p).$$

4.1 Fiscal Inefficiency and Rent Seeking

We start from a raw (pre-policy) output process Y_t^r that evolves exogenously, subject to a policy-induced volatility channel:

$$\frac{Y_{t+1}^r}{Y_t^r} = \exp\left(\mu_y - \frac{\sigma_y^2(x_t)}{2} + \sigma_y(x_t)\varepsilon_{t+1}\right), \qquad \varepsilon_{t+1} \sim \mathcal{N}(0,1), \tag{5}$$

where the growth shock is i.i.d. for tractability.

The government is mandated to allocate a constant share of raw output to public goods, $g \equiv \frac{G_t}{Y_t^r}$, $g \in (0,1)$, which we treat as exogenous. This spending enters production, so a constant share is consistent with a standard Cobb–Douglas technology. Beyond mandated spending, the government can extract rents by expanding outlays X_t. Define the rent share $x_t \equiv \frac{X_t}{Y_t^r}$, $x_t \geq 0$. This spending is inefficient and associated with policy uncertainty (see, for example, Pástor and Veronesi (2013)) and increases the risk of output growth. We capture this channel in reduced form by letting conditional volatility increase with x_t:

$$\sigma_y^2(x_t) = \sigma^2(1 + x_t), \qquad \sigma > 0. \tag{6}$$

This channel disciplines the government because higher volatility raises required compensation on claims to future fiscal surpluses, worsening debt valuation and increasing the cost of refinancing.

The government finances spending by collecting taxes that generate deadweight losses:

$$\tau_t \equiv \frac{T_t}{Y_t^r}, \qquad \theta(\tau_t) = \frac{1}{2}c^\tau \tau_t^2, \qquad c^\tau > 0. \tag{7}$$

Consistent with the public finance literature, $\theta(\tau)$ is smooth, convex, and increasing, which ensures interior optimal policies. Following Barro (1979), we adopt the quadratic specification.

Rent extraction also wastes resources. To extract one dollar of rents, the economy loses an additional $c^x > 0$ dollars of final output. Final output is therefore $Y_t = Y_t^r\big(1 - \theta(\tau_t) - c^x x_t\big)$. The government chooses (τ_t, x_t) and implements $T_t = \tau_t Y_t^r$, $X_t = x_t Y_t^r$. Households consume all resources left after mandated public goods and rent extraction, so goods-market clearing implies

$$C_t = Y_t - G_t - X_t = Y_t^r\Big(1 - \theta(\tau_t) - g - (1 + c^x)x_t\Big). \tag{8}$$

In asset markets, households hold (i) private equity as a claim on dividends and (ii) public debt as a claim on net fiscal surpluses. Dividends equal post-tax output:

$$D_t = Y_t - T_t = Y_t^r\Big(1 - \theta(\tau_t) - c^x x_t - \tau_t\Big). \tag{9}$$

Net surplus is $NS_t \equiv T_t - G_t - X_t = Y_t^r(\tau_t - g - x_t)$, so $C_t = D_t + NS_t$.

4.2 Government Objective: Debt Management, Rent Seeking, and Re-election

At date t, the incumbent observes the current approval rate $p_t \in (0,1)$ and chooses a *term policy* implemented over the next T periods. Concretely, the incumbent sets next-period fiscal rates (τ_{t+1}, x_{t+1}) and commits to keep them constant over the term:

$$(\tau_{t+i}, x_{t+i}) = (\tau_{t+1}, x_{t+1}), \qquad i = 1, \ldots, T.$$

At date $t+T$, economic outcomes are realized and the incumbent remains in office with probability p_{t+T}.

We assume a risk-free asset exists and, under no arbitrage, households price payoffs using a stochastic discount factor that loads on the i.i.d. output shock:

$$m_{t+1} \equiv \frac{M_{t+1}}{M_t} = \exp\left(-r - \frac{\eta^2 \sigma_y^2(x_t)}{2} + \eta \sigma_y(x_t)\varepsilon_{t+1}\right), \tag{10}$$

where η governs the exposure of marginal utility to output shocks and r is the (log) risk-free rate. The discount factor from t to $t+i$ is $m_{t+i} \equiv M_{t+i}/M_t$. The government takes Eq. (10) as given, including the mapping from policy-induced risk $\sigma_y(\cdot)$ to valuations.

Let B_t denote the market value of outstanding public debt. The government's intertemporal budget constraint is

$$B_t \le \mathbb{E}_t\left[\sum_{s=t+1}^{\infty} \frac{M_s}{M_t}\big(T_s - G_s - X_s\big)\right], \tag{11}$$

which requires that debt be backed by the present value of future net surpluses. Equivalently, debt

evolves according to the flow budget constraint

$$B_{t+1} = B_t(1 + R^B_{t+1}) - \big(T_{t+1} - G_{t+1} - X_{t+1}\big),$$

where the government takes debt valuation as given and manages the cash flows.

A benevolent government that only values household welfare would choose policy to maximize the discounted value of household consumption,

$$W_t = \mathbb{E}_t\left[\sum_{s=t}^{\infty} \frac{M_s}{M_t} C_s\right], \tag{12}$$

subject to Eq. (11). In this benchmark, rent extraction is never optimal: x_t reduces household resources directly (via $(1 + c^x)x_t$ in Eq. (8)) and increases macro risk through $\sigma_y(x_t)$, tightening debt valuation. Hence a benevolent government sets $x_t = 0$ and chooses τ_t to finance mandated spending g and service outstanding debt. Our model in this nested case reduces to the benchmark framework that government manage the aggregate output risk by smoothing tax (e.g. Barro (1979) and Jiang et al. (2024a)).

We depart from this benchmark by allowing the incumbent to extract private rents from discretionary spending $X_t = x_t Y^r_t$. In addition to valuing household consumption, the incumbent receives a private payoff $(1 + r^x)X_t = (1 + r^x)x_t Y^r_t$ each period. This payoff is outside can be interpreted as resources captured by the office holder and affiliated networks (e.g., patronage transfers, influence rents, or private enrichment). Combining the household resource cost with the private benefit, the incumbent's per-period flow payoff can be written as

$$C_t + (1 + r^x)X_t = Y^r_t\Big(1 - \theta(\tau_t) - g + \kappa x_t\Big), \qquad \kappa \equiv r^x - c^x, \qquad \theta(\tau) = \tfrac{1}{2}c^\tau \tau^2.$$

The net payoff parameter $\kappa > 0$. Our setting borrows the spirits in Barro (1973) and assumes linear function format for simplicity. The incumbent therefore chooses term policy (τ_{t+1}, x_{t+1}) to maximize expected discounted payoffs over the remaining term and the continuation value of holding power, subject to Eq. (11). Here (τ_{t+1}, x_{t+1}) denotes the constant rates chosen at date t

and implemented over $t+1, \ldots, t+T$:

$$V(B_t, Y_t^r, p_t) = \max_{\tau_{t+1},\, x_{t+1}} \mathbb{E}_t\left[\sum_{i=1}^{T} m_{t+i} Y_{t+i}^r\Big(1 - \theta(\tau_{t+1}) - g + \kappa x_{t+1}\Big)\right] + \mathbb{E}_t p_{t+T} \mathbb{E}_t\Big[m_{t+T}\, V(B_{t+T}, Y_{t+T}^r, p_{t+T})\Big]. \tag{13}$$

Eq. (13) highlights the incumbent's trade-off. A higher rent share x_{t+1} raises private payoffs, but it lowers net surpluses and increases macro risk through $\sigma_y(x)$, tightening the intertemporal budget constraint. Further, the re-election adds an extra consideration of incumbent: it consider how fiscal policy might depress performance reduce approval at the election date and thus lower the continuation value from holding office. We interpret p_{t+T} as a reduced-form measure of expected political influence disciplined by the election, and model the continuation value as proportional to p_{t+T}. This captures a two-party environment in which political control is not purely winner-take-all and the opposition retains meaningful influence over policy outcomes. A discontinuous majority-threshold election rule would require specifying an explicit out-of-office value and regime-switching dynamics, making the model less tractable; we therefore deviate from that specification.

4.3 Role of PART: Electoral Discipline on Fiscal Policy

As in the empirical section, we model the observed approval rate as a logistic transformation of a latent approval index:

$$p_{t+1} = \Lambda(\ell_{t+1}), \qquad \Lambda(\ell) = \frac{1}{1 + \exp(-\ell)}.$$

The latent index loads on an incumbent-performance measure q_{t+1}:

$$\ell_{t+1} - \bar{\ell} = \phi\big(\ell_t - \bar{\ell}\big) + \gamma\big(q_{t+1} - \bar{q}\big) + \epsilon_{t+1}^{\ell}, \qquad \epsilon_{t+1}^{\ell} \text{ i.i.d.}, \tag{14}$$

where $\gamma \geq 0$ governs how strongly voters translate performance into approval, and ϵ_{t+1}^{ℓ} captures non-economic shocks (e.g., political events) orthogonal to fundamentals. For simplicity, we assume

approval shocks are independent of output growth.[8]

A key link between performance and the fiscal rates implemented at $t+1$ is

$$q_{t+1} = \bar{q} + q_\tau\big(\tau_{t+1} - \bar{\tau}\big) + q_x\big(x_{t+1} - \bar{x}\big), \qquad q_\tau < 0, \quad q_x < 0. \tag{15}$$

This reduced-form specification captures that higher taxes and greater rent extraction reduce household allocations and worsen observable outcomes, consistent with findings in Table 9.

Voters do not directly observe the decomposition of outlays into mandated public goods G and discretionary rents X. Instead, they infer performance from observable aggregates—realized output, household allocations, tax collections, and the market value of government liabilities—which embed the consequences of the term policy. Modeling strategic voting under hidden actions would introduce an additional informational layer and turn the political block into a dynamic game. We abstract from this complexity to isolate the mechanism of interest: PART as a state-dependent *sensitivity* of approval to fundamentals.

The strength of election discipline is governed by approval tightness. Since $\Lambda'(\ell) = \Lambda(\ell)\big(1 - \Lambda(\ell)\big)$, then, for any term-policy component $a_{t+1} \in \{\tau_{t+1}, x_{t+1}\}$,

$$\frac{\partial p_{t+1}}{\partial a_{t+1}} = \Lambda'(\ell_{t+1})\frac{\partial \ell_{t+1}}{\partial a_{t+1}} = \gamma\, d_{t+1}\, \frac{\partial q_{t+1}}{\partial a_{t+1}}, \qquad \frac{\partial q_{t+1}}{\partial \tau_{t+1}} = q_\tau, \quad \frac{\partial q_{t+1}}{\partial x_{t+1}} = q_x. \tag{16}$$

The marginal electoral effect of policy is scaled by d_{t+1}. Approval is most responsive at intermediate levels (high tightness) and least responsive near the bounds (low tightness). This state dependence links PART to equilibrium fiscal distortions and risk premia.

In the full model, the incumbent's optimal term policy (τ_{t+1}, x_{t+1}) depends on political conditions (PAR and PART) and on the intertemporal budget constraint. We begin with a steady-state characterization that fixes the political state at a constant approval level p and tightness $d(p) \equiv p(1-p)$. This conditional steady state can be interpreted as (i) the long-run average policy under persistent approval dynamics in Eq. (14), and (ii) a local optimality condition evaluated at

[8] Under this benchmark, the continuation term separates into an expected-survival component and a market-valuation component, $\mathbb{E}_t[m_{t+T}p_{t+T}V_{t+T}] = \mathbb{E}_t[p_{t+T}]\,\mathbb{E}_t[m_{t+T}V_{t+T}]$. More generally, $\mathbb{E}_t[m_{t+T}p_{t+T}V_{t+T}] = \mathbb{E}_t[p_{t+T}]\,\mathbb{E}_t[m_{t+T}V_{t+T}] + \mathrm{Cov}_t(p_{t+T}, m_{t+T}V_{t+T})$. A nonzero covariance term introduces an interaction between political risk and discount-rate risk. It would reinforce the same electoral-discipline mechanism but is not needed for our main results.

time-varying (p_t, b_t, g_t), which we later use for time-series implications.

Fixing p and restricting attention to stationary fiscal policy delivers tractability. Under i.i.d. raw output growth and time-invariant policy-induced risk, all term payoffs and discounted net surpluses scale with current raw output Y_t^r. In a stationary equilibrium, the government maintains a constant market debt-to-raw-output ratio and chooses constant policy:

$$\tau_t = \tau, \qquad x_t = x, \qquad b_t = b \equiv \frac{B_t}{Y_t^r} \text{ constant.}$$

Under these restrictions, the value function inherits balanced-growth homogeneity,

$$V(B_t, Y_t^r, p) = Y_t^r \, v(b, p), \qquad b \equiv \frac{B_t}{Y_t^r},$$

so the problem reduces to stationary objects $(\tau, x, b, v(b,p), \partial v(b,p)/\partial b)$. The conditional steady state therefore delivers a mapping $(p, b) \mapsto (\tau, x)$, which we later evaluate at observed (p_t, b_t) to obtain testable time-series implications for next-term policy (τ_{t+1}, x_{t+1}).

Proposition 1 (Steady-State System)**.** *Under the T-period election timing and the conditional steady state with constant approval inputs $\big(p, d(p)\big)$ and constant debt ratio $b = B/Y^r$, the stationary policy (τ, x) and normalized value $v(b,p)$ solve*

$$c^\tau \tau = -p \, \frac{\partial v(b,p)}{\partial b} + \Gamma_T \, d(p) \, v(b,p) \, q_\tau \, \Psi_T(x), \tag{17}$$

$$\kappa\big(1 - \lambda x \, \Phi_T(x)\big) = -p \, \frac{\partial v(b,p)}{\partial b} - \Gamma_T \, d(p) \, v(b,p) \, q_x \, \Psi_T(x), \tag{18}$$

$$\tau - g - x = b\big(E(x) - 1\big), \tag{19}$$

$$v(b,p)\Big(1 - p \, E(x)^{-T}\Big) = \big(1 - \theta(\tau) - g + \kappa x\big) \, A_T(x), \qquad \theta(\tau) = \frac{1}{2} c^\tau \tau^2, \tag{20}$$

where $d(p) \equiv p(1-p)$, $q_\tau < 0$, $q_x < 0$, and $\Gamma_T \equiv \gamma(1 - \phi^T)/(1 - \phi)$. The net discount rate is

$$E(x) \equiv e^{\delta(x)} = \exp(\delta_0 + \lambda x), \qquad \lambda = \eta\sigma^2, \qquad \delta_0 = r + \lambda - \mu_y,$$

and $A_T(x)$, $\Psi_T(x)$, and $\Phi_T(x)$ are discounting objects determined by $E(x)$ and T.

Proposition 1 isolates two political channels that enter the marginal conditions differently. The

approval level p scales the continuation motive through $p\,\partial v(b,p)/\partial b$, the shadow value of fiscal space conditional on political survival. Dispersion d scales the electoral-performance channel: when approval is fragile (high d), policy distortions have larger marginal effects on survival, tightening discipline and shifting the optimal mix between taxes and rents.

To highlight the role of approval sensitivity, we shut down this channel by setting $\gamma = 0$ and show that optimal policy is then independent of approval.

Proposition 2 (No Sensitivity Implies No Re-election Effects)**.** *When $\gamma = 0$, approval is insensitive to performance in Eq. (14)–Eq. (15), so fiscal choices do not depend on expected approval conditions:*

$$\frac{\partial x}{\partial p} = 0, \qquad \frac{\partial \tau}{\partial p} = 0.$$

Moreover, the optimal (τ, x) coincides with the solution under certain political survival (i.e., $p \equiv 1$, hence $d \equiv 0$).

In other words, without the performance-to-approval linkage, re-election does not discipline rent seeking.

Approval sensitivity, summarized by PART, is therefore central for determining inefficient spending x. This feature helps rationalize the nonlinear relation between political conditions and asset returns. We illustrate this point in the following propositions:

Proposition 3 (Convexity of Rent Extraction in PAR)**.** *When $q_\tau + q_x < 0$, the optimal rent share x is a convex function of approval p.*

Convexity implies that rent seeking is lowest at intermediate approval levels and higher near the extremes. The mechanism is the concavity of approval sensitivity $d(p)$: when tightness is high, marginal improvements in performance have larger electoral returns, increasing the incentive to restrain rent extraction.

Moreover, changes in x feed into economic fundamentals:

Proposition 4 (Rent Seeking Lowers Output, Consumption, and Dividends)**.** *When $\kappa^2\lambda \le c^\tau c^x$,*

$$\frac{\partial}{\partial x}\Big(\frac{Y_t(x)}{Y_t^r}\Big) < 0, \quad \frac{\partial}{\partial x}\Big(\frac{C_t(x)}{Y_t^r}\Big) < 0, \quad \frac{\partial}{\partial x}\Big(\frac{D_t(x)}{Y_t^r}\Big) < 0 \quad \textit{for all } x \in (0, 1/\lambda],$$

so higher rent extraction strictly lowers the shares of final output, household consumption, and dividends in raw output.

Proposition 4 links PART-driven fiscal discipline to real allocations. To visualize these implications, we fix the debt-to-output ratio b at its historical average and solve the stationary system in Proposition 1 on a grid of approval levels p. Parameters are chosen to match average tax and spending shares in the data (details in Appendix A).

Figure 5 reports comparative statics for rent extraction x, policy-induced volatility $\sigma_y(x)$, and the implied shares of GDP and consumption in pre-allocation resources. Consistent with Proposition 3, $x(p)$ is convex in PAR and reaches its minimum around intermediate approval levels, where tightness is highest and electoral discipline is strongest. The green dashed line is the benevolent benchmark ($x \equiv 0$). The red dashed line is the no-re-election benchmark, in which electoral discipline is shut down and the government chooses a high and essentially constant rent share. The equilibrium (black) lies between these benchmarks: when d is high, rent extraction is pushed toward the benevolent benchmark; when d is low (very high or very low PAR), rent extraction approaches the unconstrained benchmark. The top-right panel shows the uncertainty implication. Because rent extraction raises policy-induced risk through $\sigma_y(x)$ in Eq. (6), the same discipline channel that restrains x also reduces volatility. The bottom panels translate these policy patterns into real allocations. GDP and consumption peak around intermediate approval levels—precisely when tightness is high and rent extraction is most disciplined—and fall back toward the low-output, high-rent benchmark as approval moves toward the bounds. Together, these comparative statics summarize the mechanism linking PART to real outcomes and risk premia.

4.4 Linkage to Predictability Results

For empirical purposes, we interpret Proposition 1 as a *local* mapping from state variables to next-term fiscal policy. At date t, the government observes debt-to-output b_t, mandated spending g_t, and political conditions (p_t), and chooses a term policy (τ_{t+1}, x_{t+1}) that is held fixed over $t+1, \ldots, t+T$. The comparative static in Figure 5 therefore illustrates how current PAR and PART forecast the term policy (τ_{t+1}, x_{t+1}) and, through Eq. (8)–Eq. (9), forecast consumption and dividend growth. In turn, expected asset returns are pinned down by the pricing kernel

Eq. (10). Following the same logic, we also analyze how changes in the fiscal state—debt-to-GDP b and mandated spending g (often found counter-cyclical)—shift optimal fiscal choices and hence fundamentals and asset prices (see Appendix A). Overall, the model links predictability to the channel through which initial conditions (p_t, b_t, g_t) map into next-term policy wedges (τ_{t+1}, x_{t+1}), which then propagate to future growth and asset pricing.

We summarize the key prediction as follows.

Proposition 5 (Cash-Flow and Return Responses)**.** *Consider a deviation of next-term fiscal policy* (τ_{t+1}, x_{t+1}) *from its long-run steady-state values* $(\bar{\tau}, \bar{x})$*. Define*

$$\Delta x_{t+1} \equiv x_{t+1} - \bar{x}, \qquad \Delta \tau_{t+1} \equiv \tau_{t+1} - \bar{\tau}, \qquad \Delta \theta_{t+1} \equiv \theta(\tau_{t+1}) - \theta(\bar{\tau}),$$

and let $\Delta \log Z_{t+1} \equiv \log(Z_{t+1}/\bar{Z})$ *denote log deviations from steady state for any aggregate* $Z \in \{Y, Y^r, C, D\}$*. A first-order log-linearization of Eq. (8)–Eq. (9) implies the cash-flow responses*

$$\Delta \log Y_{t+1} = \Delta \log Y^r_{t+1} - c^x \, \Delta x_{t+1} - \Delta \theta_{t+1}, \tag{21}$$

$$\Delta \log C_{t+1} = \Delta \log Y^r_{t+1} - (c^x + 1) \, \Delta x_{t+1} - \Delta \theta_{t+1}, \tag{22}$$

$$\Delta \log D_{t+1} = \Delta \log Y^r_{t+1} - c^x \, \Delta x_{t+1} - \Delta \theta_{t+1} - \Delta \tau_{t+1}. \tag{23}$$

The corresponding equity-return response obeys

$$r^E_{t+1} = \Delta \log D_{t+1} + \delta(\bar{x}) - \frac{\rho}{1-\rho} \, \lambda \, \Delta x_{t+1}, \tag{24}$$

where $\rho \equiv \frac{1}{1+\bar{D}/\bar{P}}$ *is a near-one constant pinned down by the average dividend–price ratio, and* $\lambda = \eta\sigma^2$ *is the policy-to-discount-rate slope in* $\delta(x) = \delta_0 + \lambda x$*.*

Proposition 5 provides a convenient decomposition from policy deviations to fundamentals and returns that is directly usable for predictability. A rise in rent extraction ($\Delta x_{t+1} > 0$) lowers output, consumption, and dividends through the resource wedges in Eq. (8)–Eq. (9). At the same time, it increases required compensation by raising the discount rate through $\delta(x)$, captured by the $\lambda \, \Delta x_{t+1}$ term. Movements in τ_{t+1} further depress dividends through the $-\Delta \tau_{t+1}$ term.

Within our framework, the government must finance spending according to

$$\tau(x) = g + x + b\big(E(x) - 1\big), \qquad E(x) = e^{\delta(x)}.$$

Debt financing is therefore costly because the primary surplus is effectively exposed to output risk (see also Jiang et al. (2024a, 2026)). Since $E(x)$ increases with x through the policy-induced discount-rate channel, a higher x implies not only larger discretionary outlays but also higher refinancing pressure, and thus higher taxes and a larger primary surplus share in steady state. As a result, dividends are negatively affected by x both through the cash-flow wedge and through the higher discount rate. Taken together, these channels imply lower equity valuations and, through Eq. (24), lower (log-linearized) equity returns when Δx_{t+1} rises. Consequently, political states in which PART predicts a reduction in rent extraction x translate into higher subsequent dividend growth and higher equity returns in the model. Our framework also implies bond-return predictability, yet this result is more involved with other confounding factors on debt valuation and we provide a detailed discussion in Appendix B.

To connect these objects to PAR and PART, we use the steady-state mapping in Proposition 1 together with its local comparative statics. Holding the non-political conditions fixed, the optimal next-term rent choice admits a first-order representation around the long-run mean political state:

$$\Delta x_{t+1} = x_{t+1} - \bar{x} = A_p\big(\mathrm{PAR}_t - \overline{\mathrm{PAR}}\big) + A_d\big(\mathrm{PART}_t - \overline{\mathrm{PART}}\big), \tag{25}$$

where $\overline{\mathrm{PAR}}$ and $\overline{\mathrm{PART}}$ denote long-run (sample) means, and A_p, A_d are local slopes implied by the steady-state system. The same mapping implies that $\Delta\tau_{t+1}$ and $\Delta\theta_{t+1}$ are likewise spanned by $\mathrm{PAR}_t - \overline{\mathrm{PAR}}$ and $\mathrm{PART}_t - \overline{\mathrm{PART}}$. Substituting these relationships into Proposition 5 yields a tight theoretical interpretation of predictability regressions: PAR and PART forecast future growth and returns because they forecast next-term policy wedges relative to their long-run benchmarks.

In practice, political conditions co-move with other slow-moving fundamentals, so predictability reflects the joint influence of political and macro–finance channels. Our empirical strategy therefore treats PAR and PART as political state variables that complement canonical predictors for each outcome. Guided by the model's first-order conditions—which distinguish level effects (through p_t)

from sensitivity effects (through d_t)—we estimate reduced-form regressions of the form

$$y_{t+h} = \alpha_h + \beta_h^{\text{PAR}} \, \text{PAR}_t + \beta_h^{\text{PART}} \, \text{PART}_t + \Gamma_h' Z_t + \varepsilon_{t+h},$$

where Z_t collects standard controls such as valuation ratios, macro variables, and fiscal controls such as debt-to-GDP. Including both PAR and PART is essential for separating approval-level effects from approval-sensitivity effects in the data, consistent with the mechanism in which electoral discipline depends not only on the level of approval but also on the tightness/sensitivity of political support.

5 Conclusion

This paper uses presidential job approval as a real-time measure of political support and traces how it feeds into U.S. macroeconomic outcomes, fiscal behavior, and asset prices. Our central object is the presidential approval rate (PAR) and a derived measure of approval tightness that is highest when support is at a knife edge, that is, when approval is around 50 percent. Taken together, our evidence shows that political support is not just background noise for the economy, but a state variable that helps organize variation in growth, risk premia, and uncertainty.

Empirically, we document pronounced nonlinear links between approval and both macro and financial outcomes. Future real output and consumption growth and the equity risk premium are highest when presidential approval is in an intermediate range and lowest when it is either very high or very low, whereas macroeconomic and financial uncertainty display the opposite U-shaped pattern. We summarize these regularities with approval tightness, $\text{PART} = \text{PAR}(1 - \text{PAR})$, which peaks when electoral competition is tight. This simple transformation captures the hump-shaped relation between approval and growth/returns and the mirror-image pattern for volatility, and it delivers forecasting power for these variables over and above standard predictors and measures of belief or policy uncertainty.

We then show that approval tightness has a natural interpretation as the sensitivity of political support to fundamentals. Modeling presidential approval as a probability-like object, we estimate logit and probit specifications in which real activity, labor-market conditions, inflation, fiscal vari-

ables, and major political events feed into approval. In this framework, the derivative of approval with respect to fundamentals is exactly proportional to approval tightness and is empirically steepest at intermediate approval levels. Moreover, periods with high approval tightness are followed by lower federal spending growth and better scores on perception-based indicators of control of corruption and government capacity. These results support viewing approval tightness as an empirical measure of electoral responsiveness that tightens fiscal discipline and strengthens state capacity.

To interpret these patterns in a unified way, we develop a structural political-economy model in which a government can extract rents through inefficient fiscal spending that both wastes resources and raises macroeconomic risk. Elections discipline this behavior because policy choices feed into approval, which in turn determines the probability of remaining in office. The strength of this discipline is governed by how sensitive approval is to fundamentals—captured in reduced form by approval tightness. In equilibrium, rent extraction and political uncertainty are U-shaped in approval, while real activity, private-sector cash flows, and expected returns are hump-shaped, mirroring our empirical findings for growth, the equity premium, and volatility.

Overall, the paper highlights approval tightness as a simple, observable proxy for the intensity of electoral discipline, with clear implications for fiscal policy, governance outcomes, and risk pricing. As a potential direction for future research, our framework can be extended to study interactions between political cycles and fiscal risk cycles, government debt management, and related issues. On the empirical side, a natural next step is to construct analogous measures of approval tightness in other democracies and to embed approval-sensitive electoral discipline into richer quantitative models of the joint behavior of macroeconomic and financial variables.

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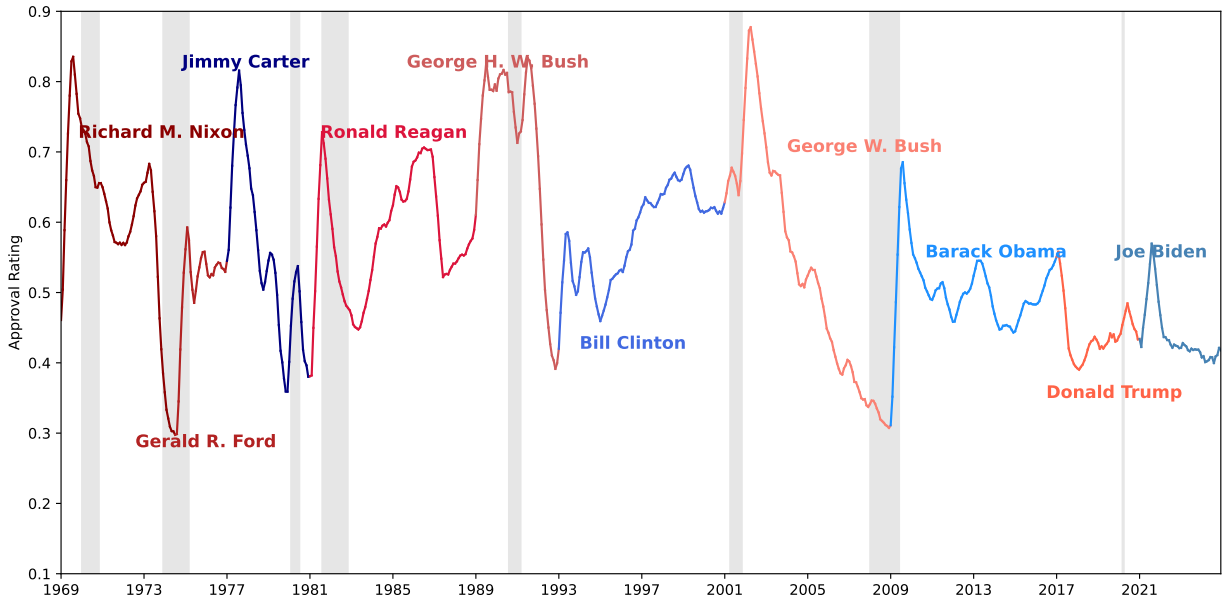


FIGURE 1: **Presidential Approval Rating.** This figure plots the time series of Gallup's Presidential Approval Rating (PAR), smoothed using a six-month moving average. Shaded bands denote NBER-dated recession periods. The sample spans January 1969 to December 2024. Colored segments correspond to presidential administrations and party affiliation. Warm tones denoting Republican presidents and cool tones denoting Democratic presidents.

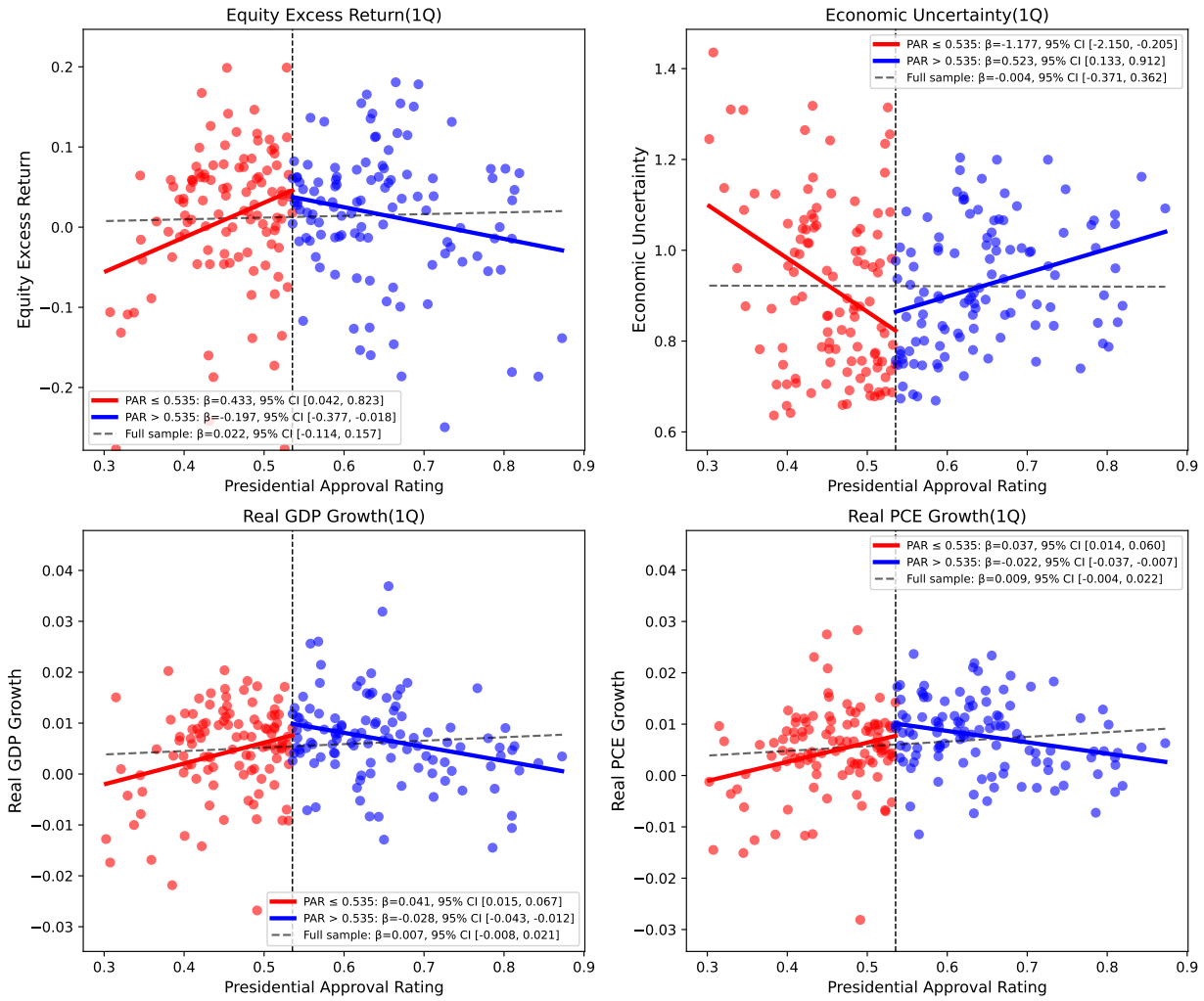


FIGURE 2: **Median-Threshold Predictability of Presidential Approval Rating.** Each panel plots a one-quarter-ahead outcome against the presidential approval rating PAR_t. The sample is split at the median of PAR_t (vertical dashed line): observations with PAR_t below (above) the median are shown in red (blue). Solid lines report within-subsample OLS fits, and the dashed gray line reports the full-sample linear fit. The legend reports slope estimates and 95% confidence intervals based on Newey–West standard errors. Panels show equity excess returns, economic uncertainty as measured by Jurado et al. (2015), real GDP growth, and real PCE growth.

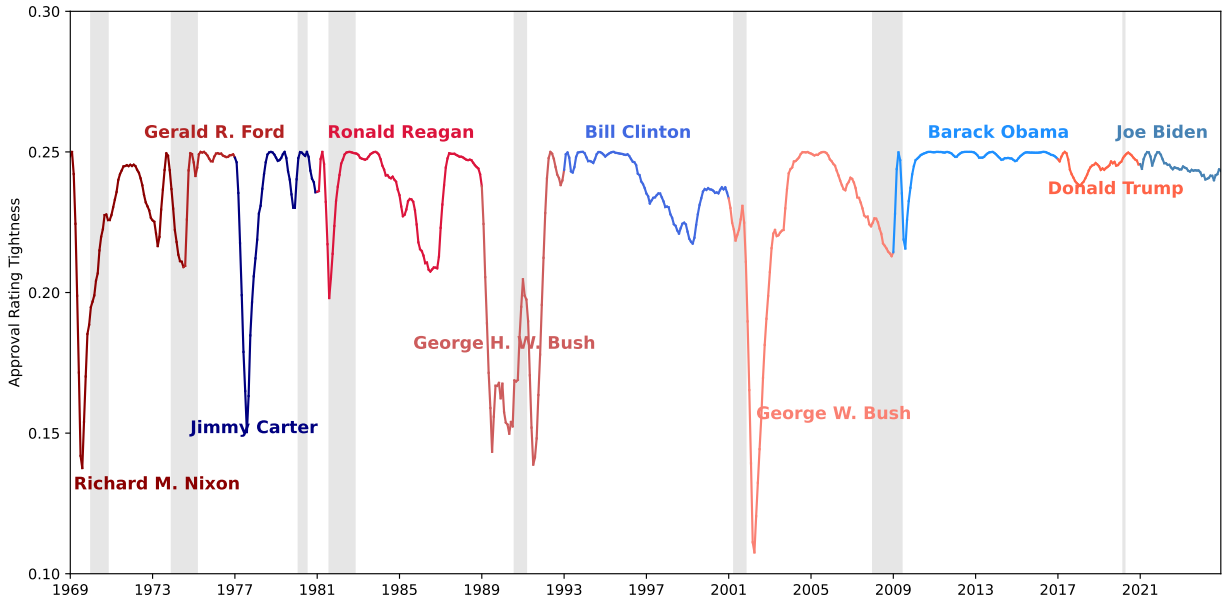


FIGURE 3: **Presidential Approval Rating Tightness.** This figure plots the time series of presidential approval rating tightness,

$$PART_t = PAR_t\big(1 - PAR_t\big).$$

, over January 1969 to December 2024. Shaded bands denote NBER-dated recession periods. Colored segments correspond to presidential administrations and party affiliation. Warm tones denoting Republican presidents and cool tones denoting Democratic presidents.

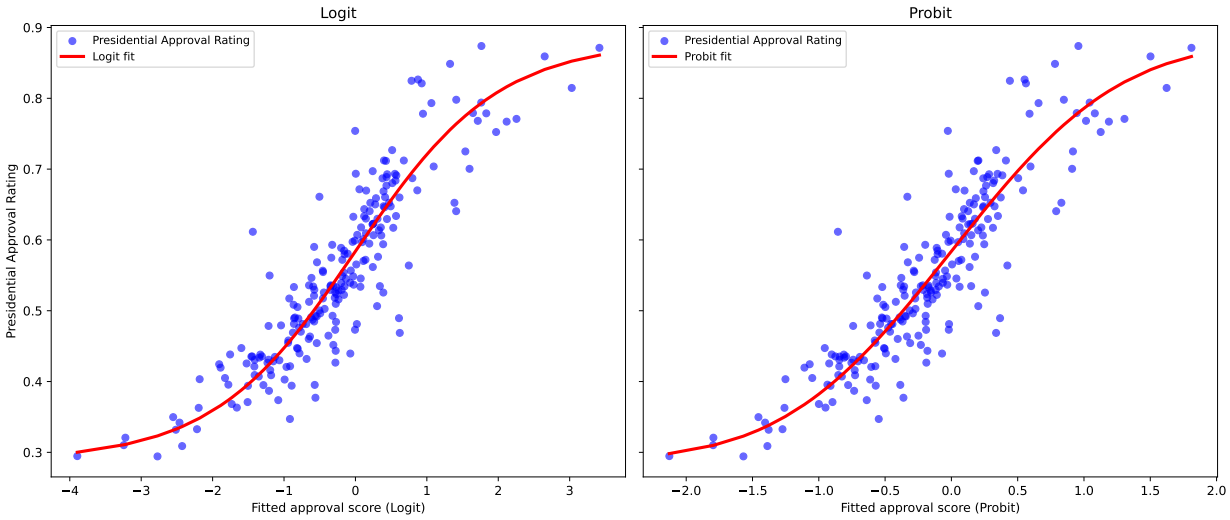


FIGURE 4: **Response of the Presidential Approval Rate to Economic Performance.** This figure relates observed presidential approval to fitted values of the latent approval index under two alternative link functions. The left panel uses a scaled logit specification, $PAR_t = a + b\,\sigma(\ell_t)$ with $\sigma(x) = 1/(1+e^{-x})$. The right panel uses a scaled probit specification, $PAR_t = a + b\,\Phi(\ell_t)$, where $\Phi(\cdot)$ is the standard normal cumulative distribution function. In each panel, the horizontal axis reports the fitted economic performance index $\widehat{\ell_t}$ from the linear regression

$$\ell_t = \alpha + \beta\,\ell_{t-1} + \delta^{\top}\text{Politics}_{t-1} + \gamma\,X_t + \varepsilon_{t+1},$$

where X_t includes measures of economic performance, namely real GDP growth, real PCE growth, the unemployment rate, $\log(\text{Tax/GDP})_t$, $\log(\text{Spending/GDP})_t$, and inflation, and Politics_{t-1} collects indicators for major political events. Blue markers plot the realized approval rate PAR_t against the fitted index, and the red curve overlays the implied mapping from the fitted index to approval under the corresponding link function.

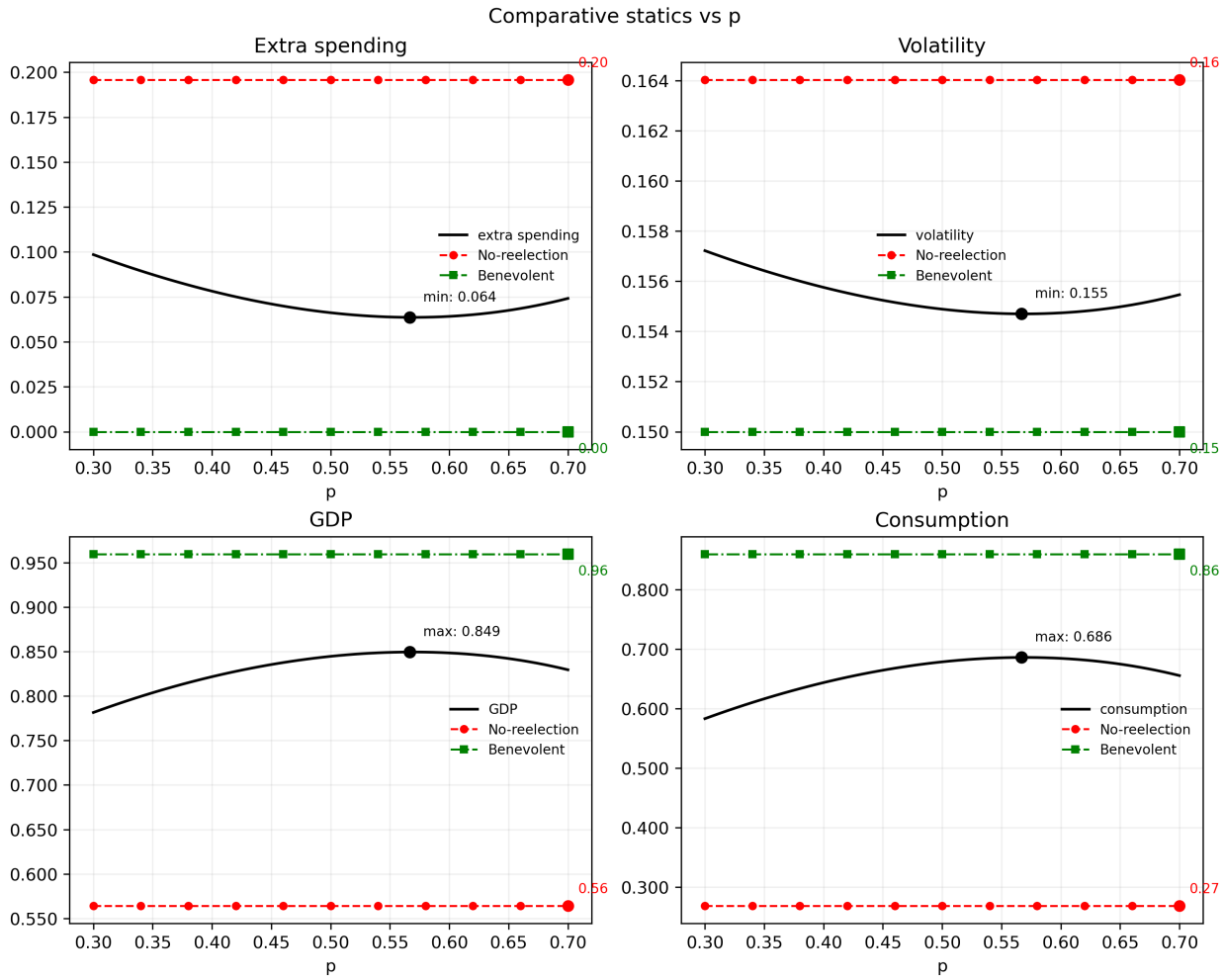


FIGURE 5: **Rent Extraction and Real Allocations Across Presidential Approval Levels.** This plot shows how the model-implied optimal rent extraction x and associated macroeconomic outcomes vary with the presidential approval rate p, as characterized in Proposition 1. The top-left panel plots equilibrium rent extraction $x(p)$ (black) against the benevolent benchmark $x \equiv 0$ (green dashed) and the no–re-election benchmark (red dashed). The top-right panel reports the implied policy-induced volatility $\sigma_y(x(p))$ as defined in Eq. (6). The bottom panels translate these policy choices into the implied shares of GDP and consumption in pre-allocation resources defined by equation Eq. (5).

TABLE 1: **Summary Statistics of Predictors, Returns, and Growth Rates.** This table reports summary statistics for the variables used in the empirical analysis. Panel A reports predictors, and Panel B reports the log excess equity return (r^E) and log growth rates of macroeconomic variables—ΔGDP for real GDP, ΔPCE for real personal consumption expenditures, and ΔG for federal government spending. Summary statistics are reported at each series' native frequency (monthly for PAR and $PART$, quarterly otherwise). Mean and standard deviation are annualized for returns and growth rates only. Correlations are computed using quarterly-aligned series. Panel A also reports Augmented Dickey–Fuller (ADF), Phillips–Perron (PP), and Dickey–Fuller GLS (DF–GLS) unit-root tests with associated p-values; significance stars indicate rejection of the unit-root null at the 1%, 5%, and 10% levels. The sample spans 1969–2024.

	Mean, Std, AR1, Stationarity Tests					Correlation				
Panel A: Predictors										
	PAR	$PART$	dp	log(Debt/GDP)	nsb		$PART$	dp	log(Debt/GDP)	nsb
Mean	0.550	0.233	-3.696	-0.968	-0.019	PAR	-0.698	0.113	-0.324	0.329
Std	0.121	0.025	0.402	0.508	0.057	$PARD$		-0.095	0.262	-0.266
AR1	0.987	0.975	0.974	0.997	0.889	dp			-0.600	0.013
ADF	-4.72***	-5.48***	-1.46	-0.81	-3.28**	log(Debt/GDP)				-0.463
p-val	0.000	0.000	0.552	0.815	0.016					
PP	-3.95***	-4.77***	-1.44	-0.21	-3.48***					
p-val	0.002	0.000	0.563	0.937	0.009					
DF_GLS	-3.95***	-4.75***	-1.25	0.15	-2.47**					
p-val	0.000	0.000	0.199	0.743	0.014					
Panel B: Equity Returns and Cash Flow Growth										
	r^E	ΔGDP	ΔPCE	ΔG			ΔGDP	ΔPCE	ΔG	
Mean	5.196	2.212	2.460	6.412		r^E	0.059	0.065	0.032	
Std	17.620	2.278	2.258	11.990		ΔGDP		0.861	-0.476	
AR1	0.033	-0.011	-0.072	-0.256		ΔPCE			-0.511	
ADF	-6.93***	-5.95***	-5.66***	-6.25***						
p-val	0.000	0.000	0.000	0.000						
PP	-14.39***	-15.10***	-15.96***	-20.05***						
p-val	0.000	0.000	0.000	0.000						
DF_GLS	-3.63***	-4.22***	-5.66***	-3.98***						
p-val	0.000	0.000	0.000	0.000						

TABLE 2: **PAR and Economic Growth**: This table reports predictive regressions of real output and consumption growth on presidential approval (PAR_t), presidential approval tightness ($PART_t$), and related controls. The dependent variables are real GDP growth and real PCE growth, reported in the left and right blocks, respectively. Panels differ by sampling frequency and forecast horizon. Panel A uses quarterly data and forecasts one-quarter-ahead growth. Panel B uses quarterly data and forecasts cumulative four-quarter growth. Panel C uses annual data and forecasts one-year-ahead growth. Columns (1) and (2) report univariate specifications that include PAR_t or $PART_t$ separately, column (3) includes both jointly, and column (4) adds controls, including lagged growth and survey expectations from the Survey of Professional Forecasters. The expectations control is horizon-matched: when forecasting one-quarter-ahead growth, we use the SPF expectation of next-quarter growth; when forecasting cumulative four-quarter growth, we use the SPF expectation of four-quarter growth; and when forecasting one-year-ahead growth in annual data, we use the SPF expectation of next-year growth. Newey-West (HAC) t-statistics are reported in parentheses using four lags. Significance stars denote statistical significance at the 1%, 5%, and 10% levels.

Panel A: Quarterly, h=1	Y: GDP Growth					Y: PCE Growth			
	(1)	(2)	(3)	(4)		(1)	(2)	(3)	(4)
PAR_t	0.01		0.03***	0.01	PAR_t	0.01		0.03***	0.02**
	(0.94)		(4.41)	(1.43)		(1.39)		(4.91)	(2.49)
$PART_t$		0.06***	0.15***	0.07**	$PART_t$		0.04*	0.13***	0.05*
		(2.63)	(5.7)	(2.35)			(1.8)	(5.43)	(1.92)
$E_t[GDPG_{t+1}]$				1.29***	$E_t[PCEG_{t+1}]$				1.68***
				(5.32)					(5.88)
$GDPG_t$				-0.17**	$PCEG_t$				-0.11*
				(-2.12)					(-1.78)
$R^2_{adj}(\%)$	0.09	1.01	5.46	26.22	$R^2_{adj}(\%)$	0.5	0.19	4.37	28.06
Panel B: Quarterly, h=4	Y: GDP Growth					Y: PCE Growth			
PAR_t	0.03		0.11***	0.08***	PAR_t	0.04		0.1***	0.09***
	(1.1)		(5.72)	(3.58)		(1.58)		(5.63)	(4.17)
$PART_t$		0.19**	0.55***	0.41***	$PART_t$		0.12	0.47***	0.38***
		(2.14)	(5.51)	(4.21)			(1.56)	(4.89)	(4.05)
$E_t[GDPG_{t:t+4}]$				1.01***	$E_t[PCEG_{t:t+4}]$				1.24***
				(4.54)					(3.6)
$GDPG_t$				-0.13	$PCEG_t$				-0.11
				(-0.96)					(-0.98)
$R^2_{adj}(\%)$	1.51	3.23	17.67	36.73	$R^2_{adj}(\%)$	2.97	1.28	16	38.01
Panel C: Annual, h=1	Y: GDP Growth					Y: PCE Growth			
PAR_t	0.02		0.12***	0.08***	PAR_t	0.03		0.12***	0.08***
	(0.79)		(5.51)	(4.14)		(1.18)		(5.05)	(3.91)
$PART_t$		0.19**	0.65***	0.38***	$PART_t$		0.12	0.56***	0.36***
		(2.16)	(5.68)	(4.37)			(1.57)	(5.0)	(2.94)
$E_t[GDPG_{t+1}]$				0.95***	$E_t[PCEG_{t+1}]$				0.98***
				(3.29)					(3.68)
$GDPG_t$				-0.04	$PCEG_t$				-0.0
				(-0.54)					(-0.04)
$R^2_{adj}(\%)$	-0.19	2.01	18.57	47.81	$R^2_{adj}(\%)$	1.53	-0.17	16.94	40.05

TABLE 3: **PAR and Equity Returns**: This table reports predictive regressions of equity excess returns on presidential approval (PAR_t), approval tightness ($PART_t$), and controls. The forecasting specification is

$$R_{t:t+h} = \alpha + \beta^{PAR} PAR_t + \beta^{PART} PART_t + \beta^{Control} Control_t + \varepsilon_{t:t+h},$$

where $R_{t:t+h}$ denotes the cumulative equity excess return from t to $t+h$. Panel A reports monthly regressions and Panel B reports quarterly regressions. In each panel, the left block uses the one-period horizon and the right block uses the longer cumulative horizon, with h = 12 months in Panel A and h = 4 quarters in Panel B. Columns (1) and (2) report univariate specifications that include PAR_t or $PART_t$ as the sole predictor. Column (3) includes both variables jointly. The final column adds the full control set, including the standard return-predictor controls of Welch and Goyal (2008), especially the dividend-to-price ratio dp_t, and macro variables following Baker and Wurgler (2006). The row labeled *Controls* indicates whether this full control set is included. Newey-West (HAC) t-statistics are reported in parentheses, using six lags for the monthly one-period horizon, twelve lags for the monthly 12-month horizon, and four lags for quarterly horizons. Significance stars denote rejection at the 1%, 5%, and 10% levels.

Panel A: Monthly	h=1				h=12			
	(1)	(2)	(3)	(4)	(1)	(2)	(3)	(4)
PAR_t	0.0		0.07**	0.1***	-0.02		0.28	0.51**
	(0.2)		(2.56)	(3.37)	(-0.14)		(1.16)	(2.15)
$PART_t$		0.22***	0.46***	0.52***		1.14**	2.08**	2.64***
		(3.28)	(3.81)	(4.15)		(2.35)	(2.2)	(2.78)
dp_t				0.03***				0.27***
				(3.0)				(2.62)
Controls:	N	N	N	Y	N	N	N	Y
R^2_{adj}(%)	-0.14	1.30	2.90	5.53	-0.13	2.75	4.69	22.24
Panel B: Quarterly	h=1				h=4			
	(1)	(2)	(3)	(4)	(1)	(2)	(3)	(4)
PAR_t	0.02		0.23***	0.39***	-0.02		0.29	0.58**
	(0.31)		(2.83)	(4.0)	(-0.14)		(1.2)	(2.5)
$PART_t$		0.67***	1.45***	1.96***		1.18**	2.16**	2.88***
		(2.73)	(3.77)	(4.71)		(2.32)	(2.25)	(3.11)
dp_t				0.11***				0.25**
				(2.94)				(2.3)
Controls:	N	N	N	Y	N	N	N	Y
R^2_{adj}(%)	-0.36	3.14	7.85	15.24	-0.43	2.37	3.98	21.6

TABLE 4: **Presidential Approval and Equity Returns: Controlling for Alternative Predictors**: This table examines whether presidential approval (PAR_t) and approval tightness ($PART_t$) predict equity excess returns after controlling for commonly used alternative predictors. We estimate

$$R_{t:t+h} = \alpha + \beta^{PAR} PAR_t + \beta^{PART} PART_t + \beta^{Controls} Controls_t + \varepsilon_{t:t+h},$$

where $R_{t:t+h}$ denotes the cumulative equity excess return between t and $t+h$ (Panel A: h = 1 month; Panel B: h = 1 quarter). Each row augments the baseline specification with the control variable(s) listed in the *Controls* column. *Disagreement* is the economic belief disagreement index of Huang et al. (2021) and *EPU* is the Economic Policy Uncertainty index of Baker et al. (2016). Both series are standardized at the monthly frequency. *Party* is an indicator equal to one if the incumbent president is a Democrat and zero otherwise. $Midterm^{Before}$, $Midterm^{After}$, $Election^{Before}$, and $Election^{After}$ are indicators equal to one during the five months immediately preceding or following the corresponding midterm or presidential election and zero otherwise. *Honeymoon* is an indicator equal to one during the first twelve months of a new presidency and zero otherwise. The row labeled *All* includes all controls jointly. All control variables are constructed at the monthly frequency. For Panel B, quarterly values of *Disagreement* and *EPU* are computed as quarterly averages of the monthly series, and quarterly values of the dummy variables equal one if any of the three months in the quarter takes the value one and zero otherwise. The sample spans from January 1970 to December 2020 for *Disagreement* and from January 1969 to December 2024 for all other variables. Newey–West (HAC) t-statistics are reported in parentheses, computed with six lags for monthly regressions and four lags for quarterly regressions. Significance stars denote rejection of the null hypothesis at the 1%, 5%, and 10% levels.

Panel A:	1 Month						
Controls	β^{PAR}	(t-stat)	β^{PART}	(t-stat)	$\beta^{Controls}$	(t-stat)	R^2_{adj}(%)
/	0.07**	(2.56)	0.46***	(3.81)			2.9
Disagreement	0.07***	(2.72)	0.42***	(3.69)	-0.01***	(-3.05)	4.43
EPU	0.08***	(3.11)	0.48***	(3.94)	0.01***	(3.62)	4.43
Party	0.07**	(2.53)	0.43***	(3.62)	0.0	(1.13)	2.91
$Midterm^{Before}$	0.07***	(2.59)	0.47***	(3.82)	0.0	(0.47)	2.79
$Midterm^{After}$	0.07***	(2.63)	0.45***	(3.83)	0.02***	(4.3)	4.22
$Election^{Before}$	0.07***	(2.6)	0.46***	(3.76)	-0.0	(-0.34)	2.77
$Election^{After}$	0.07***	(2.58)	0.46***	(3.83)	-0.0	(-0.3)	2.76
Honeymoon	0.07**	(2.57)	0.46***	(3.76)	0.0	(0.07)	2.75
All	0.08***	(3.11)	0.43***	(3.48)			6.91
Panel B:	1 Quarter						
Controls	β^{PAR}	(t-stat)	β^{PART}	(t-stat)	$\beta^{Controls}$	(t-stat)	R^2_{adj}(%)
/	0.23***	(2.83)	1.45***	(3.77)			7.85
Disagreement	0.23***	(3.09)	1.39***	(3.81)	-0.01**	(-2.01)	9.87
EPU	0.28***	(3.33)	1.51***	(3.95)	0.02***	(4.72)	12.47
Party	0.23***	(2.79)	1.39***	(3.6)	0.01	(0.76)	7.63
$Midterm^{Before}$	0.24***	(2.99)	1.53***	(3.94)	0.03**	(2.36)	9.26
$Midterm^{After}$	0.23***	(2.98)	1.43***	(3.92)	0.02*	(1.85)	8.68
$Election^{Before}$	0.23***	(2.85)	1.47***	(3.83)	-0.01	(-0.76)	7.7
$Election^{After}$	0.24***	(3.0)	1.48***	(3.89)	-0.01	(-0.8)	7.71
Honeymoon	0.24***	(2.99)	1.41***	(3.46)	-0.02	(-0.86)	7.96
All	0.29***	(3.99)	1.55***	(3.84)			15.06

TABLE 5: **Presidential Approval and Volatility Indices**: This table reports predictive regressions of volatility indices on presidential approval (PAR_t), presidential approval tightness ($PART_t$), and lagged volatility. Each block uses a different volatility measure as the dependent variable. The proxies include the CBOE Volatility Index (VIX), constructed from S&P 500 index options; the Economic Policy Uncertainty (EPU) index of Baker et al. (2016); and the financial, macro, and real uncertainty measures of Jurado et al. (2015). For each proxy, the dependent variable is the average volatility level over the window from month t to month $t+h$, with horizons $h=1$, $h=6$, and $h=12$. All volatility proxies are standardized. The regressions include Y_t, the lagged value of the uncertainty proxy, to account for persistence. Newey-West (HAC) t-statistics are reported in parentheses. The HAC lag length equals 6 for horizons $h=1$ and $h=6$, and equals 12 for horizon $h=12$. Significance stars denote statistical significance at the 1%, 5%, and 10% levels.

	Y: VIX				Y: EPU		
	h=1	h=6	h=12		h=1	h=6	h=12
PAR_t	-0.54	-1.41	-1.67	PAR_t	-0.59**	-1.08**	-1.23**
	(-1.08)	(-1.27)	(-1.21)		(-2.47)	(-2.5)	(-2.03)
$PART_t$	-4.49**	-11.99**	-12.43**	$PART_t$	-1.45	-3.71**	-4.26*
	(-2.12)	(-2.53)	(-2.16)		(-1.46)	(-2.08)	(-1.78)
Y_t	0.83***	0.55***	0.42***	Y_t	0.78***	0.62***	0.54***
	(27.22)	(12.74)	(6.4)		(21.18)	(12.47)	(9.77)
$R^2_{adj}(\%)$	73.18	54.73	44.49	$R^2_{adj}(\%)$	62.84	55.96	51.04
	Y: Financial Uncertainty				Y: Macro Uncertainty		
	h=1	h=6	h=12		h=1	h=6	h=12
PAR_t	-0.40**	-1.19**	-1.48*	PAR_t	-0.34**	-1.45***	-2.30***
	(-2.57)	(-2.33)	(-1.67)		(-2.18)	(-2.72)	(-2.61)
$PART_t$	-2.46***	-7.75***	-9.85***	$PART_t$	-1.28*	-5.71***	-8.32***
	(-3.74)	(-3.66)	(-2.65)		(-1.87)	(-2.79)	(-2.76)
Y_t	0.97***	0.83***	0.68***	Y_t	0.98***	0.83***	0.67***
	(75.22)	(23.6)	(12.66)		(48.69)	(19.89)	(12.13)
$R^2_{adj}(\%)$	96.33	81.28	65.52	$R^2_{adj}(\%)$	97.09	81.97	67.57
	Y: Real Uncertainty						
	h=1	h=6	h=12				
PAR_t	-0.34**	-1.48***	-2.28***				
	(-2.5)	(-3.18)	(-3.19)				
$PART_t$	-1.45**	-6.15***	-8.46***				
	(-2.56)	(-3.41)	(-3.58)				
Y_t	0.97***	0.82***	0.67***				
	(32.67)	(16.46)	(17.02)				
$R^2_{adj}(\%)$	95.82	78.37	65.35				

TABLE 6: **PAR and Equity Returns: Small-Sample Bias-Adjusted Estimates**: This table reports predictive regressions of equity excess returns on presidential approval (PAR_t), approval tightness ($PART_t$), and controls. Coefficient estimates for PAR_t and $PART_t$ are adjusted for small-sample bias following Amihud and Hurvich (2004). The forecasting specification is

$$R_{t:t+h} = \alpha + \beta^{PAR} PAR_t + \beta^{PART} PART_t + \beta^{Control} Control_t + \varepsilon_{t:t+h},$$

where $R_{t:t+h}$ denotes the cumulative equity excess return from t to $t+h$. Panel A reports monthly regressions and Panel B reports quarterly regressions. In each panel, the left block uses the one-period horizon and the right block uses the longer cumulative horizon, with h = 12 months in Panel A and h = 4 quarters in Panel B. Columns (1) and (2) report univariate specifications that include PAR_t or $PART_t$ as the sole predictor. Column (3) includes both variables jointly. The final column adds the full control set, including the standard return-predictor controls of Welch and Goyal (2008) and macro variables following Baker and Wurgler (2006). The row labeled *Controls* indicates whether this full control set is included. Reported t-statistics are based on HAC covariance estimates that use Newey-West lags of 6 for the monthly one-period horizon, 12 for the monthly 12-month horizon, and 4 for quarterly horizons, and incorporate the Amihud-Hurvich heuristic variance adjustment under a diagonal predictor-dynamics correction. Significance stars denote rejection at the 1%, 5%, and 10% levels.

Panel A: Monthly	h=1				h=12			
	(1)	(2)	(3)	(4)	(1)	(2)	(3)	(4)
PAR_t	0.00		0.07**	0.10***	-0.02		0.32	0.55**
	(0.19)		(2.53)	(3.34)	(-0.13)		(1.4)	(2.4)
$PART_t$		0.22***	0.46***	0.52***		1.17**	2.23**	2.80***
		(3.28)	(3.76)	(4.16)		(2.57)	(2.49)	(3.0)
dp_t				0.03***				0.26***
				(2.99)				(2.61)
Controls:	N	N	N	Y	N	N	N	Y
$R^2_{adj}(\%)$	-0.14	1.31	2.88	5.53	-0.13	2.74	4.71	22.24
Panel B: Quarterly	h=1				h=4			
	(1)	(2)	(3)	(4)	(1)	(2)	(3)	(4)
PAR_t	0.02		0.24***	0.39***	-0.02		0.41*	0.73***
	(0.29)		(2.7)	(3.86)	(-0.11)		(1.87)	(3.23)
$PART_t$		0.67***	1.47***	1.97***		1.22**	2.60***	3.44***
		(2.76)	(3.67)	(4.68)		(2.43)	(2.92)	(3.79)
dp_t				0.12***				0.26**
				(3.12)				(2.35)
Controls:	N	N	N	Y	N	N	N	Y
$R^2_{adj}(\%)$	-0.36	3.14	7.86	15.24	-0.44	2.34	4.05	21.6

TABLE 7: **PAR and Economic Growth: Small-Sample Bias-Adjusted Estimates**: This table reports predictive regressions of real output and consumption growth on presidential approval (PAR_t), presidential approval tightness ($PART_t$), and related controls. Coefficient estimates for PAR_t and $PART_t$ are adjusted for small-sample bias following Amihud and Hurvich (2004). The dependent variables are real GDP growth and real PCE growth, reported in the left and right blocks, respectively. Panels differ by sampling frequency and forecast horizon. Panel A uses quarterly data and forecasts one-quarter-ahead growth. Panel B uses quarterly data and forecasts cumulative four-quarter growth. Panel C uses annual data and forecasts one-year-ahead growth. Columns (1) and (2) report univariate specifications that include PAR_t or $PART_t$ separately, column (3) includes both jointly, and column (4) adds controls, including lagged growth and survey expectations from the Survey of Professional Forecasters. The expectations control is horizon-matched: when forecasting one-quarter-ahead growth, we use the SPF expectation of next-quarter growth; when forecasting cumulative four-quarter growth, we use the SPF expectation of four-quarter growth; and when forecasting one-year-ahead growth in annual data, we use the SPF expectation of next-year growth. Reported t-statistics are based on HAC covariance estimates that use Newey-West lags of four, and incorporate the Amihud-Hurvich heuristic variance adjustment under a diagonal predictor-dynamics correction. Significance stars denote statistical significance at the 1%, 5%, and 10% levels.

Panel A: Quarterly, h=1	Y: GDP Growth					Y: PCE Growth			
	(1)	(2)	(3)	(4)		(1)	(2)	(3)	(4)
PAR_t	0.01		0.03***	0.01	PAR_t	0.01		0.03***	0.01**
	(0.88)		(3.98)	(0.96)		(1.37)		(5.34)	(2.21)
$PART_t$		0.06**	0.15***	0.06*	$PART_t$		0.04*	0.13***	0.04
		(2.57)	(5.2)	(1.78)			(1.79)	(5.69)	(1.47)
$E_t[GDPG_{t+1}]$				1.30***	$E_t[PCEG_{t+1}]$				1.70***
				(5.36)					(6.39)
$GDPG_t$				-0.16**	$PCEG_t$				-0.11*
				(-2.03)					(-1.79)
$R^2_{adj}(\%)$	0.04	1.05	5.39	26.18	$R^2_{adj}(\%)$	0.51	0.19	4.41	28.16
Panel B: Quarterly, h=4	Y: GDP Growth					Y: PCE Growth			
PAR_t	0.03		0.12***	0.09***	PAR_t	0.04		0.11***	0.09***
	(1.1)		(6.16)	(3.96)		(1.58)		(6.19)	(4.39)
$PART_t$		0.19**	0.58***	0.44***	$PART_t$		0.12	0.50***	0.39***
		(2.25)	(6.14)	(4.72)			(1.62)	(5.5)	(4.27)
$E_t[GDPG_{t:t+4}]$				1.00***	$E_t[PCEG_{t:t+4}]$				1.26***
				(4.55)					(3.66)
$GDPG_t$				-0.15	$PCEG_t$				-0.11
				(-1.17)					(-0.96)
$R^2_{adj}(\%)$	1.54	3.19	17.72	37.05	$R^2_{adj}(\%)$	3.06	1.25	16.2	38.88
Panel C: Annual, h=1	Y: GDP Growth					Y: PCE Growth			
PAR_t	0.01		0.11***	0.08***	PAR_t	0.03		0.12***	0.09***
	(0.34)		(5.68)	(4.55)		(0.9)		(5.21)	(3.66)
$PART_t$		0.25***	0.66***	0.38***	$PART_t$		0.16**	0.61***	0.37***
		(3.08)	(7.05)	(3.97)			(1.98)	(5.15)	(2.73)
$E_t[GDPG_{t+1}]$				0.94***	$E_t[PCEG_{t+1}]$				0.93***
				(3.13)					(4.19)
$GDPG_t$				-0.03	$PCEG_t$				0.02
				(-0.41)					(0.31)
$R^2_{adj}(\%)$	-1.55	7.92	29.93	49.49	$R^2_{adj}(\%)$	2.25	1.58	28.55	41.64

TABLE 8: **Out-of-sample predictability**: This table reports out-of-sample forecasting performance for equity excess returns, bond excess returns, and key macroeconomic outcomes using presidential approval (PAR), approval tightness ($PART$), and their combination (PAR+$PART$) as predictors. Forecasts are generated from recursively estimated predictive regressions at the horizon and frequency indicated in each panel. The in-sample initialization window is 1969-1996, and the out-of-sample evaluation period runs from 1997 through 2024. R^2_{OS} is the out-of-sample R^2 of Campbell and Thompson (2008). CW-test is the MSFE-adjusted Clark-West statistic of Clark and West (2007). *, **, and *** indicate significance at the 10%, 5%, and 1% levels, respectively.

	R^2_{OS}(%)	CW-t	R^2_{OS}(%)	CW-t
Panel A: Equity Return	1 Month		1 Quarter	
PAR	-0.30%	-0.85	-1.10%	-0.92
PART	2.02%	2.89***	5.06%	2.33***
PAR+PART	3.47%	2.49***	8.98%	2.30**
Panel B: GDP Growth	1 Quarter		1 Year	
PAR	0.34%	0.54	-3.96%	-0.65
PART	-0.57%	0.51	0.82%	0.9
PAR+PART	1.93%	2.21**	24.46%	1.42*
Panel C: PCE Growth	1 Quarter		1 Year	
PAR	0.85%	0.86	0.32%	0.37
PART	-0.66%	-0.01	-1.96%	-0.01
PAR+PART	1.88%	2.45***	18.64%	1.40*

TABLE 9: **Response of the Presidential Approval Rate to Economic Performance**: This table reports logit and probit regressions of the presidential approval rate on economic performance measures. Panel A uses a scaled logit specification, $PAR_t = a + b\,\sigma(\ell_t)$ with $\sigma(x) = 1/(1+e^{-x})$. Panel B uses a scaled probit specification, $PAR_t = a + b\,\Phi(\ell_t)$, where $\Phi(\cdot)$ is the standard normal cumulative distribution function. The baseline specification is

$$\ell_t = \alpha + \beta\,\ell_{t-1} + \delta^{\top}\text{Politics}_{t-1} + \gamma\,X_t + \varepsilon_{t+1},$$

where X_t denotes the economic-performance measure that varies across columns and Politics_t collects indicators for major political events. Columns (1) through (5) include one economic measure at a time together with inflation, and column (6) includes the set of economic measures jointly. The economic measures include real GDP growth, real PCE growth, the unemployment rate, $\log(Tax/GDP)_t$, and $\log(Spending/GDP)_t$. Newey-West standard errors with four lags are used for inference and the corresponding t-statistics are reported in parentheses. Significance stars denote rejection of the null hypothesis at the 1%, 5%, and 10% levels.

Panel A: Logit	(1)	(2)	(3)	(4)	(5)	(6)
Real GDP Growth$_t$	0.07**					0.04
	(2.35)					(0.52)
Real PCE Growth$_t$		0.05**				0.00
		(2.4)				(0.06)
Unemployment$_t$			-0.12***			-0.14***
			(-3.16)			(-3.39)
log(Tax/GDP)$_t$				-0.02		-0.10***
				(-0.5)		(-2.88)
log(Spending/GDP)$_t$					-0.08***	-0.06*
					(-2.68)	(-1.77)
Inflation$_t$	-0.16***	-0.16***	-0.18***	-0.18***	-0.16***	-0.17***
	(-3.09)	(-3.19)	(-3.94)	(-3.64)	(-3.09)	(-3.16)
ℓ_t	0.53***	0.53***	0.51***	0.53***	0.52***	0.50***
	(9.29)	(9.37)	(9.52)	(9.52)	(9.46)	(9.61)
R^2_{adj}(%)	77.37	77.27	77.93	77.1	77.31	78.18
Panel B: Probit	(1)	(2)	(3)	(4)	(5)	(6)
Real GDP Growth$_t$	0.04**					0.02
	(2.24)					(0.44)
Real PCE Growth$_t$		0.03**				0.00
		(2.37)				(0.12)
Unemployment$_t$			-0.07***			-0.08***
			(-3.11)			(-3.3)
log(Tax/GDP)$_t$				-0.01		-0.05***
				(-0.47)		(-2.8)
log(Spending/GDP)$_t$					-0.04***	-0.04*
					(-2.69)	(-1.75)
Inflation$_t$	-0.08***	-0.09***	-0.10***	-0.09***	-0.08***	-0.09***
	(-3.01)	(-3.09)	(-3.89)	(-3.56)	(-2.99)	(-3.07)
ℓ_t	0.56***	0.56***	0.54***	0.56***	0.55***	0.53***
	(10.3)	(10.42)	(10.55)	(10.57)	(10.47)	(10.51)
R^2_{adj}(%)	78.69	78.61	79.23	78.43	78.64	79.48

TABLE 10: **Presidential Approval and Government Expenditure Growth**: This table reports predictive regressions of real federal government expenditure growth on presidential approval (PAR_t), presidential approval tightness ($PART_t$), and related controls. In each panel, we estimate

$$dG_{t:t+h} = \alpha + \beta^{PAR} PAR_t + \beta^{PART} PART_t + \beta^{Controls} Controls_t + \varepsilon_{t:t+h},$$

where $dG_{t:t+h}$ denotes cumulative real federal expenditure growth from t to $t+h$. Panel A uses quarterly data. The left block forecasts one-quarter-ahead expenditure growth with $h = 1$, and the right block forecasts cumulative four-quarter expenditure growth with $h = 4$. Panel B uses annual data and forecasts one-year-ahead expenditure growth with $h = 1$. Across all specifications we control for lagged expenditure growth, dG_t. Additional controls are introduced across columns and include inflation and one fiscal-conditions variable at a time. The fiscal-conditions variables are the net surplus-to-debt ratio nsb_t, $\log(Tax/GDP)_t$, and $\log(Expenditure/GDP)_t$, where Tax denotes federal tax revenues and $Expenditure$ denotes federal government expenditures. Newey-West standard errors are computed with four lags and the corresponding t-statistics are reported in parentheses. Significance stars denote statistical significance at the 1%, 5%, and 10% levels.

Panel A: Quarterly	h=1					h=4				
	(1)	(2)	(3)	(4)	(5)	(1)	(2)	(3)	(4)	(5)
PAR_t	-0.06**	-0.07**	-0.09**	-0.07**	-0.11***	-0.21**	-0.22**	-0.27***	-0.21**	-0.31***
	(-1.96)	(-2.28)	(-2.3)	(-2.31)	(-2.59)	(-2.46)	(-2.52)	(-3.1)	(-2.55)	(-3.46)
$PART_t$	-0.33***	-0.37***	-0.33***	-0.37***	-0.31***	-1.06***	-1.11***	-1.01***	-1.11***	-0.95***
	(-2.6)	(-3.06)	(-3.02)	(-3.09)	(-2.86)	(-3.97)	(-4.04)	(-4.09)	(-4.02)	(-4.14)
dG_t	-0.25***	-0.28***	-0.25***	-0.28***	-0.22***	-0.24**	-0.28**	-0.18**	-0.28**	-0.11*
	(-3.32)	(-3.62)	(-3.18)	(-3.69)	(-2.99)	(-2.26)	(-2.5)	(-2.14)	(-2.54)	(-1.65)
$Inflation_t$		-1.14*	-1.26**	-1.10*	-1.25**		-1.29	-1.62	-1.22	-1.56*
		(-1.82)	(-1.96)	(-1.8)	(-2.03)		(-1.34)	(-1.63)	(-1.36)	(-1.69)
nsb_t			0.18*					0.49**		
			(1.79)					(2.09)		
$log(Tax/GDP)_t$				-0.05					-0.08	
				(-1.33)					(-0.66)	
$log(Expenditure/GDP)_t$					-0.12***					-0.31***
					(-2.72)					(-3.34)
R^2_{adj}(%)	5.84	7.8	9.81	7.58	12.34	6.03	7.03	14.95	6.86	22.36
Panel B: Annual	h=1									
	(1)	(2)	(3)	(4)	(5)					
PAR_t	-0.23***	-0.23***	-0.26***	-0.23***	-0.30***					
	(-3.23)	(-3.19)	(-4.0)	(-3.1)	(-3.94)					
$PART_t$	-1.28***	-1.28***	-1.25***	-1.28***	-1.27***					
	(-4.4)	(-4.34)	(-4.74)	(-4.24)	(-4.99)					
dG_t	-0.11	-0.11	-0.05	-0.12	-0.02					
	(-1.04)	(-1.08)	(-0.48)	(-0.98)	(-0.27)					
$Inflation_t$		0.01	-0.02	0.01	-0.05					
		(0.04)	(-0.14)	(0.05)	(-0.28)					
nsb_t			0.25***							
			(2.65)							
$log(Tax/GDP)_t$				-0.03						
				(-0.24)						
$log(Expenditure/GDP)_t$					-0.20***					
					(-3.46)					
R^2_{adj}(%)	9.66	7.86	11.56	6.12	19.22					

TABLE 11: **Presidential Approval and Federal Defense vs. Nondefense Spending Growth**: This table reports predictive regressions of real federal government expenditure growth on presidential approval (PAR_t), approval tightness ($PART_t$), and controls, estimated separately for defense and nondefense spending. We obtain annual national defense spending from NIPA Table 3.17 and define nondefense spending as total federal expenditure minus the national defense component. We estimate

$$dG_{t+1} = \alpha + \beta^{PAR} PAR_t + \beta^{PART} PART_t + \beta^{Controls} Controls_t + \varepsilon_{t+1},$$

where dG_{t+1} denotes next-year real expenditure growth and corresponds to $dG_{t+1}^{Defense}$ in the defense block and $dG_{t+1}^{Nondefense}$ in the nondefense block. All specifications include the lagged dependent variable, $dG_t^{Defense}$ for defense spending and $dG_t^{Nondefense}$ for nondefense spending. Additional controls are introduced across columns and include inflation and one fiscal-conditions variable at a time. The fiscal-conditions variables are the net surplus-to-debt ratio nsb_t, $\log(Tax/GDP)_t$, and $\log(Expenditure/GDP)_t$, where Tax denotes federal tax revenues and $Expenditure$ denotes federal government expenditures. Newey-West standard errors are computed with four lags and the corresponding t-statistics are reported in parentheses. Significance stars denote statistical significance at the 1%, 5%, and 10% levels.

	Defense Spending Growth						Nondefense Spending Growth				
	(1)	(2)	(3)	(4)	(5)		(1)	(2)	(3)	(4)	(5)
PAR_t	-0.05	-0.04	-0.08	-0.04	-0.09*	PAR_t	-0.24***	-0.23**	-0.28***	-0.23**	-0.34***
	(-0.93)	(-0.87)	(-1.55)	(-0.98)	(-1.66)		(-2.62)	(-2.56)	(-3.14)	(-2.51)	(-3.52)
$PART_t$	-0.17	-0.14	-0.13	-0.10	-0.15	$PART_t$	-1.61***	-1.60***	-1.59***	-1.62***	-1.63***
	(-0.64)	(-0.56)	(-0.51)	(-0.42)	(-0.59)		(-4.41)	(-4.25)	(-4.6)	(-4.21)	(-4.9)
$dG_t^{Defense}$	0.70***	0.75***	0.81***	0.81***	0.78***	$dG_t^{Nondefense}$	-0.07	-0.06	-0.01	-0.07	0.00
	(8.41)	(11.2)	(12.04)	(12.61)	(10.63)		(-0.63)	(-0.52)	(-0.08)	(-0.64)	(0.04)
$Inflation_t$		0.38***	0.34***	0.37***	0.34**	$Inflation_t$		0.15	0.09	0.16	0.05
		(2.84)	(2.93)	(3.81)	(2.52)			(0.45)	(0.27)	(0.48)	(0.15)
nsb_t			0.25***			nsb_t			0.32***		
			(4.22)						(3.35)		
$log(Tax/GDP)_t$				0.17***		$log(Tax/GDP)_t$				-0.07	
				(4.48)						(-0.41)	
$log(Expenditure/GDP)_t$					-0.11***	$log(Expenditure/GDP)_t$					-0.27***
					(-3.73)						(-3.98)
R^2_{adj}(%)	49.19	53.57	59.27	58.02	57.23	R^2_{adj}(%)	7.7	6.26	9.54	4.76	17.98

TABLE 12: **Presidential Approval and Governance Capacity**: This table reports predictive regressions of government governance capacity on presidential approval and related controls. The dependent variable is the *Control of Corruption* index (left block), or the *Governance Capacity* index (right block), both from the Worldwide Governance Indicators. The regressors of interest are presidential approval (PAR_t) and presidential approval tightness ($PART_t$). Specifications progressively add controls, including the lagged dependent variable (Y_t). Newey-West (HAC) t-statistics, computed with four lags, are reported in parentheses. Significance stars denote statistical significance at the 1%, 5%, and 10% levels.

	Y_{t+1}: Control of Corruption				Y_{t+1}: Government Capacity			
	(1)	(2)	(3)	(4)	(1)	(2)	(3)	(4)
PAR_t	0.01		0.07*	0.06*	-0.0		0.03	0.03
	(0.35)		(1.73)	(1.65)	(-0.03)		(1.38)	(1.42)
$PART_t$		0.39***	0.63**	0.59**		0.22***	0.31***	0.32***
		(2.84)	(2.44)	(2.27)		(2.99)	(3.1)	(3.06)
Y_t				-0.08				0.07
				(-0.85)				(0.52)
$R^2_{adj}(\%)$	-3.66	7.98	13.67	10.63	-3.99	14.79	19.37	16.41

Appendix A Derivations and Proofs in Theoretical Framework

A1 Analysis of Steady-State Solution

At date t, the incumbent chooses a pair (τ_t, x_t) that is implemented for the next T periods, $t+1, \dots, t+T$. To keep notation compact, we refer to this T-period policy block simply as (τ_t, x_t). Recall the approval-rate dynamics in the main text Eq. (14):

$$p_{t+1} = \Lambda(\ell_{t+1}), \qquad \Lambda(\ell) = \frac{1}{1+\exp(-\ell)}, \qquad \ell_{t+1} - \bar{\ell} = \phi\big(\ell_t - \bar{\ell}\big) + \gamma\big(q_{t+1} - \bar{q}\big) + \epsilon^{\ell}_{t+1}, \qquad \epsilon^{\ell}_{t+1} \text{ i.i.d..}$$

We study a conditional steady state evaluated at the average latent index $\ell_t = \bar{\ell}$, so that the approval level and dispersion are constant:

$$p \equiv \Lambda(\bar{\ell}), \qquad d(p) \equiv p(1-p) = \Lambda'(\bar{\ell}).$$

Under this simplified environment, we derive steady-state fiscal policy and value objects.

Proof of Proposition 1. Fix an election horizon $T \geq 1$. Mandated spending g is exogenous and constant within the policy block. Raw output follows Eq. (5), and the stochastic discount factor follows Eq. (10). Policy affects risk through Eq. (6), i.e.,

$$\sigma_y^2(x) = \sigma^2(1+x), \qquad \sigma^2 > 0.$$

Discounting induced by policy risk. Combine the raw-output growth process and the pricing kernel. Under constant x_t within the block, lognormality implies

$$\mathbb{E}_t\left[m_{t+1}\frac{Y^r_{t+1}}{Y^r_t}\right] = \exp\big(\mu_y - r + \eta\sigma_y^2(x_t)\big) \equiv e^{-\delta(x_t)}.$$

Define

$$\delta(x) = r + \eta\sigma^2(1+x) - \mu_y = \delta_0 + \lambda x, \qquad \lambda \equiv \eta\sigma^2 > 0, \qquad \delta_0 \equiv r + \lambda - \mu_y, \qquad E(x) \equiv e^{\delta(x)}. \tag{A1}$$

Because growth shocks are i.i.d. and policy-induced risk is time-invariant within the block,

$$\mathbb{E}_t[m_{t+i}Y^r_{t+i}] = Y^r_t e^{-i\delta(x_t)} = Y^r_t E(x_t)^{-i}, \qquad i = 1, \dots, T. \tag{A2}$$

Discount-sum objects. Define the T-period discount sums and the duration ratio:

$$A_T(x) \equiv \sum_{i=1}^{T} e^{-i\delta(x)} = \sum_{i=1}^{T} E(x)^{-i} = \frac{1 - E(x)^{-T}}{E(x) - 1}, \tag{A3}$$

$$D_T(x) \equiv \sum_{i=1}^{T} i\, e^{-i\delta(x)} = \sum_{i=1}^{T} i\, E(x)^{-i}, \tag{A4}$$

$$\Phi_T(x) \equiv \frac{D_T(x)}{A_T(x)}. \tag{A5}$$

Also define the continuation-to-flow valuation ratio:

$$\Psi_T(x) \equiv \frac{\mathbb{E}_t[m_{t+T}Y^r_{t+T}]}{\mathbb{E}_t[\sum_{i=1}^{T} m_{t+i}Y^r_{t+i}]} = \frac{E(x)^{-T}}{A_T(x)} = \frac{E(x) - 1}{E(x)^T - 1}. \tag{A6}$$

Approval sensitivity at the election date. In the conditional steady state $\ell_t = \bar{\ell}$, we have $d(p) = \Lambda'(\bar{\ell})$. Assume q depends smoothly on policy and define local derivatives evaluated at the steady state:

$$q_\tau \equiv \frac{\partial q}{\partial \tau} < 0, \qquad q_x \equiv \frac{\partial q}{\partial x} < 0.$$

Iterating the latent law forward T periods and using that the policy block chosen at t affects $q_{t+1}, \dots, q_{t+T}$ with the same marginal derivatives gives

$$\frac{\partial \ell_{t+T}}{\partial a_t} = \gamma \sum_{j=0}^{T-1} \phi^j \, q_a \equiv \Gamma_T \, q_a, \qquad a_t \in \{\tau_t, x_t\}, \qquad \Gamma_T \equiv \gamma \frac{1-\phi^T}{1-\phi}.$$

Therefore, by the chain rule,

$$\frac{\partial p_{t+T}}{\partial a_t} = \Lambda'(\bar{\ell}) \frac{\partial \ell_{t+T}}{\partial a_t} = \Gamma_T \, d(p) \, q_a, \qquad a_t \in \{\tau_t, x_t\}. \tag{A7}$$

Stationary budget identity and bond pricing. Let the net surplus share be $\Delta_t \equiv \tau_t - g - x_t$. Debt evolves according to the flow constraint in the main text:

$$B_{t+1} = B_t(1 + R^B_{t+1}) - \Delta_t \, Y^r_{t+1}. \tag{A8}$$

Consider a stationary candidate equilibrium with constant policy $(\tau_t, x_t) \equiv (\tau, x)$ and constant debt-to-raw-output ratio $b_t \equiv B_t / Y^r_t \equiv b$. Divide Eq. (A8) by Y^r_{t+1} and impose $B_{t+1}/Y^r_{t+1} = b$ and $B_t / Y^r_t = b$:

$$b = b \, \frac{1 + R^B_{t+1}}{Y^r_{t+1}/Y^r_t} - \Delta, \qquad \Delta \equiv \tau - g - x.$$

Hence

$$1 + R^B_{t+1} = \frac{Y^r_{t+1}}{Y^r_t}\Big(1 + \frac{\Delta}{b}\Big). \tag{A9}$$

No-arbitrage implies $\mathbb{E}_t[m_{t+1}(1 + R^B_{t+1})] = 1$. Substitute Eq. (A9) and use $\mathbb{E}_t[m_{t+1} Y^r_{t+1}/Y^r_t] = E(x)^{-1}$ (from Eq. (A1)) to obtain

$$1 = \Big(1 + \frac{\Delta}{b}\Big) E(x)^{-1} \quad \Longrightarrow \quad \Delta = b\big(E(x) - 1\big).$$

Equivalently,

$$\tau - g - x = b\big(E(x) - 1\big). \tag{A10}$$

Combining Eq. (A9) and Eq. (A10) yields

$$1 + R^B_{t+1} = E(x) \, \frac{Y^r_{t+1}}{Y^r_t}. \tag{A11}$$

Homogeneity of the value function and the steady-state value identity. Under T-period election timing, the incumbent solves the objective in the main text Eq. (13), which we write (with the same factorization assumption) as

$$V(B_t, Y^r_t, p_t) = \max_{\tau_t, x_t} \Big\{ \mathbb{E}_t\Big[\sum_{i=1}^{T} m_{t+i} Y^r_{t+i} \big(1 - \theta(\tau_t) - g + \kappa x_t\big) \Big] + \mathbb{E}_t[p_{t+T}] \, \mathbb{E}_t\big[m_{t+T} V(B_{t+T}, Y^r_{t+T}, p_{t+T}) \big] \Big\}, \tag{A12}$$

where $\theta(\tau) = \frac{1}{2} c^\tau \tau^2$ and $\kappa > 0$.

In the conditional steady state we take $\mathbb{E}_t[p_{t+T}] = p$ and $d(p)$ constant. Under i.i.d. growth, all present values scale with Y^r_t by Eq. (A2), motivating the homogeneous guess

$$V(B_t, Y^r_t, p) = Y^r_t \, v(b, p), \qquad b \equiv \frac{B_t}{Y^r_t}, \qquad V_B(B_t, Y^r_t, p) = \frac{\partial v}{\partial b}(b, p) \equiv v_b(b, p). \tag{A13}$$

Under constant policy, $\mathbb{E}_t[m_{t+T} Y^r_{t+T}] = Y^r_t \, E(x)^{-T}$ and $\mathbb{E}_t[\sum_{i=1}^{T} m_{t+i} Y^r_{t+i}] = Y^r_t \, A_T(x)$ by Eq. (A2)–Eq. (A3). Substitute Eq. (A13) into Eq. (A12) and divide by Y^r_t to obtain

$$v(b, p) = \big(1 - \theta(\tau) - g + \kappa x\big) A_T(x) + p \, E(x)^{-T} \, v(b, p),$$

or equivalently

$$v(b,p)\Big(1-p\,E(x)^{-T}\Big)=\big(1-\theta(\tau)-g+\kappa x\big)\,A_T(x). \tag{A14}$$

A key collapse identity for the debt-state derivative. Iterating Eq. (A8) over the T-block implies B_{t+T} depends on the cumulated product of one-period bond returns. Differentiating B_{t+T} with respect to (τ_t, x_t) gives

$$\frac{\partial B_{t+T}}{\partial \tau_t}=-\sum_{i=1}^{T} Y^r_{t+i}\prod_{j=i+1}^{T}(1+R^B_{t+j}),\qquad \frac{\partial B_{t+T}}{\partial x_t}=+\sum_{i=1}^{T} Y^r_{t+i}\prod_{j=i+1}^{T}(1+R^B_{t+j}). \tag{A15}$$

Using Eq. (A11), for any $i \le T$,

$$\prod_{j=i+1}^{T}(1+R^B_{t+j})=E(x)^{T-i}\frac{Y^r_{t+T}}{Y^r_{t+i}}.$$

Therefore

$$\sum_{i=1}^{T} Y^r_{t+i}\prod_{j=i+1}^{T}(1+R^B_{t+j})=Y^r_{t+T}\,\Theta_T(x),\qquad \Theta_T(x)\equiv\sum_{i=1}^{T}E(x)^{T-i}=\frac{E(x)^T-1}{E(x)-1}.$$

Multiplying by $\mathbb{E}_t[m_{t+T}Y^r_{t+T}]=Y^r_t E(x)^{-T}$ yields

$$\mathbb{E}_t[m_{t+T}Y^r_{t+T}]\,\Theta_T(x)=Y^r_t\,\frac{1-E(x)^{-T}}{E(x)-1}=Y^r_t A_T(x),$$

so

$$\mathbb{E}_t\left[m_{t+T}\sum_{i=1}^{T} Y^r_{t+i}\prod_{j=i+1}^{T}(1+R^B_{t+j})\right]=Y^r_t A_T(x). \tag{A16}$$

FOC for τ_t. Differentiate Eq. (A12) with respect to τ_t. The flow-payoff derivative is $-\theta'(\tau_t)=-c^\tau\tau_t$. Using Eq. (A15), the product rule for the continuation term, and the steady-state substitutions $\mathbb{E}_t[p_{t+T}]=p$, $V_{B,t+T}=v_b(b,p)$, and $\mathbb{E}_t[m_{t+T}V_{t+T}]=v(b,p)\,Y^r_t E(x)^{-T}$, we obtain

$$0=-c^\tau\tau\,Y^r_t A_T(x)+p\,v_b(b,p)\,\mathbb{E}_t\left[m_{t+T}\frac{\partial B_{t+T}}{\partial\tau_t}\right]+\left(\frac{\partial}{\partial\tau_t}\mathbb{E}_t[p_{t+T}]\right)v(b,p)\,Y^r_t E(x)^{-T}.$$

By Eq. (A15) and Eq. (A16), $\mathbb{E}_t[m_{t+T}\partial B_{t+T}/\partial\tau_t]=-Y^r_t A_T(x)$. Divide by $Y^r_t A_T(x)>0$, use $\Psi_T(x)=E(x)^{-T}/A_T(x)$ from Eq. (A6), and $\partial p_{t+T}/\partial\tau_t=\Gamma_T d(p)q_\tau$ from Eq. (A7) to get

$$c^\tau\tau=-p\,v_b(b,p)+\Gamma_T\,d(p)\,v(b,p)\,q_\tau\,\Psi_T(x). \tag{A17}$$

FOC for x_t. Differentiate Eq. (A12) with respect to x_t. Because x affects discounting through $\delta(x)$, the rent term contributes the derivative of $xA_T(x)$. Using $\delta_x=\lambda$ and $A'_T(x)=-\lambda D_T(x)$ (from Eq. (A3)–Eq. (A4)),

$$\frac{\partial}{\partial x}\big(xA_T(x)\big)=A_T(x)-\lambda x D_T(x)=A_T(x)\big(1-\lambda x\Phi_T(x)\big).$$

Hence the flow derivative is $\kappa\,Y^r_t A_T(x)\big(1-\lambda x\Phi_T(x)\big)$. For the continuation term, Eq. (A15) and Eq. (A16) imply $\mathbb{E}_t[m_{t+T}\partial B_{t+T}/\partial x_t]=+Y^r_t A_T(x)$. Proceeding exactly as for τ_t and using $\partial p_{t+T}/\partial x_t=\Gamma_T d(p)q_x$ yields

$$\kappa\big(1-\lambda x\Phi_T(x)\big)=-p\,v_b(b,p)-\Gamma_T\,d(p)\,v(b,p)\,q_x\,\Psi_T(x). \tag{A18}$$

In summary, the constant policy satisfies the first-order conditions Eq. (A17) and Eq. (A18). The steady state is pinned down jointly by the stationary budget identity Eq. (A10) and the value identity Eq. (A14).

The steady-state system in Proposition 1 provides a numerical mapping from the state variables (b,p,g) to next-term fiscal policy (τ,x) under a T-period policy block. We next establish two implications that isolate approval dispersion $d(p)=p(1-p)$ as the electoral-discipline channel: (i) when performance does not

affect approval ($\gamma = 0$), the optimal policy mix coincides with the benchmark without re-election discipline; (ii) under regularity, the optimal rent share $x(p)$ is U-shaped in approval p and locally convex at its interior minimum.

Proof of Proposition 2. Set $\gamma = 0$. Then $\Gamma_T = \gamma(1-\phi^T)/(1-\phi) = 0$, so policy does not affect election-date approval: $\partial p_{t+T}/\partial a_t = 0$ for $a_t \in \{\tau_t, x_t\}$. In the steady-state system, the political terms drop out and the two FOCs reduce to

$$c^\tau \tau = -p\, v_b(b,p), \qquad \kappa\big(1-\lambda x \Phi_T(x)\big) = -p\, v_b(b,p),$$

together with the stationary budget identity Eq. (A10). Define the composite shadow object $u \equiv -p\, v_b(b,p)$. Then

$$c^\tau \tau = u, \qquad \kappa\big(1-\lambda x \Phi_T(x)\big) = u,$$

which pins down (τ, x) jointly with Eq. (A10) as a function of (b, g) and structural parameters only. The approval level p affects only the scaling of v_b through $v_b = -u/p$, but does not affect the optimal policy mix:

$$\frac{\partial \tau}{\partial p} = 0, \qquad \frac{\partial x}{\partial p} = 0 \qquad (\gamma = 0).$$

Moreover, because p disappears from the policy system entirely, the resulting (τ, x) coincide with the no-re-election benchmark (equivalently $p \equiv 1$, hence $d(p) = 0$).

Proof of Proposition 3. Fix (b, g) and parameters. Work with the steady-state system in Proposition 1. Subtract the x-FOC from the τ-FOC to eliminate the common shadow term $-p\, v_b(b,p)$:

$$c^\tau \tau - \kappa\big(1-\lambda x \Phi_T(x)\big) = \Gamma_T\, d(p)\, v(b,p)\, \Psi_T(x)\, (q_\tau + q_x). \tag{A19}$$

Use the stationary budget identity Eq. (A10) to express taxes as a deterministic function of (x, b, g):

$$\tau(x) = g + x + b\big(E(x) - 1\big), \qquad E(x) = e^{\delta(x)}. \tag{A20}$$

Define net per-period resources in the incumbent's flow payoff:

$$N(x) \equiv 1 - \theta(\tau(x)) - g + \kappa x, \qquad \theta(\tau) = \tfrac{1}{2} c^\tau \tau^2. \tag{A21}$$

From the value identity Eq. (A14) and $\Psi_T(x)$ in Eq. (A6), the product $v(b,p)\Psi_T(x)$ simplifies to

$$v(b,p)\Psi_T(x) = \frac{N(x)}{E(x)^T - p}. \tag{A22}$$

Substituting Eq. (A20) and Eq. (A22) into Eq. (A19) yields a single equation in (x, p):

$$F(x,p) \equiv c^\tau \tau(x) - \kappa\big(1-\lambda x \Phi_T(x)\big) - \Gamma_T\, d(p)\, \frac{N(x)}{E(x)^T - p}\, (q_\tau + q_x) = 0. \tag{A23}$$

Assume an interior solution $x(p)$ such that $E(x(p))^T > p$ and the regularity condition $F_x(x(p), p) \neq 0$, so the Implicit Function Theorem applies:

$$x_p(p) = -\frac{F_p(x(p),p)}{F_x(x(p),p)}. \tag{A24}$$

Holding x fixed, F depends on p only through $d(p) = p(1-p)$ and $(E(x)^T - p)^{-1}$. Let $\Sigma q \equiv q_\tau + q_x < 0$ and $\tilde{E} \equiv E(x)^T$. Differentiating Eq. (A23) gives

$$F_p(x,p) = -\Gamma_T\, \Sigma q\, \frac{N(x)}{\tilde{E} - p} \left[(1-2p) + \frac{p(1-p)}{\tilde{E} - p} \right]. \tag{A25}$$

Define $S(p; \tilde{E}) \equiv (1-2p) + \frac{p(1-p)}{\tilde{E}-p}$. An interior turning point p^* (i.e., $x_p(p^*) = 0$) satisfies $S(p^*; \tilde{E}) = 0$,

equivalently

$$(p^*)^2 - 2\tilde{E}\,p^* + \tilde{E} = 0, \qquad \tilde{E} = E(x(p^*))^T. \tag{A26}$$

Since $S(0;\tilde{E}) = 1 > 0$ and $S(1;\tilde{E}) = -1 < 0$, $S(\cdot;\tilde{E})$ crosses zero once on $(0,1)$, so the turning point is unique whenever the solution curve is well-defined and F_x does not change sign.

To establish local convexity at p^*, differentiate Eq. (A24) again. At p^* we have $x_p(p^*) = 0$, so

$$x_{pp}(p^*) = -\frac{F_{pp}(x(p^*), p^*)}{F_x(x(p^*), p^*)}.$$

Using Eq. (A25) and the fact that $S(p^*;\tilde{E}) = 0$, one obtains

$$F_{pp}(x, p^*) = 2\Gamma_T\,\Sigma q\,\frac{N(x)}{\tilde{E} - p^*}.$$

Hence

$$x_{pp}(p^*) = -\frac{2\Gamma_T\,\Sigma q}{F_x(x(p^*), p^*)} \cdot \frac{N(x(p^*))}{E(x(p^*))^T - p^*}. \tag{A27}$$

Under $\Gamma_T > 0$, $\Sigma q < 0$, $N(x(p^*)) > 0$, $E(x(p^*))^T > p^*$, and $F_x(x(p^*), p^*) > 0$, Eq. (A27) implies $x_{pp}(p^*) > 0$. Thus $x(p)$ is locally convex at its interior minimum.

We now prove Proposition 4, i.e., holding all else fixed, how the shares of GDP, consumption, and dividends in pre-allocation output Y^r vary with rent extraction.

Proof of Proposition 4. From the resource identities in the main text,

$$\frac{Y_t}{Y_t^r} = 1 - \theta(\tau) - c^x x, \qquad \frac{C_t}{Y_t^r} = 1 - \theta(\tau) - g - (1 + c^x)x, \qquad \frac{D_t}{Y_t^r} = 1 - \theta(\tau) - c^x x - \tau.$$

Fix a policy experiment in which x varies and $\tau = \tau(x)$ adjusts according to the fiscal optimality relation in the non-political benchmark. In the one-period case ($T = 1$, so $\Phi_1(x) = 1$), eliminating the common shadow value between the fiscal FOCs yields

$$\tau(x) = \frac{\kappa}{c^\tau}(1 - \lambda x), \qquad x \in \Big[0, \frac{1}{\lambda}\Big]. \tag{A28}$$

Substituting Eq. (A28) into Y_t/Y_t^r gives

$$\frac{Y_t(x)}{Y_t^r} = 1 - c^x x - \frac{1}{2}c^\tau \tau(x)^2 = \Big(1 - \frac{\kappa^2}{2c^\tau}\Big) + \Big(-c^x + \frac{\kappa^2\lambda}{c^\tau}\Big)x - \frac{\kappa^2\lambda^2}{2c^\tau}x^2,$$

so

$$\frac{\partial}{\partial x}\Big(\frac{Y_t(x)}{Y_t^r}\Big) = -c^x + \frac{\kappa^2\lambda}{c^\tau} - \frac{\kappa^2\lambda^2}{c^\tau}x. \tag{A29}$$

The right-hand side is strictly decreasing in x on $[0, 1/\lambda]$ and attains its maximum at $x = 0$. Hence a necessary and sufficient condition for monotonicity on $(0, 1/\lambda]$ is

$$\frac{\partial}{\partial x}\Big(\frac{Y_t(x)}{Y_t^r}\Big)\Big|_{x=0} = -c^x + \frac{\kappa^2\lambda}{c^\tau} \le 0 \quad \Longleftrightarrow \quad \kappa^2\lambda \le c^\tau c^x,$$

which implies $\partial(Y_t(x)/Y_t^r)/\partial x < 0$ for all $x \in (0, 1/\lambda]$.

Consumption satisfies $\frac{C_t(x)}{Y_t^r} = \frac{Y_t(x)}{Y_t^r} - g - x$, so

$$\frac{\partial}{\partial x}\Big(\frac{C_t(x)}{Y_t^r}\Big) = \frac{\partial}{\partial x}\Big(\frac{Y_t(x)}{Y_t^r}\Big) - 1 < 0 \quad \text{for all } x \in (0, 1/\lambda],$$

under the same condition.

For dividends, $\frac{D_t(x)}{Y_t^r} = \frac{Y_t(x)}{Y_t^r} - \tau(x)$, so

$$\frac{\partial}{\partial x}\Big(\frac{D_t(x)}{Y_t^r}\Big) = \frac{\partial}{\partial x}\Big(\frac{Y_t(x)}{Y_t^r}\Big) - \tau_x(x), \qquad \tau_x(x) = -\frac{\kappa\lambda}{c^\tau}.$$

Thus,

$$\frac{\partial}{\partial x}\Big(\frac{D_t(x)}{Y_t^r}\Big) = -c^x + \frac{\kappa^2\lambda}{c^\tau} - \frac{\kappa^2\lambda^2}{c^\tau}x + \frac{\kappa\lambda}{c^\tau}.$$

This derivative is also decreasing in x, and a sufficient condition for it to be strictly negative on $(0, 1/\lambda]$ is the strengthened inequality

$$(\kappa^2 + \kappa)\lambda \le c^\tau c^x.$$

In the baseline text we state the conservative condition $\kappa^2\lambda \le c^\tau c^x$ to guarantee that rent extraction lowers real production and household allocations; dividend monotonicity holds whenever taxes do not fall too rapidly with x (including under the debt-stabilizing financing rule implicit in the stationary budget identity).

A2 Role of Debt Management and Mandated Spending

The steady-state system in Proposition 1 delivers a numerical mapping from (b, p, g) to optimal term policy (τ, x). This subsection isolates how the fiscal state (b, g) shifts policy, holding the political environment fixed at $(p, d(p))$.

Throughout, we use the following objects defined above: $E(x) = e^{\delta(x)}$ from Eq. (A1), $\Phi_T(x)$ from Eq. (A5), and $\Psi_T(x)$ from Eq. (A6). The stationary budget identity Eq. (A10) implies that for any candidate x, the associated tax share must satisfy

$$\tau(x; b, g) = g + x + b\big(E(x) - 1\big). \tag{A30}$$

Define the flow-payoff resource term (net of tax distortions and mandated spending, including private rents)

$$N(x; b, g) \equiv 1 - \theta\big(\tau(x; b, g)\big) - g + \kappa x, \qquad \theta(\tau) = \tfrac{1}{2}c^\tau\tau^2. \tag{A31}$$

Combining the value identity in Proposition 1 (equivalently Eq. (A14)) with $\Psi_T(x)$ implies the closed-form product

$$v(b, p)\,\Psi_T(x) = \frac{N(x; b, g)}{E(x)^T - p}. \tag{A32}$$

Next, subtract the x-FOC Eq. (19) from the τ-FOC Eq. (18) to eliminate the common continuation shadow term:

$$c^\tau\tau - \kappa\big(1 - \lambda x\Phi_T(x)\big) = \Gamma_T\, d(p)\, v(b, p)\, \Psi_T(x)\,(q_\tau + q_x).$$

Substituting Eq. (A30) and Eq. (A32) yields a single equilibrium restriction determining x as a function of (b, g, p):

$$F(x; b, g, p) \equiv c^\tau\,\tau(x; b, g) - \kappa\big(1 - \lambda x\Phi_T(x)\big) - \Gamma_T\, d(p)\,(q_\tau + q_x)\,\frac{N(x; b, g)}{E(x)^T - p} = 0. \tag{A33}$$

Given a solution $x(b, g, p)$ to Eq. (A33), taxes follow from Eq. (A30).

Debt management: how policy moves with b (holding (g, p) fixed). Assume an interior solution satisfying the regularity condition $F_x(x(b, g, p); b, g, p) \ne 0$. By the Implicit Function Theorem,

$$\frac{\partial x}{\partial b} = -\frac{F_b}{F_x}.$$

Holding x fixed, Eq. (A30) implies $\partial\tau/\partial b = E(x) - 1 > 0$. Moreover, since N depends on b only through τ,

$$\frac{\partial N}{\partial b} = -\theta'(\tau)\frac{\partial \tau}{\partial b} = -(c^\tau \tau)\big(E(x) - 1\big).$$

Differentiating Eq. (A33) with respect to b (holding x fixed) yields

$$F_b(x; b, g, p) = c^\tau \left(E(x) - 1\right)\left[1 + \Gamma_T\, d(p)\, (q_\tau + q_x)\, \frac{\tau(x; b, g)}{E(x)^T - p}\right]. \tag{A34}$$

Since $E(x) > 1$ and $c^\tau > 0$, the leading factor is positive. Because $(q_\tau + q_x) < 0$, the bracketed correction is typically below 1 and remains positive provided

$$1 + \Gamma_T\, d(p)\, (q_\tau + q_x)\, \frac{\tau(x; b, g)}{E(x)^T - p} > 0. \tag{A35}$$

Under Eq. (A35) and the stability case $F_x > 0$ at the selected root, we obtain $\partial x/\partial b < 0$: higher inherited debt disciplines rent extraction.

Taxes satisfy $\tau(x; b, g) = g + x + b(E(x) - 1)$, so

$$\frac{\partial \tau}{\partial b} = (E(x) - 1) + \big(1 + b\, E_x(x)\big)\frac{\partial x}{\partial b}, \qquad E_x(x) = \lambda E(x) > 0.$$

The first term is strictly positive; the second term is negative when $\partial x/\partial b < 0$. Therefore, $\partial\tau/\partial b$ is generically positive, and it can only fail to be positive if x falls so sharply that the induced decline in $E(x)$ overwhelms the direct refinancing pressure.

Mandated spending: how policy moves with g (holding (b, p) fixed). Assume again an interior solution with $F_x \neq 0$. By the Implicit Function Theorem,

$$\frac{\partial x}{\partial g} = -\frac{F_g}{F_x}.$$

Holding x fixed, Eq. (A30) gives $\partial\tau/\partial g = 1$, and

$$\frac{\partial N}{\partial g} = -\theta'(\tau)\frac{\partial \tau}{\partial g} - 1 = -(c^\tau \tau) - 1 < 0.$$

Differentiating Eq. (A33) with respect to g (holding x fixed) yields

$$F_g(x; b, g, p) = c^\tau + \Gamma_T\, d(p)\, (q_\tau + q_x)\, \frac{c^\tau \tau(x; b, g) + 1}{E(x)^T - p}. \tag{A36}$$

Because $(q_\tau + q_x) < 0$, the second term is negative. Thus $F_g > 0$ (hence $\partial x/\partial g < 0$ under $F_x > 0$) provided

$$c^\tau + \Gamma_T\, d(p)\, (q_\tau + q_x)\, \frac{c^\tau \tau(x; b, g) + 1}{E(x)^T - p} > 0. \tag{A37}$$

This delivers the core implication: higher mandated spending tends to reduce rent extraction because cutting x both frees resources one-for-one and lowers refinancing pressure through $E(x)$.

Taxes respond according to

$$\frac{\partial \tau}{\partial g} = 1 + \big(1 + bE_x(x)\big)\frac{\partial x}{\partial g}, \qquad E_x(x) = \lambda E(x) > 0,$$

so $\partial\tau/\partial g$ is not sign-restricted: the direct +1 effect can be offset by a sufficiently strong endogenous rent cut. If the substitution between x and g is strong, taxes may fall even as g rises.

In the political steady state, both higher inherited debt b and higher mandated spending g tend to

discipline rent seeking (x falls), because cutting rents relaxes the budget mechanically and reduces refinancing costs by lowering $E(x)$. The key contrast is on the tax margin: debt increases create robust positive refinancing pressure on τ, whereas higher g can be financed largely by cutting x and lowering discount-rate pressure, so τ need not rise.

A3 Linkage to Predictability Results

Proposition 1 delivers a static mapping from the date-t state (b_t, g_t, p_t, d_t) to an optimal *next-term* policy choice (τ_{t+1}, x_{t+1}). The empirical object, however, is predictability: how political conditions at date t forecast future cash flows and returns. The link is the term-policy timing. Since (τ_{t+1}, x_{t+1}) applies over $t+1, \ldots, t+T$, movements in political conditions at t shift the wedges faced by households starting at $t+1$, thereby moving cash flows through Eq. (8)–Eq. (9) and discount rates through Eq. (10).

Proof of Proposition 5. Start from the resource wedges implied by Eq. (8)–Eq. (9):

$$\frac{Y_{t+1}}{Y^r_{t+1}} = 1 - \theta(\tau_{t+1}) - c^x x_{t+1}, \qquad \frac{C_{t+1}}{Y^r_{t+1}} = 1 - \theta(\tau_{t+1}) - g - (1 + c^x) x_{t+1}, \qquad \frac{D_{t+1}}{Y^r_{t+1}} = 1 - \theta(\tau_{t+1}) - c^x x_{t+1} - \tau_{t+1}.$$

Under the steady-state benchmark at date t, the corresponding objects equal

$$\frac{\bar{Y}}{\bar{Y}^r} = 1 - \theta(\bar{\tau}) - c^x \bar{x}, \qquad \frac{\bar{C}}{\bar{Y}^r} = 1 - \theta(\bar{\tau}) - g - (1 + c^x) \bar{x}, \qquad \frac{\bar{D}}{\bar{Y}^r} = 1 - \theta(\bar{\tau}) - c^x \bar{x} - \bar{\tau}.$$

For output,

$$\log\left(\frac{Y_{t+1}}{\bar{Y}}\right) = \log\left(\frac{Y^r_{t+1}}{\bar{Y}^r}\right) + \log\left(\frac{1 - \theta(\tau_{t+1}) - c^x x_{t+1}}{1 - \theta(\bar{\tau}) - c^x \bar{x}}\right).$$

Using $\log(1-u) \approx -u$ yields

$$\log\left(\frac{1 - \theta(\tau_{t+1}) - c^x x_{t+1}}{1 - \theta(\bar{\tau}) - c^x \bar{x}}\right) \approx -\big(\theta(\tau_{t+1}) - \theta(\bar{\tau})\big) - c^x \left(x_{t+1} - \bar{x}\right) = -\Delta\theta_{t+1} - c^x \, \Delta x_{t+1},$$

which delivers the stated expression for $\log(Y_{t+1}/\bar{Y})$. The derivations for $\log(C_{t+1}/\bar{C})$ and $\log(D_{t+1}/\bar{D})$ are identical, noting that g is time-invariant in the benchmark environment and that dividends also load on τ_{t+1}.

For returns, let P_t denote the equity price and M_{t+1} the pricing kernel. Equity is a claim to dividends $\{D_{t+j}\}_{j \geq 1}$, so

$$P_t = \mathbb{E}_t\Big[\sum_{j=1}^{\infty} \frac{M_{t+j}}{M_t} D_{t+j}\Big].$$

In the model, the risk-adjusted discount rate is $\delta(x) = \delta_0 + \lambda x$, so at the steady-state benchmark $x_t = \bar{x}$ the equity price-dividend ratio is constant and satisfies

$$P_t = \frac{e^{-\delta(\bar{x})}}{1 - e^{-\delta(\bar{x})}} \, D_t.$$

Equivalently,

$$1 + \frac{D_t}{P_t} = e^{\delta(\bar{x})} \quad \Rightarrow \quad dp_t \equiv \log\left(1 + \frac{D_t}{P_t}\right) = \delta(\bar{x}).$$

Using the Campbell–Shiller present-value identity in the form reported by Campbell et al. (2023),

$$(1 - \rho)\, r^E_{t+1} - (1 - \rho)\, \Delta d_{t+1} = dp_t - \rho \, dp_{t+1}, \qquad \rho \equiv \frac{1}{1 + \bar{D}/\bar{P}},$$

where $\Delta d_{t+1} \equiv \log(D_{t+1}/D_t)$. Under the steady-state benchmark $D_t = \bar{D}$, $\Delta d_{t+1} = \log(D_{t+1}/\bar{D}) = \Delta \log D_{t+1}$. Moreover, since $dp_{t+1} = \delta(x_{t+1}) = \delta(\bar{x}) + \lambda(x_{t+1} - \bar{x}) = dp_t + \lambda \, \Delta x_{t+1}$,

$$dp_t - \rho \, dp_{t+1} = (1 - \rho)\, dp_t - \rho \, \lambda \, \Delta x_{t+1}.$$

Substituting into the identity and dividing by $1-\rho$ yields

$$r^E_{t+1} = \Delta \log D_{t+1} + dp_t - \frac{\rho}{1-\rho}\lambda \Delta x_{t+1} = \Delta \log D_{t+1} + \delta(\bar{x}) - \frac{\rho}{1-\rho}\lambda \Delta x_{t+1}.$$

A4 Numerical Calibration

Based on the comparative statics, we calibrate the key parameters in Proposition 1. We set the debt-to-GDP ratio and PAR to their historical means, $p = 0.55$ and $b = 0.4$. We set volatility to $\sigma = 0.15$. To match the equity expected return, we set $\lambda = \eta\sigma^2 = 0.1$ and $\delta_0 = r + \lambda - \mu_y = 0.1$. For simplicity, we set $\Gamma_T \equiv \gamma(1-\phi^T)/(1-\phi) = 1$, and $T = 8$ to capture the typical horizon of a two-term incumbent. Lastly, we choose the remaining parameters to match the historical mean of the spending-to-GDP ratio, $g + x \approx 0.17$, when (p, b) are also at their historical means: $\kappa = 1.5$, $c^\tau = 4$, $q_\tau = -4$, $c^x = 1$, and $q_x = -1$. These parameter values generate the comparative statics for x as a function of PAR shown in Figure 5.

Mapping to our analysis of debt-management motives, we also report how x and the associated variables vary with b in Figure A.1 and with g in Figure A.2. The debt channel is straightforward: higher debt reduces rent extraction and therefore predicts higher future equity returns, consistent with Proposition 5. This mechanism aligns with empirical findings in the literature (e.g., Liu (2023)). The impact of g depends on calibration. Under the baseline parameterization, higher mandated spending reduces x and raises τ. This tax response is stronger than the debt channel because the primary surplus must adjust directly to finance higher mandated spending. As a result, the dividend and consumption shares of pre-allocation output may decrease even as x falls, because higher taxes and the associated distortions are needed to finance g, as illustrated by the non-monotonic patterns in Figure A.2. In other words, even benevolent spending may not always increase private-sector allocations in this stylized environment.

Appendix B Additional Results and Robustness

B1 Bond Return Predictability

For bond returns, the same log-linear argument as in the proof of Proposition 5 yields

$$r^B_{t+1} = \Delta \log NS_{t+1} + \delta(\bar{x}) - \frac{\rho}{1-\rho}\lambda \Delta x_{t+1}.$$

In steady state, the net surplus share satisfies

$$\frac{NS}{Y^r} = \tau(\bar{x}; b, g) - \left(g + \bar{x}\right) = b\left(e^{\delta(\bar{x})} - 1\right).$$

Log-differentiating implies

$$\Delta \log NS_{t+1} = \Delta \log Y^r_{t+1} + \Delta \log b_{t+1} + \lambda \Delta x_{t+1}.$$

Substituting into the expression for r^B_{t+1} and collecting terms yields

$$r^B_{t+1} = \Delta \log Y^r_{t+1} + \Delta \log b_{t+1} + \lambda \Delta x_{t+1}. \tag{B1}$$

A higher rent share x raises the discount rate $\delta(x)$. At the same time, the government budget identity $\tau(x) - g - x = b(E(x) - 1)$ implies that the primary surplus—and hence fiscal backing—co-moves with x, making the overall valuation effect a priori ambiguous. To clarify this mechanism, we decompose Eq. (B1) into two components. The first is a direct policy-induced risk component, $\lambda \Delta x_{t+1}$, which captures the effect holding the debt ratio fixed. Under this channel, a higher PART—by lowering x—reduces the riskiness of the net surplus and predicts lower Treasury returns. The second component, $\Delta \log b_{t+1}$, reflects valuation changes driven by movements in the debt-to-GDP ratio, which may arise from channels beyond political conditions.

As a remark, a large literature documents that U.S. government debt can be valued above its fiscal backing because it provides safe-asset services (e.g., Krishnamurthy and Vissing-Jorgensen, 2012; He et al., 2019; Jiang et al., 2024c; Chen et al., 2025). Put differently, the U.S. government may borrow at lower effective cost and sustain larger deficits than in the benchmark environment considered here. Our model abstracts from this channel and instead focuses on how PART disciplines rent extraction. Empirically, we therefore test whether higher PART predicts lower Treasury returns after controlling for standard debt-valuation factors.

We obtain annualized continuously compounded zero-coupon yields with maturities ranging from 1 month to 360 months from Liu and Wu (2021). Following Cochrane and Piazzesi (2005) and Ludvigson and Ng (2009), we construct bond excess returns over a 12-month holding period. Let

$$y_t^{(n)} = -\frac{p_t^{(n)}}{n},$$

where $p_t^{(n)}$ and $y_t^{(n)}$ denote the log price and log yield of an n-year zero-coupon bond at time t. Suppose an investor purchases an n-year bond at time t for price $p_t^{(n)}$, holds it for 12 months, and sells it at time $t+12$ for price $p_{t+12}^{(n-1)}$. The 12-month log return is

$$r_{t:t+12}^{(n)} = p_{t+12}^{(n-1)} - p_t^{(n)},$$

and the corresponding 12-month excess return is

$$rx_{t:t+12}^{(n)} = r_{t:t+12}^{(n)} - y_t^{(1)}.$$

We focus on excess returns of shorter-maturity bonds, which behave less like safe assets than long-term bonds—specifically, 2-year and 5-year Treasuries. Table B.1 reports predictive regressions of bond excess returns on presidential approval, PAR_t, approval tightness, PART_t, and control variables. We estimate

$$rx_{t:t+12}^{(i)} = \alpha + \beta^{PAR} PAR_t + \beta^{PART} PART_t + \beta^{Controls} Controls_t + \varepsilon_{t:t+12},$$

where $rx_{t:t+12}^{(i)}$ denotes the cumulative 12-month excess return on the i-year Treasury. The left block reports results for 2-year excess returns, $rx_{t:t+12}^{(2)}$, and the right block reports results for 5-year excess returns, $rx_{t:t+12}^{(5)}$. All specifications include the Cochrane–Piazzesi factor CP_t following Cochrane and Piazzesi (2005). Columns (1) and (2) report univariate specifications in which PAR_t or PART_t is the sole political predictor. Column (3) includes both jointly, and the final column adds a full set of macroeconomic controls following Ludvigson and Ng (2009). These macroeconomic variables are entered with a one-month publication lag.

The left block shows that CP_t is positive and highly significant across all specifications, consistent with Cochrane and Piazzesi (2005). Conditional on CP_t, PAR_t exhibits little predictive content for 2-year excess returns. In contrast, PART_t enters negatively and is statistically significant, and it remains significant when PAR_t and PART_t are included jointly. Adding the macroeconomic control set does not eliminate this predictability: the coefficient on PART_t remains negative and statistically significant, indicating that approval tightness contains incremental information about bond risk premia beyond that captured by forward-rate information and standard macroeconomic conditions.

The right block for 5-year excess returns delivers the same qualitative pattern. The CP_t factor remains strongly significant, PAR_t does not exhibit robust predictive power conditional on CP_t, and PART_t remains negative and statistically significant in both the univariate and joint specifications, including after adding the full control set. Overall, Table B.1 shows that approval tightness is a robust predictor of bond excess returns across maturities. The negative predictive relation between PART_t and bond excess returns is consistent with the model implication that higher approval tightness reduces rent extraction and lowers the exposure of net fiscal surpluses to aggregate output risk.

B2 Unsmoothed Presidential Approval Rate

Our baseline results use smoothed level and tightness of approval rate and we ensure this noise-reduction procedure do not make these time-series non-stationary. As an extra robustness check, Table B.2 repeats the predictive regressions using the unsmoothed presidential approval rate, PAR_t, and the unsmoothed approval tightness, $PART_t$. For each outcome, we estimate

$$Y_{t:t+h} = \alpha + \beta^{PAR} PAR_t + \beta^{PART} PART_t + \beta^{Control} Control_t + \varepsilon_{t:t+h}.$$

The dependent variables span three domains: equity returns, real economic growth, and uncertainty. For equity, we consider both monthly and quarterly frequencies, using short- and longer-horizon outcomes with $h \in \{1, 12\}$ months and $h \in \{1, 4\}$ quarters. For real activity, we study quarterly real GDP growth and real PCE growth at horizons $h \in \{1, 4\}$ and include lagged growth Y_t to account for persistence. For uncertainty, we focus on the average uncertainty level over the subsequent 12 months for the VIX, the EPU index, and the financial, macro, and real uncertainty measures of Jurado et al. (2015), and we include lagged uncertainty Y_t to account for persistence.

Across panels, the results are qualitatively similar to the baseline estimates based on smoothed approval measures. In Panel A, $PART_t$ remains strongly predictive of equity excess returns at both monthly and quarterly horizons, and PAR_t retains positive predictive content once $PART_t$ is included. Panel B shows that both PAR_t and $PART_t$ remain positively associated with subsequent GDP and PCE growth at one- and four-quarter horizons, even after controlling for lagged growth. Panel C shows that both approval measures continue to forecast lower future uncertainty, with negative coefficients for PAR_t and $PART_t$ across the uncertainty proxies. Overall, Table B.2 indicates that the predictive relations documented in the baseline analysis are robust to whether the approval series are smoothed or left in raw form.

B3 Extreme Values in Presidential Approval Rate

Figure 1 shows that the presidential approval rating curve is smooth for most periods, but it occasionally exhibits abrupt shifts—phenomena commonly labeled "jumps" in the literature. While the jumps can serve as the primary driver of the predictability, in this subsection, we identify the observations that contain jumps and then examine whether the predictive power persists after removing them.

Lee and Mykland (2008) proposed a test for the presence of jumps in time series. Under the null hypothesis of "no jumps", the process follows a pure diffusion:

$$dX(t) = \mu(t)dt + \sigma(t)dW(t)$$

where $W(t)$ is a standard Brownian motion. The alternative hypothesis allows for discontinuities:

$$dX(t) = \mu(t)dt + \sigma(t)dW(t) + Y(t)dJ(t)$$

where $J(t)$ denotes a jump-counting process of intensity $\lambda(t)$ and the jump size $Y(t)$ at time t. The test assesses whether the jump component $Y(t)\,dJ(t)$ is statistically distinguishable from zero.

We consider an economy in which a fraction p of agents approve of the incumbent president, whereas the remaining $1 - p$ disapprove. At each period t, a sample of n agents is surveyed and asked whether they approve or disapprove of the incumbent president. We define the i-th agent, $X_{(i,t)} = 1$, if the agent approves of the current president, and $X_{(i,t)} = 0$ if the agent disapproves. Then the presidential approval rating $P(t) = \frac{1}{n}\sum_{i=1}^{n} X_{(i,t)}$ follows a binomial distribution:

$$P(t) \sim \text{Binomial}(n, p)$$

When n is large enough, $P(t)$ approximates a normal distribution: .

$$P(t) \sim N(p, \frac{p(1-p)}{n})$$

Thus, the dynamics of $P(t)$ can be described by the following equation:

$$dP(t) = \mu(t)\,dt + \sigma(t)\,dW(t)$$

Following Lee and Mykland (2008), we can test for the presence of jumps in $P(t)$.

Figure B.1 demonstrates the jumps in Presidential Approval Rating. The red upward triangle and the green downward triangle represent positive and negative jumps, respectively. Because moving-average smoothing attenuates fluctuations—and thus diminishes the number of jumps detected—we employ the raw, unsmoothed Presidential Approval Rating series as the input to the jump-detection test. Figure B.1 reveals that the PAR series contains several jumps—most are positive, whereas only a few are negative.

To assess whether extreme movements in presidential approval drive our baseline predictability results, Table B.3 repeats the predictive regressions after excluding periods associated with jumps in PAR_t, identified using the procedure of Lee and Mykland (2008). In the monthly sample, we drop the jump month and the subsequent five months. In the quarterly sample, we drop the jump quarter and the subsequent quarter, and we classify a quarter as a jump quarter if any of its three constituent months contains a jump. For each outcome, we estimate

$$Y_{t:t+h} = \alpha + \beta^{PAR} PAR_t + \beta^{PART} PART_t + \beta^{Control} Control_t + \varepsilon_{t:t+h}.$$

The dependent variables span three domains: equity returns, real economic growth, and uncertainty. For equity, we consider both monthly and quarterly frequencies, using horizons $h \in \{1, 12\}$ months and $h \in \{1, 4\}$ quarters. For real activity, we study quarterly real GDP growth and real PCE growth at horizons $h \in \{1, 4\}$ and include lagged growth Y_t to account for persistence. For uncertainty, we focus on the average uncertainty level over the subsequent 12 months for the VIX, the EPU index, and the financial, macro, and real uncertainty measures of Jurado et al. (2015), and we include lagged uncertainty Y_t to account for persistence.

The results remain qualitatively similar to the baseline estimates. For equity excess returns, both PAR_t and $PART_t$ retain positive and statistically significant coefficients across monthly and quarterly horizons, and the explanatory power remains economically meaningful. For real activity, PAR_t and $PART_t$ continue to forecast higher future GDP and PCE growth at both one- and four-quarter horizons after controlling for lagged growth. For uncertainty, both approval measures continue to predict lower future uncertainty, with negative and generally significant coefficients across the uncertainty proxies. Overall, the persistence of these patterns after excluding approval-jump windows indicates that our main predictability results are not driven by a small number of extreme movements in PAR_t.

Additional Figures and Tables

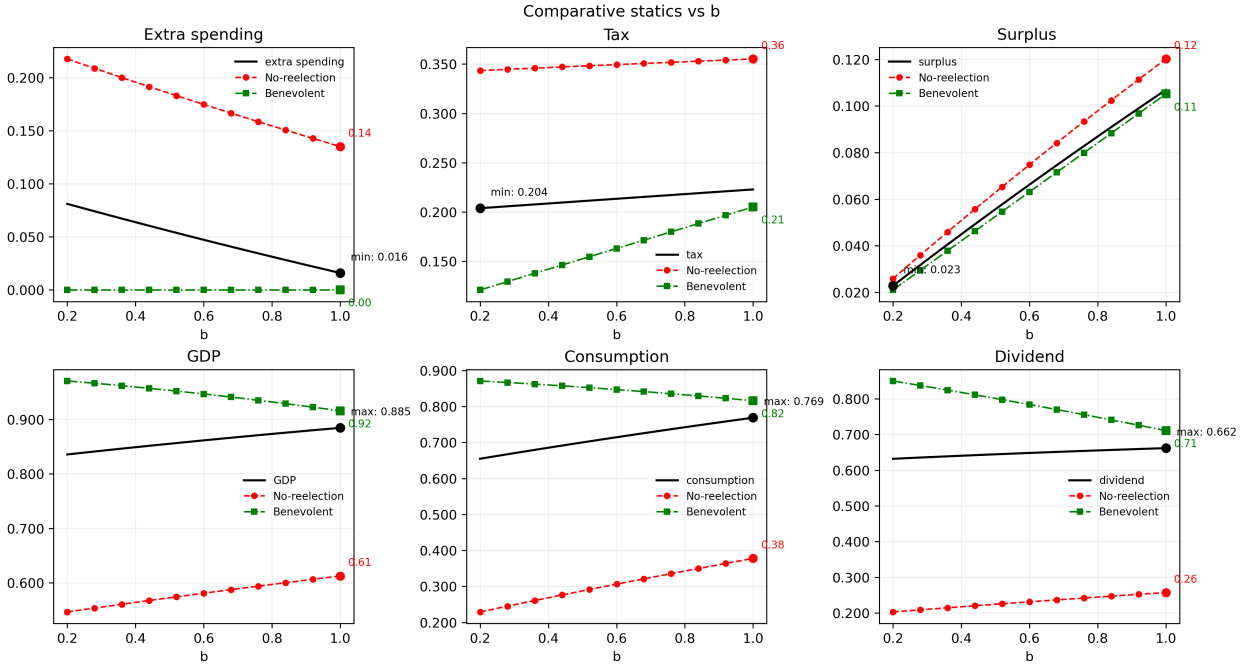


FIGURE A.1: **Debt Management, Rent Extraction, and Real Allocations.** This plot shows how the model-implied optimal rent extraction x and associated macroeconomic outcomes vary with government debt b, as characterized in Proposition 1. The top-left panel plots equilibrium rent extraction $x(b)$ (black) against the benevolent benchmark $x \equiv 0$ (green dashed) and the no–re-election benchmark (red dashed). The top-middle and top-right panels report the implied shares of taxes and surpluses in pre-allocation resources. The bottom panels translate these policy choices into the implied shares of GDP, consumption, and dividends in pre-allocation resources defined by equation Eq. (5).

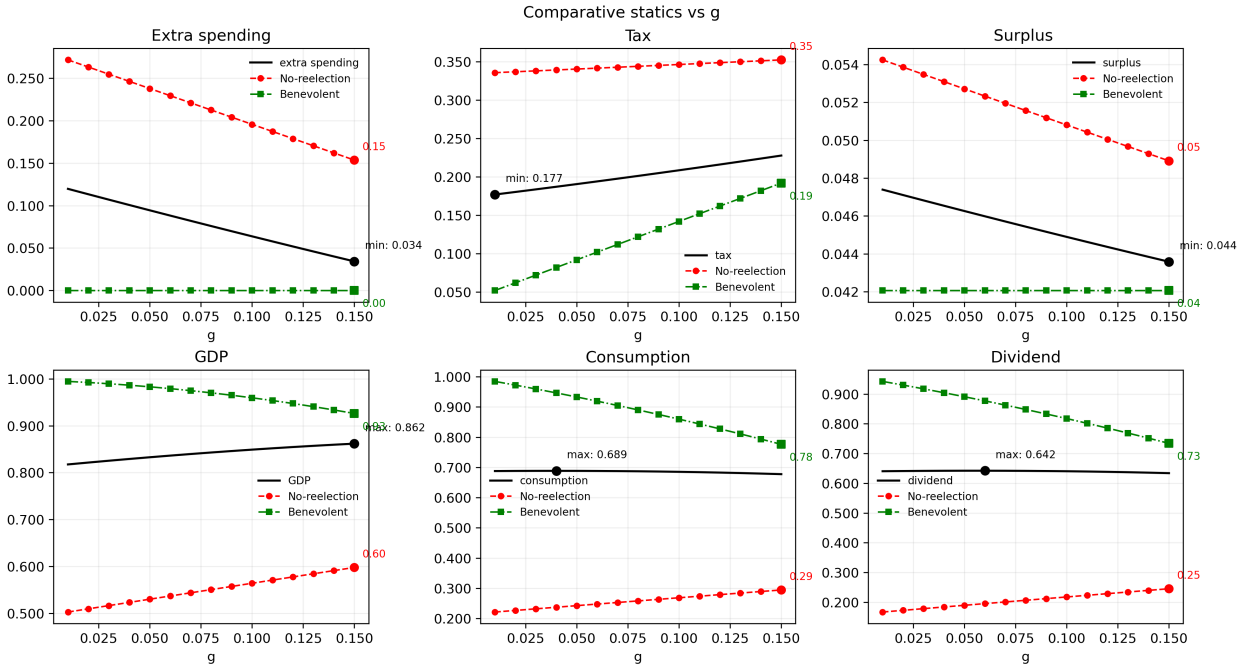


FIGURE A.2: **Mandated Spending, Rent Extraction, and Real Allocations.** This plot shows how the model-implied optimal rent extraction x and associated macroeconomic outcomes vary with mandated spending g, as characterized in Proposition 1. The top-left panel plots equilibrium rent extraction $x(g)$ (black) against the benevolent benchmark $x \equiv 0$ (green dashed) and the no–re-election benchmark (red dashed). The top-middle and top-right panels report the implied shares of taxes and surpluses in pre-allocation resources. The bottom panels translate these policy choices into the implied shares of GDP, consumption, and dividends in pre-allocation resources defined by equation Eq. (5).

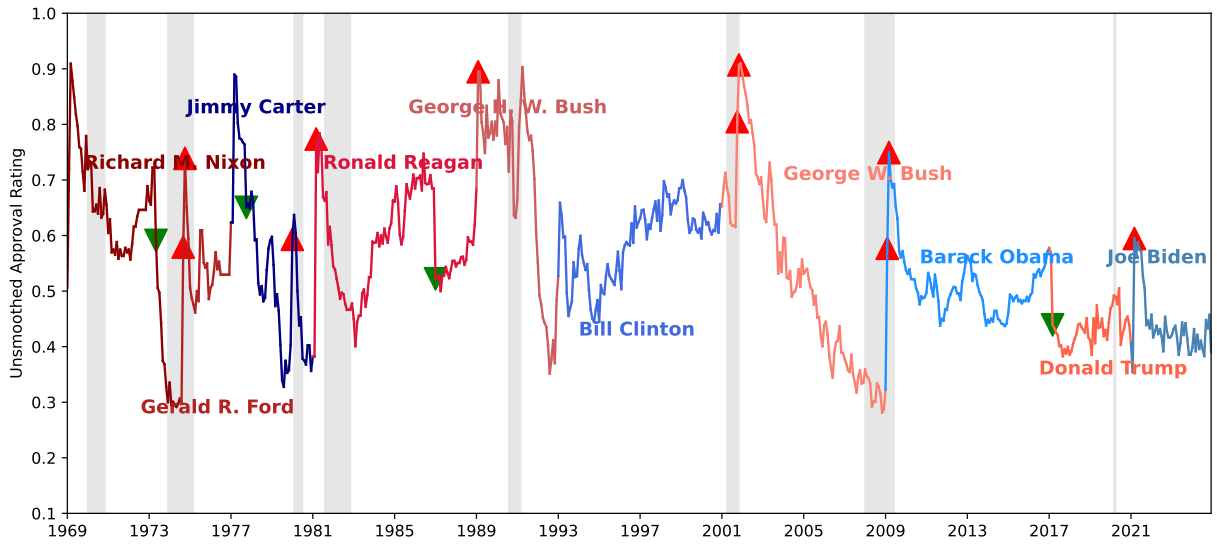


FIGURE B.1: **Unsmoothed Presidential Approval Rating and Jumps** This figure plots the unsmoothed monthly time series of Gallup's presidential approval rating (PAR). Shaded bands denote NBER-dated recession periods. The sample spans January 1969 to December 2024. Colored segments indicate presidential administrations and party affiliation, with warm tones denoting Republican presidents and cool tones denoting Democratic presidents. Red upward triangles and green downward triangles mark detected positive and negative jumps, respectively.

TABLE B.1: **PAR and Bond Returns**: This table reports predictive regressions of bond excess returns on presidential approval (PAR_t), approval tightness (PART_t), and controls. The left block uses excess returns on the 2-year Treasury over the next 12 months, $rx^{(2)}_{t:t+12}$, and the right block uses excess returns on the 5-year Treasury over the next 12 months, $rx^{(5)}_{t:t+12}$. All specifications include the Cochrane–Piazzesi factor, CP_t, following Cochrane and Piazzesi (2005). Columns (1) and (2) report univariate specifications that include PAR_t or PART_t as the sole political predictor, and column (3) includes both jointly. The final column adds the full control set, including the macroeconomic factors of Ludvigson and Ng (2009). The row labeled Controls indicates whether this full control set is included. Newey-West (HAC) t-statistics are reported in parentheses using 18 lags. Significance stars denote rejection of the null hypothesis at the 1%, 5%, and 10% levels.

	Y:$rx^{(2)}_{t:t+12}$				Y:$rx^{(5)}_{t:t+12}$			
PAR_t	0.01		-0.02	-0.01	0.01		-0.07	-0.03
	(0.82)		(-1.07)	(-0.35)	(0.28)		(-1.23)	(-0.45)
PART_t		-0.15***	-0.21***	-0.15**		-0.32**	-0.55***	-0.44**
		(-3.15)	(-3.17)	(-2.06)		(-2.21)	(-2.72)	(-2.01)
CP_t	0.39***	0.39***	0.43***	0.40***	1.40***	1.36***	1.51***	1.34***
	(4.43)	(4.31)	(4.81)	(3.94)	(4.46)	(4.3)	(4.76)	(4.03)
Controls	N	N	N	Y	N	N	N	Y
$R^2_{adj}(\%)$	15.74	19.61	20.34	30.51	17.31	19.35	20.33	30.72

TABLE B.2: **Unsmoothed PAR and Predictability**: This table repeats the baseline predictive regressions using the unsmoothed presidential approval (PAR_t) and unsmoothed approval tightness ($PART_t$). For each outcome, we estimate

$$Y_{t:t+h} = \alpha + \beta^{PAR} PAR_t + \beta^{PART} PART_t + \beta^{Control} Control_t + \varepsilon_{t:t+h},$$

and include PAR_t and $PART_t$ jointly in all specifications. Panel A reports equity excess-return regressions at monthly horizons $h \in \{1, 12\}$ and quarterly horizons $h \in \{1, 4\}$. Panel B reports quarterly GDP and PCE growth regressions at horizons $h \in \{1, 4\}$. Panel C reports monthly uncertainty regressions at monthly horizons $h = 12$ for the VIX, EPU, and the financial, macro, and real uncertainty measures of Jurado et al. (2015). For equity, $Y_{t:t+h}$ denotes cumulative excess returns from t to $t+h$; for growth, it denotes cumulative real growth from t to $t+h$; and for uncertainty, it denotes the average uncertainty level over t to $t+h$. The control variable includes the lagged dependent variable Y_t in Panels B and C and is omitted in Panel A. Newey-West (HAC) t-statistics are reported in parentheses, using six lags for the monthly one-period horizon, twelve lags for the monthly 12-month horizon, and four lags for quarterly horizons. Significance stars denote rejection of the null hypothesis at the 1%, 5%, and 10% levels.

Panel A: Equity	Monthly		Quarterly	
	h=1	h=12	h=1	h=4
PAR_t	0.06**	0.36*	0.22***	0.39*
	(1.98)	(1.84)	(2.58)	(1.87)
$PART_t$	0.32***	2.40***	1.32***	2.72***
	(2.91)	(3.4)	(3.7)	(3.39)
$R^2_{adj}(\%)$	2.01	9.44	7.82	8.38
Panel B: Economic Growth	GDP Growth		PCE Growth	
Quarterly	h=1	h=4	h=1	h=4
PAR_t	0.02***	0.02***	0.10***	0.10***
	(3.04)	(4.33)	(5.55)	(5.91)
$PART_t$	0.12***	0.12***	0.52***	0.45***
	(3.73)	(5.37)	(5.8)	(5.4)
Y_t	-0.04	-0.09	0.08	-0.01
	(-0.22)	(-0.6)	(0.32)	(-0.02)
$R^2_{adj}(\%)$	3.12	3.61	18.52	16.16

Panel C: Uncertainty	VIX	EPU	Fin U	Macro U	Real U
Monthly	h=12	h=12	h=12	h=12	h=12
PAR_t	-1.67	-1.23**	-1.48*	-2.30***	-2.28***
	(-1.21)	(-2.03)	(-1.67)	(-2.61)	(-3.19)
$PART_t$	-12.43**	-4.26*	-9.85***	-8.32***	-8.46***
	(-2.16)	(-1.78)	(-2.65)	(-2.76)	(-3.58)
Y_t	0.42***	0.54***	0.68***	0.67***	0.67***
	(6.4)	(9.77)	(12.66)	(12.13)	(17.02)
$R^2_{adj}(\%)$	44.49	51.04	65.52	67.57	65.35

TABLE B.3: **Predictability After Excluding Approval Jumps**: This table repeats the baseline predictive regressions after excluding periods associated with jumps in presidential approval (PAR_t). We detect jumps in monthly unsmoothed PAR_t following Lee and Mykland (2008). In the monthly sample, we remove the month in which a jump occurs and the subsequent five months. In the quarterly sample, we remove the quarter in which a jump occurs and the subsequent one quarter, and we classify a quarter as a jump quarter if any of its three constituent months contains a jump. For each outcome, we estimate

$$Y_{t:t+h} = \alpha + \beta^{PAR} PAR_t + \beta^{PART} PART_t + \beta^{Control} Control_t + \varepsilon_{t:t+h},$$

and include PAR_t and $PART_t$ jointly in all specifications. Panel A reports equity excess-return regressions at monthly horizons $h \in \{1, 12\}$ and quarterly horizons $h \in \{1, 4\}$. Panel B reports quarterly GDP and PCE growth regressions at horizons $h \in \{1, 4\}$. Panel C reports monthly uncertainty regressions at monthly horizons $h = 12$ for the VIX, EPU, and the financial, macro, and real uncertainty measures of Jurado et al. (2015). For equity, $Y_{t:t+h}$ denotes cumulative excess returns from t to $t+h$; for growth, it denotes cumulative real growth from t to $t+h$; and for uncertainty, it denotes the average uncertainty level over t to $t+h$. The control variable includes the lagged dependent variable Y_t in Panels B and C and is omitted in Panel A. Newey-West (HAC) t-statistics are reported in parentheses, using six lags for the monthly one-period horizon, twelve lags for the monthly 12-month horizon, and four lags for quarterly horizons. Significance stars denote rejection of the null hypothesis at the 1%, 5%, and 10% levels.

Panel A: Equity	Monthly		Quarterly	
	h=1	h=12	h=1	h=4
PAR_t	0.07**	0.45*	0.18*	0.58**
	(2.43)	(1.89)	(1.87)	(2.14)
$PART_t$	0.47***	2.54***	1.22***	3.10***
	(3.74)	(2.59)	(2.91)	(2.95)
$R^2_{adj}(\%)$	3	7.88	4.99	8.82
Panel B: Economic Growth	GDP Growth		PCE Growth	
Quarterly	h=1	h=4	h=1	h=4
PAR_t	0.03***	0.03***	0.11***	0.12***
	(2.73)	(3.2)	(4.68)	(5.51)
$PART_t$	0.15***	0.13***	0.58***	0.52***
	(3.68)	(3.75)	(4.71)	(4.41)
Y_t	-0.11	-0.17	0.05	-0.09
	(-0.74)	(-1.25)	(0.17)	(-0.26)
$R^2_{adj}(\%)$	4.6	5.06	16.86	16.16

Panel C: Uncertainty	VIX	EPU	Fin U	Macro U	Real U
Monthly	h=12	h=12	h=12	h=12	h=12
PAR_t	-1.87	-1.45**	-1.95**	-2.46***	-2.38***
	(-1.31)	(-2.27)	(-2.15)	(-2.65)	(-3.21)
$PART_t$	-12.26**	-5.07**	-10.76***	-9.51***	-8.99***
	(-2.03)	(-2.18)	(-2.66)	(-2.93)	(-3.84)
Y_t	0.44***	0.55***	0.71***	0.68***	0.70***
	(6.55)	(9.73)	(12.66)	(12.85)	(16.61)
$R^2_{adj}(\%)$	42.37	52.12	65.99	66.01	64.66