

Trade War Shocks, Inflation Compensation, and Price Stability*

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Abstract

The U.S.–China trade war has important but underexplored implications for price stability. This paper examines the effects of trade-war shocks on market-based inflation compensation using daily U.S. data from 2016 to 2026. We identify structural trade-war shocks via event-day heteroskedasticity, exploiting the sharp rise in financial-market volatility on major announcement days. In a system that includes inflation compensation, equity returns, Treasury yields, the VIX, and credit spreads, we find that adverse trade-war shocks lower inflation compensation, depress equity prices and long-term yields, raise volatility, and widen credit spreads. Decomposing inflation compensation shows that the decline is driven primarily by lower expected inflation and a residual premium component, while TIPS liquidity plays a minor role. These patterns indicate that financial markets interpret trade-war shocks mainly through weaker growth expectations and higher risk premia rather than persistent cost-push inflation. The effects vary with policy uncertainty, shock persistence, news direction, and political regimes. Overall, our results highlight how trade policy uncertainty shapes inflation compensation and macro-financial stability.

Key Words: Trade War Shocks, Inflation Compensation, Event-Day Heteroskedasticity, Trade Policy Uncertainty, Price Stability

JEL Codes: F13, E31, C32, D81, E52

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1 Introduction

Inflation compensation plays an important role in modern macro-financial analysis. Market-based measures of inflation compensation are closely monitored by central banks and financial markets because they reflect how inflation is priced in asset markets across horizons. Movements in these measures matter for monetary policy, the assessment of inflation risks, and the stability of financial conditions (Woodford, 2003; Bernanke, 2007). Understanding how major geopolitical and economic disturbances affect inflation compensation is therefore important for evaluating risks to price stability and macro-financial stability.

The return of protectionism and the escalation of U.S.-China trade tensions raise important questions for macroeconomics and finance. What do trade-war shocks imply for price stability? More specifically, through which channels are trade-war shocks transmitted to market-based inflation compensation? On the one hand, tariffs may act as cost-push shocks by raising import prices and production costs (Amiti et al., 2019; Fajgelbaum et al., 2020). On the other hand, trade-war escalations may operate as adverse demand shocks by weakening growth expectations and reducing investment and activity (Caldara et al., 2020). In addition, higher trade policy uncertainty may affect risk sentiment, financial conditions, and macroeconomic activity (Bloom, 2009; Baker et al., 2016). Whether financial markets interpret trade-war shocks mainly as supply disturbances or as demand-related risk shocks has direct implications for the behavior of inflation compensation and for the broader assessment of price stability.

Identifying the effects of trade-war shocks empirically is difficult. Standard macroeconomic data are available only at low frequency and are often confounded by simultaneous monetary, fiscal, and global shocks. Direct measures of inflation expectations, such as survey data, are informative but are usually observed only at monthly or quarterly frequency. By contrast, financial asset prices respond immediately to discrete policy announcements and contain forward-looking information at high frequency. Market-based inflation compensation therefore provides a useful lens for studying how trade-war news is transmitted into financial markets. Existing studies show that trade conflicts and trade policy uncertainty affect import prices, trade flows, investment, and output (Amiti et al., 2019; Fajgelbaum et al., 2020; Caldara et al., 2020). More broadly, the literature on uncertainty shocks emphasizes their effects on financial conditions and macroeconomic activity (Bloom, 2009; Baker et al., 2016). However, much less is known about how structurally identified trade-war shocks are transmitted into financial markets in real time, how they affect market-based inflation compensation at daily frequency, and whether these effects are concentrated at short and medium horizons or extend to longer-horizon inflation pricing. This paper fills that gap by identifying trade-war shocks using event-day heteroskedasticity and tracing their effects on inflation compensation and related financial variables across maturities. We also decompose inflation compensation into expected inflation, a TIPS

liquidity component, and a residual premium component in order to interpret the source of the baseline inflation-compensation response. In addition to the baseline analysis, the paper examines whether the effects of trade-war shocks vary with policy uncertainty, shock persistence across frequencies, the sign of trade-war news, and the political regime in which the shocks occur.

The paper makes four main contributions. It provides daily-frequency evidence on how structurally identified trade-war shocks affect market-based inflation compensation across the two-year, five-year, and ten-year maturities. It then places these inflation-compensation responses within a broader macro-financial transmission mechanism by jointly examining equity returns, long-term Treasury yields, volatility, and corporate credit spreads. The paper also studies whether the transmission of trade-war shocks varies with the surrounding uncertainty environment, the persistence of the shock, the direction of trade-war news, and the political regime in which the shock occurs. Finally, it decomposes inflation compensation into model-implied expected inflation, a transaction-volume-based TIPS liquidity component, and a residual premium component, and compares the estimated response with external benchmarks, including a daily oil-price benchmark and an externally identified high-frequency monetary-policy shock. These contributions allow the paper to assess whether trade-war shocks are priced mainly as inflationary cost-push shocks or as weaker-growth and risk-premium shocks, and whether their magnitude is relevant for price stability.

This paper studies the effects of trade-war shocks on market-based inflation compensation using daily U.S. financial data. We exploit regime-dependent heteroskedasticity on major U.S.-China trade announcement days to identify structural trade-war shocks. Our approach builds on identification through heteroskedasticity (Rigobon, 2003) and is closely related to recent work that applies event-day heteroskedasticity to trade-war shocks in financial markets (Chen et al., 2023). More broadly, it is also connected to the literature on high-frequency identification in macro-finance (Gürkaynak et al., 2005; Nakamura and Steinsson, 2018). The key identifying assumption is that the variance of the trade-war shock rises on announcement days relative to non-event days, while the variances of other structural innovations remain stable. This regime-specific variance shift implies an approximately rank-one change in the covariance matrix of reduced-form residuals, which allows us to identify the contemporaneous impact vector of the trade-war shock without imposing recursive ordering restrictions. Building on this baseline identification, we then examine several extensions that help sharpen the interpretation of the baseline results, including a decomposition of inflation compensation into expected inflation, TIPS liquidity, and residual premium components, state dependence with respect to trade policy uncertainty and economic policy uncertainty, a frequency-based decomposition of the identified trade-war shock into lower- and higher-frequency components, asymmetric responses to negative and positive trade-war news, and heterogeneity across political regimes.

Using daily U.S. data from 2016 to 2026, we estimate a six-variable system including five-year inflation

compensation, equity returns, long-term nominal Treasury yields, the VIX, investment-grade corporate credit spreads, and high-yield corporate credit spreads. We interpret five-year inflation compensation as a market-based measure of medium-horizon inflation compensation rather than as a pure measure of expected inflation, since breakeven rates may reflect expected inflation, inflation risk premia, and liquidity effects. We further decompose five-year inflation compensation to examine whether the baseline response is driven by expected inflation, TIPS liquidity, or residual premium movements. By examining the joint responses of inflation compensation and other macro-financial variables, we assess whether trade-war shocks are transmitted mainly through a cost-push channel or through weaker growth expectations and higher risk premia.

We summarize our results in three main findings. First, adverse trade-war shocks generate economically meaningful declines in inflation compensation, particularly at short and medium horizons, while depressing equity prices, lowering long-term nominal yields, raising volatility, and widening credit spreads. The joint pattern across variables suggests that financial markets interpret trade-war shocks primarily as weaker growth expectations and higher risk premia rather than persistent cost-push inflationary pressures. Second, the transmission is heterogeneous. The effects vary systematically with the level of policy uncertainty, the frequency (high- versus low-frequency components) of the shock, the sign of trade-war news, with negative news producing stronger and more disruptive responses, and the political regime in which the shocks occur. Third, the decomposition analysis shows that the decline in inflation compensation is driven mainly by revisions in expected inflation and movements in the residual premium component, while the TIPS liquidity component plays only a minor role. This evidence reinforces the view that trade-war shocks affect inflation compensation predominantly through weaker growth expectations and elevated risk premia.

These results highlight the macro-financial transmission of trade policy uncertainty and its implications for the behavior of market-based inflation compensation. They show that geopolitical trade shocks can exert disinflationary pressure through weaker growth expectations, residual premium movements, and tighter financial conditions, even when tariffs mechanically raise import prices. More broadly, the findings underscore the importance of understanding how financial markets price trade policy shocks when assessing risks to price stability. They also show that the effects of trade-war shocks are not uniform, but depend on the persistence of the shock, the direction of trade-war news, the surrounding uncertainty environment, and the broader political setting.

The rest of the paper proceeds as follows. Section 2 reviews the related literature. Section 3 describes the identification strategy and empirical framework, including the decomposition and extension exercises. Section 4 presents the baseline empirical results and the decomposition of the inflation-compensation response. Section 5 reports the extension results. Section 6 discusses the broader implications for price stability and policy. Section 7 concludes.

2 Related Literature

This paper lies at the intersection of three bodies of research: the effects of trade policy shocks and trade policy uncertainty, the formation and anchoring of inflation expectations, and the identification of structural shocks using financial-market data. These literatures are connected by a common issue. Trade-war announcements do not simply alter current tariff rates. They also change expectations about future policy, economic activity, inflation, and financial risk.

Research on the 2018–2019 U.S.–China trade conflict documents substantial effects on import prices, welfare, trade flows, investment, and output (Amiti et al., 2019; Fajgelbaum et al., 2020; Caldara et al., 2020). A growing part of this literature focuses on the uncertainty created by an unstable trade-policy regime. Benguria et al. (2022) construct firm-level measures of trade policy uncertainty from Chinese listed firms’ annual reports and find that greater exposure reduced investment, RD, and profits. Poilly and Tripier (2025) show that U.S. states more exposed to tariff volatility experience larger declines in output and employment. Taking a forward-looking approach, Alessandria et al. (2025) model trade policy as a Markov process and document substantial changes in beliefs about the expected duration of U.S. tariffs on China. This evidence supports treating trade-war developments as expectation and uncertainty shocks rather than as contemporaneous tariff changes alone.

Trade restrictions can also interact with monetary policy, production networks, and governments’ strategic objectives.¹ Their financial-market effects are therefore potentially broader than their direct influence on traded-goods prices. Recent evidence finds asymmetric stock-market responses to tariff increases and reductions, persistent effects of tariffs and trade policy uncertainty on major U.S. equity indices, and heterogeneous firm-level losses around the 2025 “Liberation Day” announcement.² However, this work largely concentrates on equity markets, firms, and real activity. Much less is known about how trade-war news affects the market pricing of inflation.

The second relevant literature studies inflation expectations and their importance for price stability. Expectations occupy a central role in models of inflation dynamics and monetary-policy transmission (Woodford, 2003; Galí, 2015). Empirical research has documented disagreement and information frictions in survey-based expectations (Mankiw et al., 2004; Coibion et al., 2020). More recent studies show that households, firms, and professional forecasters may form expectations differently and respond unevenly to macroeconomic news.³ The anchoring of expectations can also change with the policy environment. Bonomo et al.

¹Auray et al. (2025) show that monetary-policy rules affect the equilibrium outcome of tariff competition in a New Keynesian open-economy model. de Souza et al. (2024) examine the cost-efficient design of trade sanctions in a quantitative model with input-output linkages.

²See Gaies (2025), Yilmazkuday (2025), and Coccozza and Gallo (2025).

³Hajdini et al. (2024) introduce an indirect measure of consumer inflation expectations. Binder et al. (2024) use daily survey timing to study responses to FOMC announcements and macroeconomic releases. Candia et al. (2024) document weakly

(2024) show that an abrupt monetary-policy shift can rapidly unanchor expectations, while Crump et al. (2024) emphasize the role of expected labor-market conditions in restoring price stability after the pandemic. López-Salido and Loria (2024) further show that expectations and financial conditions affect the tails as well as the mean of the inflation distribution.

Evidence on nonlinear expectation dynamics is particularly relevant for our extension analysis. Oil-price shocks can affect inflation expectations asymmetrically, and the strength of this relationship varies across countries and monetary-policy environments (Nasir et al., 2020b,a; Nasir and Huynh, 2024). These findings make it inappropriate to assume that favorable and adverse trade-war news must produce equal and opposite responses or that transmission remains constant across policy regimes.

Our outcome variable belongs to the related literature on market-based inflation compensation. Measures derived from nominal Treasury and TIPS yields provide high-frequency information about inflation pricing, but they are not pure measures of expected inflation.⁴ Gürkaynak et al. (2010) show that inflation compensation combines expected inflation, inflation risk premia, and TIPS liquidity premia, and that its high-frequency variation is too large to represent revisions to expected inflation alone. D’Amico et al. (2018) and Kim et al. (2019) provide further evidence on the importance of inflation-risk and liquidity components. Abrahams et al. (2016) estimate a joint nominal–real term-structure model that adjusts for TIPS illiquidity and separates expected inflation, liquidity, and inflation-risk compensation. This literature motivates our reduced-form decomposition into model-implied expected inflation, a transaction-volume-based TIPS liquidity component, and a residual premium component. The last component is not treated as a pure inflation risk premium because it may also contain tax effects, embedded options, measurement error, and other unidentified wedges.

The empirical design draws on the literature identifying structural shocks from financial-market variation. Identification through heteroskedasticity uses changes in shock volatility across regimes to recover structural disturbances without imposing a recursive ordering (Rigobon, 2003). High-frequency approaches have similarly used narrow announcement windows to identify monetary-policy shocks (Gürkaynak et al., 2005; Nakamura and Steinsson, 2018). Bianchi et al. (2023) exploit the precise timing of presidential tweets about the Federal Reserve to identify the effects of political pressure on monetary-policy expectations and asset prices. Most closely related to our approach, Chen et al. (2023) use event-day heteroskedasticity to identify trade-related shocks in financial markets.

The unresolved question is how a structurally identified trade-war shock is transmitted into market-based inflation compensation in real time. Existing studies establish that trade conflicts affect real activity, firms,

anchored expectations among U.S. firms.

⁴See Ang et al. (2008), Christensen et al. (2010), Haubrich et al. (2012), and Hördahl and Tristani (2012).

and financial markets, while the TIPS literature shows that inflation compensation combines expectations, liquidity, and risk compensation. They do not establish whether trade-war announcements are priced mainly as inflationary cost shocks or as shocks to expected growth and risk premia. We address this question using a joint system of inflation compensation, equity returns, Treasury yields, volatility, and credit spreads. The analysis also examines whether the response varies across maturities, uncertainty environments, frequency components, news direction, and political regimes.⁵

3 Methodology

Our empirical strategy combines identification through heteroskedasticity (Rigobon, 2003), event-day identification of trade-related financial shocks (Chen et al., 2023), and local projections (Jordà, 2005). Relative to the existing literature, the main new element is the application of event-day heteroskedasticity to market-based inflation compensation within a joint macro-financial system. The system allows the identified trade-war disturbance to be interpreted from the simultaneous movements in inflation compensation, equity returns, Treasury yields, volatility, and corporate credit spreads. We also examine whether its transmission varies across maturities, uncertainty environments, frequency components, news direction, and political regimes, and use a decomposition of inflation compensation to clarify the underlying channels.

The empirical analysis has three main stages. We first estimate a reduced-form vector autoregression and exploit the change in residual covariance on trade-war event days to identify the contemporaneous impact vector of the trade-war shock. We then construct a standardized daily shock series and estimate its dynamic effects using local projections. Finally, we conduct a set of mechanism and heterogeneity exercises. The technical implementation of these extensions is reported in Appendix Section A.

3.1 Baseline Financial System

Let y_t denote the vector of daily financial variables,

$$y_t = \left(\Delta IC_t, R_t, \Delta Y_t^{(10)}, \Delta VIX_t, \Delta IG_t, \Delta HY_t \right)', \quad (1)$$

where ΔIC_t denotes the change in market-based inflation compensation, R_t denotes the S&P 500 return, $\Delta Y_t^{(10)}$ denotes the change in the ten-year Treasury yield, and ΔVIX_t , ΔIG_t , and ΔHY_t denote changes in the VIX, investment-grade credit spreads, and high-yield credit spreads, respectively.

⁵The frequency and state-dependent exercises are also related to Guo et al. (2022), who document time-varying and frequency-dependent spillovers across monetary, fiscal, and trade policy uncertainty.

The reduced-form VAR of order p is

$$y_t = c + \sum_{\ell=1}^p B_\ell y_{t-\ell} + Dz_t + e_t, \quad (2)$$

where c is a vector of intercepts, B_ℓ are coefficient matrices, z_t contains optional exogenous controls, and e_t is the vector of reduced-form innovations. The reduced-form innovations are related to orthogonal structural shocks through

$$e_t = Au_t, \quad (3)$$

where A is the contemporaneous impact matrix and

$$E(u_t u_t') = I. \quad (4)$$

The structural disturbance of interest is the shock associated with adverse trade-war developments.

3.2 Identification through Event-Day Heteroskedasticity

Identification through heteroskedasticity exploits changes in shock volatility across regimes without imposing a recursive ordering on the contemporaneous relationships among financial variables (Rigobon, 2003). Our implementation follows the event-day logic used by Chen et al. (2023), but differs in its focus on inflation compensation and in the broader set of asset prices used to interpret the identified shock.

Let $\mathcal{T}_0$ denote trade-war event days and $\mathcal{T}_1$ denote non-event days. The regime-specific covariance matrices of the reduced-form innovations are

$$R_s = E[e_t e_t' \mid t \in \mathcal{T}_s], \quad s \in \{0, 1\}. \quad (5)$$

Using Equation (3),

$$R_s = A \Sigma_s A', \quad (6)$$

where Σ_s is the covariance matrix of structural shocks in regime s .

The identifying assumption is that the variance of the trade-war shock increases on event days, while the variances of the other structural shocks remain stable. If a_1 denotes the contemporaneous impact vector of the trade-war shock, the population covariance difference satisfies

$$R_0 - R_1 = \tilde{a}_1 \tilde{a}_1', \quad (7)$$

where

$$\tilde{a}_1 = \sigma_{TW} a_1, \quad (8)$$

and σ_{TW}^2 measures the event-day increase in the variance of the trade-war shock. Equation (7) implies that the covariance difference is rank one in population.

We estimate the impact vector by fitting the rank-one restriction to the sample covariance difference using minimum-distance GMM:

$$\hat{\tilde{a}}_1 = \arg \min_{\tilde{a}_1} g(\tilde{a}_1)' W g(\tilde{a}_1), \quad (9)$$

where

$$g(\tilde{a}_1) = \text{vech} \left(\hat{R}_0 - \hat{R}_1 - \tilde{a}_1 \tilde{a}_1' \right), \quad (10)$$

and W is a positive-definite weighting matrix. In the empirical implementation, W is based on the estimated covariance matrix of the regime-specific covariance moments.

The impact vector is identified up to sign and scale. We normalize its sign so that a positive shock represents an adverse trade-war development, indicated by an increase in the VIX and, when required as a fallback restriction, a decline in equity prices. The impact vector is then normalized to unit length.

After estimating the impact direction, we construct the daily trade-war shock as

$$w_t = \hat{a}_1' \hat{e}_t, \quad (11)$$

where $\hat{e}_t$ denotes the estimated reduced-form innovation. The resulting shock series is standardized to have zero mean and unit variance.

3.3 Dynamic Responses

We estimate dynamic responses using local projections following Jordà (2005). For each variable $y_{i,t}$ and horizon h , the baseline specification is

$$\begin{aligned} y_{i,t+h} - y_{i,t-1} = & \alpha_{i,h} + \beta_{i,h} w_t + \gamma_{i,h} w_{t-1} + \delta_{i,h} y_{i,t-1} \\ & + \eta_{i,h}' z_t + \varepsilon_{i,t+h}. \end{aligned} \quad (12)$$

The coefficient $\beta_{i,h}$ measures the response of variable i at horizon h to a one-standard-deviation adverse trade-war shock. Statistical inference uses Newey–West standard errors with horizon-dependent lag lengths. The figures report 68% and 90% confidence bands.

3.4 Heterogeneity and Persistence

We examine whether the transmission of the identified shock depends on the environment in which trade-war news arrives. The state-dependent analysis interacts the shock with lagged standardized measures of trade policy uncertainty and economic policy uncertainty. These variables are treated as conditioning states rather than as separately identified uncertainty shocks. Using lagged uncertainty reduces the possibility that the state variable captures the contemporaneous response to the trade-war announcement itself.

We also separate the identified shock into lower- and higher-frequency components:

$$w_t = w_t^{LF} + w_t^{HF}. \quad (13)$$

The higher-frequency component captures shorter-lived announcement effects, while the lower-frequency component captures more persistent repricing. The decomposition follows the frequency-domain intuition in Baruník and Ellington (2024). The filter construction, cutoff definition, and inverse transformation are reported in Appendix Section A.2.

Political-regime heterogeneity is examined by interacting the common full-sample shock with indicators for the Trump I, Biden, and Trump II administrations. This preserves a unified shock measure while allowing its propagation to vary across political environments. Because the later regimes contain relatively few trade-war event days, these estimates are interpreted as suggestive evidence of heterogeneous transmission.

Directional asymmetry is studied using the broader Trade War Event Database, which classifies events as adverse or favorable trade-war developments. We identify a common shock from the full event set and then split it by event direction:

$$s_t^- = s_t \mathbf{1}\{\text{negative event at } t\}, \quad s_t^+ = s_t \mathbf{1}\{\text{positive event at } t\}. \quad (14)$$

Estimating both components in the same local projection allows negative and positive news to have different magnitudes and persistence without requiring them to produce mirror-image responses. The complete specifications for the state-dependent, regime-dependent, and asymmetric local projections are reported in Appendix Sections A.1–A.4.

3.5 Inflation-Compensation Decomposition

Breakeven inflation is not a pure measure of expected inflation because it may also reflect inflation risk compensation and TIPS liquidity conditions (Gürkaynak et al., 2010; D’Amico et al., 2018; Kim et al., 2019;

Abrahams et al., 2016). We therefore use the accounting relationship

$$IC_t(h) = EI_t(h) + RPC_t(h) - LP_t(h), \quad (15)$$

where $EI_t(h)$ denotes model-implied expected inflation, $LP_t(h)$ denotes the estimated TIPS liquidity-premium component, and $RPC_t(h)$ denotes the residual premium component.

The expected-inflation component is obtained from a monthly CPI forecasting system that uses information from nominal and real yield curves. The fitted system is evaluated at daily financial-market states to construct a daily expected-inflation series. The liquidity component is based on relative primary-dealer transaction volumes in nominal Treasury securities and TIPS. The residual premium component is defined as

$$\widehat{RPC}_t(h) = IC_t(h) + \widehat{LP}_t(h) - \widehat{EI}_t(h). \quad (16)$$

The residual component should not be interpreted as a pure inflation risk premium because it may also contain tax effects, embedded option effects, measurement error, and other wedges not separately identified by the reduced-form procedure.

The decomposed components are entered separately into the baseline local-projection framework. The full construction of the yield-curve factors, expected-inflation projections, transaction-volume liquidity proxy, and component-level local projections is provided in Appendix Section A.5.

3.6 Identification Diagnostics and Transmission Channels

The identifying restriction requires a meaningful event-day covariance shift that is approximately rank one. We therefore examine the eigenvalues of $\widehat{R}_0 - \widehat{R}_1$, the share of positive covariance-difference variation explained by the leading component, and the event-to-non-event variance ratio of the identified shock. The exact diagnostic definitions are reported in Appendix Section A.6.

The joint responses across financial variables help distinguish between two non-mutually exclusive channels. A cost-push interpretation would generally imply upward pressure on inflation compensation and nominal yields. By contrast, weaker expected demand and higher risk premia would be reflected in lower inflation compensation, equity prices, and Treasury yields, together with higher volatility and wider credit spreads. We therefore interpret the identified shock from the sign, magnitude, and persistence of responses across the system rather than from inflation compensation alone.

3.7 Data

The sample contains daily U.S. trading-day observations from 2016 to 2026. The baseline system includes two-year, five-year, or ten-year inflation compensation, S&P 500 returns, changes in the ten-year Treasury yield, changes in the VIX, and changes in investment-grade and high-yield credit spreads. The inflation-compensation series are obtained from the Federal Reserve Board’s TIPS Yield Curve and Inflation Compensation data. The term-premium extension uses the ten-year Adrian–Crump–Moench Treasury term premium.

The decomposition additionally uses nominal Treasury yields, TIPS yields, non-seasonally adjusted CPI-U, and primary-dealer Treasury transaction volumes. Trade-war event days are constructed from a hand-collected timeline of major U.S.–China trade announcements and related policy actions. The state-dependent analysis uses the U.S. Daily Economic Policy Uncertainty Index developed by Baker et al. (2016) and the news-based Trade Policy Uncertainty Index developed by Caldara et al. (2020).

The use of daily financial data is central to the empirical design. Standard macroeconomic variables and survey-based inflation expectations are generally observed at lower frequencies and are therefore less suitable for identifying the immediate transmission of trade-war announcements. Daily financial-market prices allow us to measure how trade-war news is incorporated into inflation compensation and broader financial conditions in real time.

4 Baseline Results

4.1 Identification Diagnostics

Before turning to the dynamic responses, we first assess whether the identifying assumption underlying the event-day heteroskedasticity approach is supported by the data.

Table 1 reports the main identification diagnostics for the baseline five-variable inflation-compensation specification. The covariance matrix of reduced-form residuals differs systematically between trade-war event days and non-event days. In particular, the covariance-difference matrix exhibits a pronounced concentration in its leading eigenvalue. The first eigenvalue is large, while the remaining eigenvalues are close to zero and mostly slightly negative, as is common in finite samples when a population rank-one restriction is approximated with sample moments. The share of the positive spectrum explained by the first component is effectively equal to one. This pattern is consistent with the view that the regime-dependent variance shift is dominated by a single structural disturbance associated with trade-war news. For completeness, analogous diagnostics for the two-year and ten-year specifications are reported in Appendix Tables 18 and 19, and they

display the same broad near-rank-one pattern.

Table 1: Identification Diagnostics: Event-Day Heteroskedasticity

Statistic	Value
Event days (residual sample)	76
Non-event days (residual sample)	2105
Largest eigenvalue of $R_0 - R_1$	4.3907
Second eigenvalue of $R_0 - R_1$	-0.0000
Eigenvalue ratio $\lambda_1/ \lambda_2 $	2600039.81
Share of first positive eigenvalue	1.0000
Frobenius norm $\ R_0 - R_1\ _F$	4.3907
Randomization p -value (variance shift)	0.1024
Randomization p -value (leading eigenvalue)	0.1024
Variance ratio of identified shock	2.1179
Randomization p -value (variance ratio)	0.1024

Notes: R_0 and R_1 denote covariance matrices of reduced-form residuals on trade-war event and non-event days, respectively. The baseline system includes five-year inflation compensation, equity returns, changes in the ten-year Treasury yield, changes in the VIX, and changes in investment-grade and high-yield credit spreads. Randomization tests are based on reshuffling event labels while preserving the number of event days.

The residual sample contains 76 trade-war event days and 2,105 non-event days. The leading eigenvalue of the covariance-difference matrix is 4.3907, whereas the second eigenvalue is essentially zero in magnitude. The corresponding eigenvalue ratio is, therefore, extremely large, indicating that the variance shift is highly concentrated in one dimension. The Frobenius norm of the covariance difference is also almost identical to the leading eigenvalue, further reinforcing the near rank-one interpretation.

To assess whether this variance shift could arise mechanically under arbitrary partitions of the data, we conduct randomization tests that reshuffle event labels while preserving the number of event days. The resulting placebo p -values are around 0.10 for the overall covariance shift, the leading eigenvalue, and the variance ratio of the identified shock. These values indicate that the observed heteroskedasticity is not overwhelmingly significant under permutation inference, but it is nonetheless unusual relative to randomized partitions and economically meaningful in magnitude.

We also examine the variance of the identified trade-war shock across regimes. The variance of the shock on event days is about 2.12 times as large as on non-event days. This increase confirms that the identified shock becomes substantially more volatile on major trade-war announcement days, even if the statistical evidence is somewhat weaker than in the earlier specification.

The diagnostics indicate that the data are broadly consistent with the identifying assumption of a dominant event-specific variance shift. The covariance difference is strongly concentrated in a single dimension, and the identified shock is materially more volatile on trade-war event days than on non-event days. While the permutation evidence is more suggestive than definitive, the overall pattern supports the use of event-day heteroskedasticity as the basis for identifying the structural trade-war shock. The appendix results for

the two-year and ten-year systems reinforce this conclusion by showing that the same broad identification pattern is not specific to the five-year specification.

4.2 Baseline Impulse Responses

Figure 1 presents the local-projection impulse responses of the baseline system to the identified trade-war shock. Several features stand out.

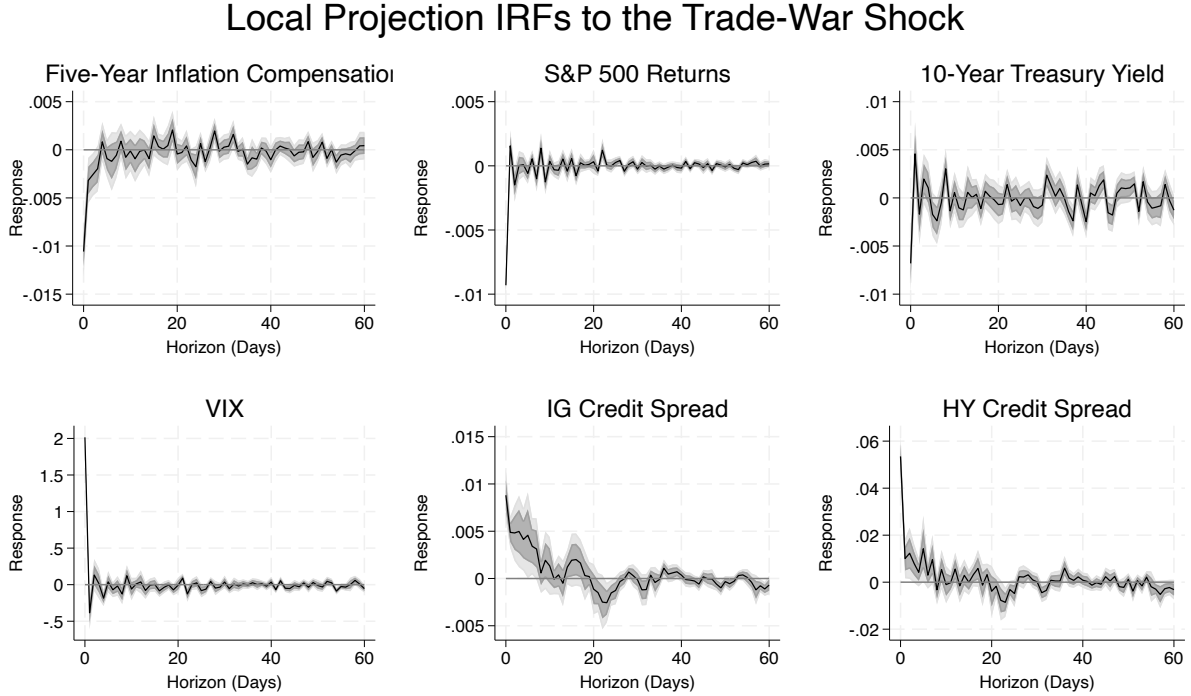


Figure 1: Impulse responses to the identified trade-war shock. The black lines show the estimated responses, and the shaded areas denote 68% and 90% confidence bands.

First, five-year inflation compensation declines on impact following the shock. The initial response is negative and economically meaningful, consistent with the view that financial markets interpret trade-war shocks as weakening growth prospects and tightening financial conditions rather than as generating persistent inflationary pressure. But it dissipates relatively quickly over subsequent trading days. This pattern indicates that trade-war news leads financial markets to mark down medium-horizon inflation compensation rather than to price in persistent inflationary pressure.

Second, equity returns fall on impact, consistent with a deterioration in risk sentiment and weaker expected economic activity. The decline is concentrated at short horizons and fades relatively quickly, suggesting that the identified shock is primarily associated with an immediate repricing of growth-sensitive

assets.

Third, the ten-year Treasury yield also declines following the shock. This response is consistent with a flight-to-safety mechanism and with downward revisions to the medium-run macroeconomic outlook. The response of long-term yields is more modest than that of volatility and credit spreads, but its sign is informative about how markets interpret trade-war escalations.

Fourth, financial risk indicators move in the opposite direction. The VIX rises sharply on impact, pointing to a sudden increase in uncertainty and risk aversion. Corporate credit spreads also widen, with the response being more pronounced for high-yield bonds than for investment-grade bonds. This cross-sectional pattern suggests that trade-war shocks trigger a tightening in financial conditions and a repricing of credit risk, especially for riskier borrowers.

The joint responses across asset classes suggest that financial markets interpret trade-war shocks primarily as adverse demand disturbances accompanied by elevated risk premia, rather than as cost-push inflationary shocks. If trade-war news were mainly perceived as a persistent supply-side inflation shock, one would expect inflation compensation and nominal yields to rise, reflecting higher expected production costs and inflationary pressure (Amiti et al., 2019; Fajgelbaum et al., 2020). Instead, the observed decline in five-year inflation compensation, together with falling equity prices and Treasury yields and rising volatility and credit spreads, points to weaker growth expectations and tighter financial conditions as the dominant transmission channels. This interpretation is consistent with the broader literature on uncertainty shocks and trade policy uncertainty. Bloom (2009) shows that increases in uncertainty can depress investment and activity, while Caldara et al. (2020) document that trade policy uncertainty has adverse effects on real economic outcomes. In this context, the responses of financial variables are consistent with a mechanism in which trade-war announcements raise uncertainty and risk premia, weaken growth expectations, and reduce market-based inflation compensation.

4.3 Economic Magnitude and Implications for Inflation Compensation

We next assess the economic magnitude of the identified trade war shock for market based inflation compensation and related financial variables.

A one standard deviation increase in the identified trade war shock lowers five year inflation compensation on impact by an economically meaningful amount. In daily financial data, a movement of this size reflects a noticeable downward repricing of medium horizon inflation compensation within a single trading day. Although the response becomes smaller over the following days, the short run cumulative effect remains negative. This suggests that the initial reassessment is not immediately reversed.

The importance of the shock is also visible in its joint effects across asset classes. The decline in five year

inflation compensation is accompanied by lower equity returns, a decline in the ten year Treasury yield, a sharp increase in the VIX, and wider corporate credit spreads, especially in the high yield segment. This pattern suggests that the identified shock is not a small or isolated movement in a single market. Instead, it is associated with a broader repricing across financial variables that are closely related to growth expectations, risk sentiment, and financial conditions.

These results suggest that trade war shocks do not generate persistent increases in market based inflation compensation. Instead, they lead to short run declines in inflation compensation alongside tighter financial conditions and weaker risk appetite. This pattern is more consistent with weaker demand and higher risk premia than with sustained cost push inflationary pressure.

4.4 Transmission Channels

We next examine the economic channels through which trade war shocks affect inflation compensation and broader financial conditions. The joint responses discussed above help us distinguish between alternative interpretations of these shocks.

A cost push interpretation of tariffs would imply an increase in inflation compensation, reflecting higher expected input costs and upward pressure on prices. Under this view, inflation compensation and nominal yields would be expected to rise after trade war announcements. The impulse responses provide little support for this mechanism. Instead, inflation compensation falls on impact and remains weak in the short run. Long term nominal yields also decline. This pattern is difficult to reconcile with an inflationary cost shock and suggests that markets do not view trade war escalations as a source of sustained upward pressure on prices.

The evidence is more consistent with a growth channel combined with a financial risk channel. The simultaneous decline in inflation compensation and long term yields suggests that changes in expected economic activity matter alongside changes in risk premia. The broader pattern of responses points to a mechanism in which trade war shocks weaken growth expectations while also increasing financial stress. Equity returns fall and long term yields decline after the shock, both of which are consistent with downward revisions to future economic activity. When growth prospects weaken, medium horizon inflation compensation is also likely to decline. The joint fall in inflation compensation and equity prices therefore suggests an important role for weaker demand expectations in the transmission mechanism.

At the same time, financial conditions tighten. Volatility rises sharply, and credit spreads widen, especially in the high yield segment. These responses suggest greater risk aversion and a higher premium for bearing credit risk. The stronger reaction of high yield spreads relative to investment grade spreads indicates that this repricing is more pronounced for riskier borrowers. This pattern is consistent with a meaningful risk premium component in asset prices during periods of trade policy uncertainty.

The evidence suggests that trade war shocks operate primarily through weaker demand and higher risk premia rather than through persistent cost driven inflationary pressure. Inflation compensation declines together with measures of economic sentiment and risk appetite. In this sense, the implications of trade war shocks for price stability appear to come mainly from weaker growth expectations and tighter financial conditions, rather than from a persistent increase in inflation compensation.

4.5 Term Premium Decomposition

To better understand the decline in long term nominal yields after trade war shocks, we examine the response of the ten year term premium using the Adrian, Crump, and Moench measure.

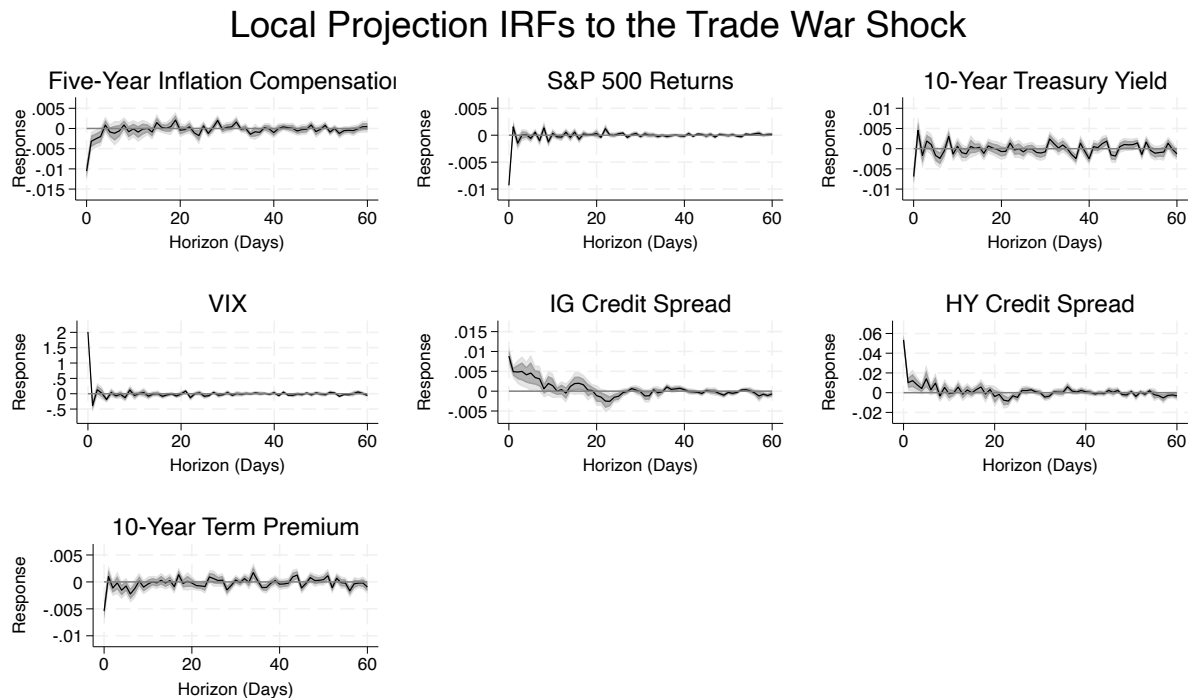


Figure 2: Local projection impulse responses from the seven variable system that augments the baseline specification with the ACM ten year term premium. The black lines denote point estimates, and the shaded areas represent 68% and 90% confidence bands.

Figure 2 shows that the ten year term premium falls on impact after the identified trade war shock. The decline is short lived, however, and the response moves back toward zero relatively quickly over subsequent trading days. This pattern suggests that term premium compression is part of the initial adjustment in long term bond yields, but it does not appear to be highly persistent.

The response of the ten year Treasury yield is also negative at short horizons. These responses suggest

that the decline in nominal yields reflects, at least in part, a flight to safety mechanism. Trade war shocks appear to increase demand for safe, long duration assets, which lowers the compensation investors require for bearing duration risk.

At the same time, the brief nature of the term premium response suggests some caution in interpretation. The evidence is consistent with a role for term premium compression in the initial decline in long term yields, but it does not imply that this channel accounts for the full yield response over time.

Taken together with the decline in inflation compensation in the baseline results, the term premium evidence remains consistent with the broader view that trade war shocks are associated with weaker growth expectations and tighter financial conditions, rather than with persistent cost push inflationary pressure.

4.6 Horizon Comparison Across Inflation Compensation Measures

We next compare the responses of inflation compensation across maturities to assess whether trade war shocks have different effects at short, medium, and long horizons.

Figure 3 presents the responses obtained when five year inflation compensation is replaced with the two year breakeven rate, while Figure 4 reports the corresponding results for the ten year breakeven rate.

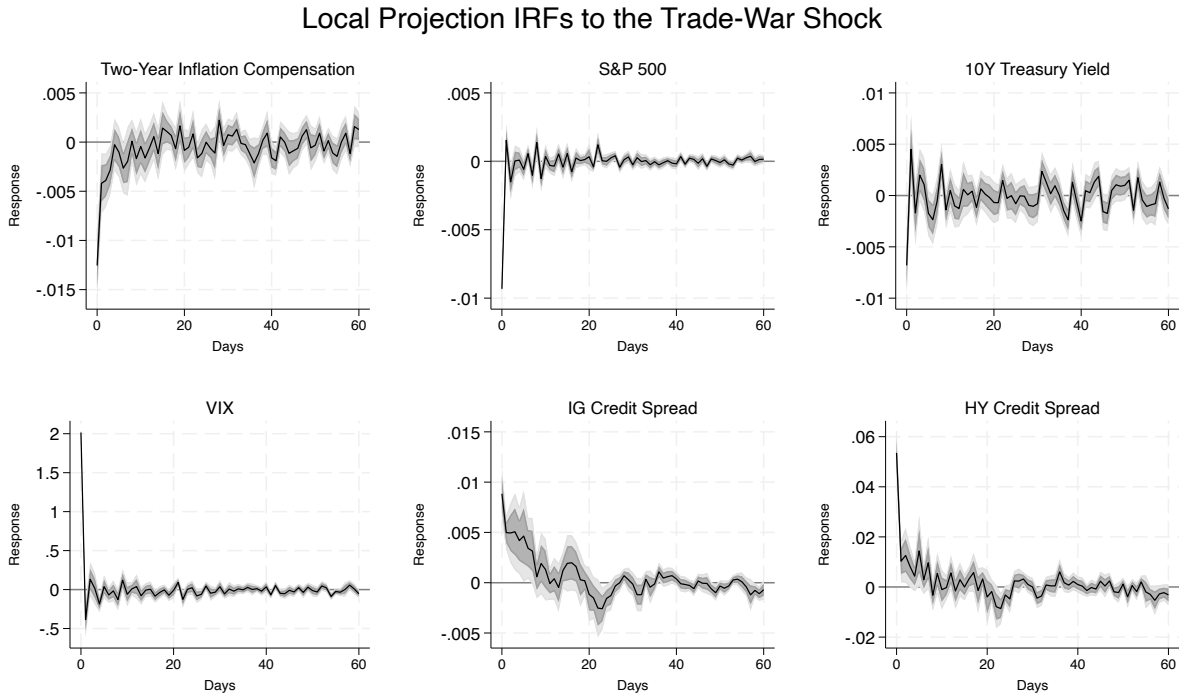


Figure 3: Impulse responses to the identified trade war shock using two year inflation compensation. The shaded areas represent confidence bands.

Local Projection IRFs to a Trade-War Shock

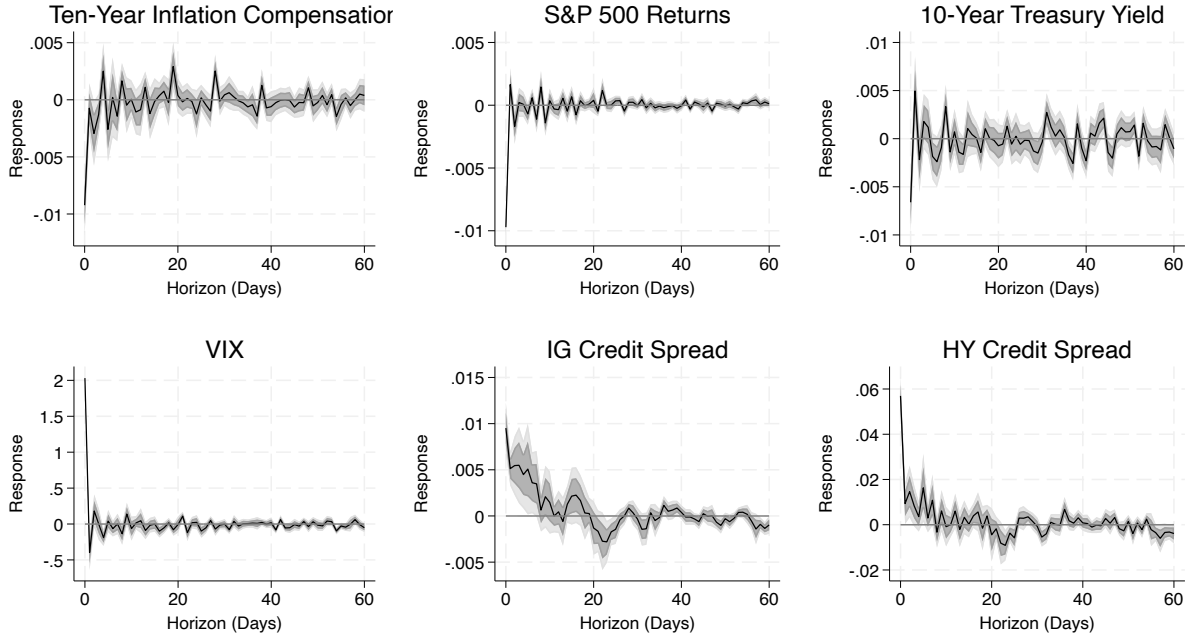


Figure 4: Impulse responses to the identified trade war shock using ten year inflation compensation. The shaded areas represent confidence bands.

The responses show a clear maturity pattern. Two year inflation compensation declines sharply on impact and then returns toward zero relatively quickly. By contrast, the response of ten year inflation compensation is also negative on impact, but it is smaller in magnitude and remains economically modest throughout. The five year baseline response lies between these two cases, with a larger decline than at the ten year horizon and a more pronounced short run response than is observed at the long end.

This pattern suggests that trade war shocks are priced primarily as short to medium run disinflationary forces rather than as shocks that materially alter long horizon inflation compensation. The stronger response at the two year horizon is consistent with a revision in near term macroeconomic conditions and risk sentiment. The more muted response at the ten year horizon suggests that longer run inflation pricing remains comparatively stable.

The broader financial responses remain similar across specifications. Equity returns fall on impact, long term Treasury yields weaken, volatility rises, and credit spreads widen. The evidence suggests that trade war shocks are associated with weaker growth expectations and tighter financial conditions across maturities, while their effect on long horizon inflation compensation remains limited.

From the perspective of price stability, these results suggest that trade war shocks can generate short

run downward pressure on inflation compensation without producing a meaningful deterioration in longer run inflation anchoring. In that sense, the main effects appear to operate through cyclical repricing at short and medium horizons rather than through persistent changes in long run inflation compensation.

5 Extension Results

5.1 State Dependence: Trade Policy Uncertainty and Economic Policy Uncertainty

We next examine whether the transmission of identified trade-war shocks varies with the uncertainty environment. Since trade-war shocks are naturally related to trade policy uncertainty, this exercise is not intended to identify an independent uncertainty shock. Instead, we interpret TPU and EPU as conditioning variables that describe the uncertainty environment in which trade-war news is priced by financial markets. To reduce simultaneity concerns, we condition on lagged standardized measures of trade policy uncertainty and economic policy uncertainty. Specifically, we interact the identified trade-war shock with lagged TPU and EPU and compute the implied impulse responses under low- and high-uncertainty states, defined as one standard deviation below and above the mean of the corresponding uncertainty measure.

Figure 5 reports the state-dependent responses based on lagged TPU. The high- and low-TPU responses are broadly similar in direction across all variables. In both states, trade-war shocks lower five-year inflation compensation on impact, depress equity returns, raise volatility, and widen corporate credit spreads. However, the two sets of responses do not display a sharp or persistent separation. Equity returns, Treasury yields, and the VIX show little evidence of systematic state dependence. Some differences are visible in corporate credit spreads, especially at short horizons, but the overall pattern does not support a strong amplification effect under high TPU. Instead, the results suggest that the baseline transmission of trade-war shocks is broadly robust across trade-policy uncertainty states.

Figure 6 reports the corresponding results using lagged EPU as the conditioning variable. The evidence for state dependence is even weaker. The high- and low-EPU responses remain close for most variables and horizons, including inflation compensation, equity returns, Treasury yields, and volatility. Some modest differences are visible in corporate credit spreads, but these differences are not sufficiently systematic to indicate a strong state-dependent transmission mechanism. This is consistent with the broader nature of EPU relative to TPU: while EPU captures general policy uncertainty, it is less directly tied to trade-war news.

The results indicate that the uncertainty environment does not fundamentally alter the transmission of

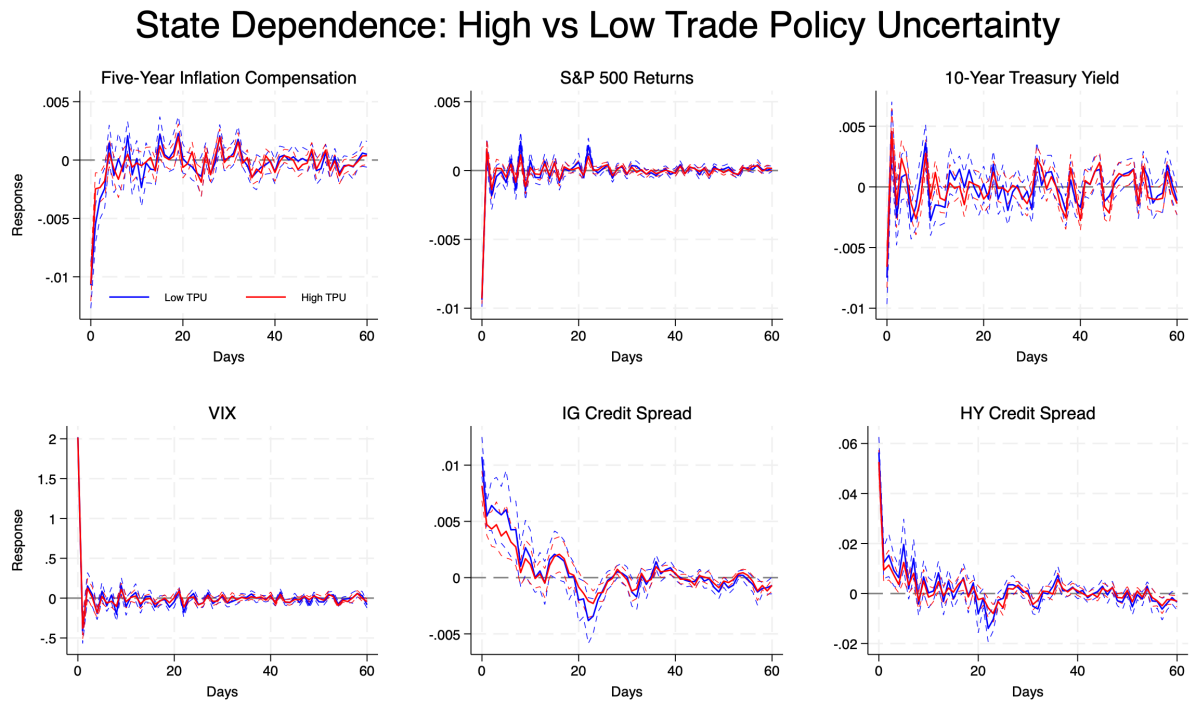


Figure 5: State dependence of trade-war shock responses under low and high trade policy uncertainty

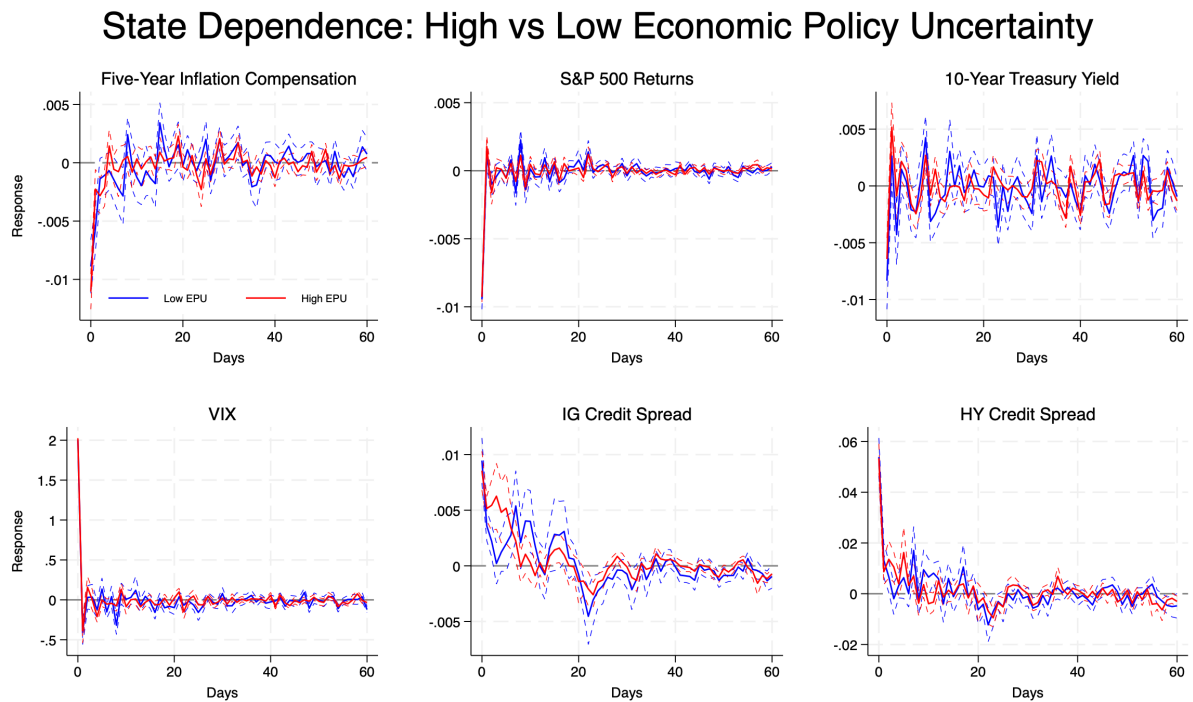


Figure 6: State dependence of trade-war shock responses under low and high economic policy uncertainty

the identified trade-war shock. If anything, TPU provides slightly more visible differences than EPU in credit spreads, but the evidence is better interpreted as robustness across uncertainty states rather than as strong uncertainty-state amplification. Using lagged uncertainty states also helps address the concern that contemporaneous TPU or EPU may be mechanically affected by trade-war announcements themselves. The results therefore indicate that the main macro-financial transmission pattern documented in the baseline analysis is not driven solely by the contemporaneous behavior of standard uncertainty measures.

5.2 Frequency-domain decomposition: low-frequency versus high-frequency trade-war shocks

Figure 7 reports the impulse responses to the low-frequency and high-frequency components of the identified trade-war shock, using a 10-trading-day cutoff. The decomposition shows that the raw daily shock is dominated by its high-frequency component, which is consistent with the event-driven nature of trade-war announcements. As reported in Table 2, the high-frequency component has a substantially larger standard deviation than the low-frequency component and is much more strongly correlated with the baseline identified shock.

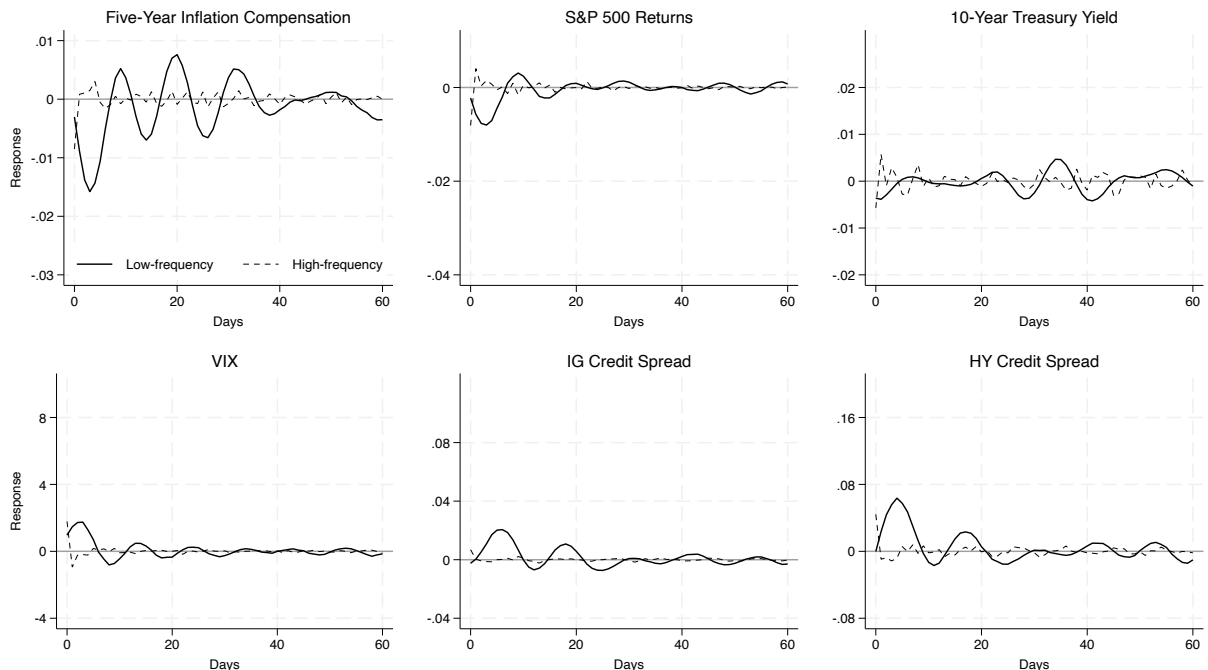


Figure 7: Impulse responses to the low-frequency and high-frequency components of the identified trade-war shock in the baseline five-year specification

Table 2: Summary statistics for the Fourier decomposition of the identified trade-war shock

	Low-frequency	High-frequency
Standard deviation	0.461	0.887
Correlation with identified shock	0.462	0.887
Correlation between components	0.001	

At the same time, the low-frequency component is more important for the persistence and economic significance of the responses in the baseline five-year specification. High-frequency shocks account mainly for the immediate repricing around the event date, whereas the low-frequency component contributes more to the subsequent adjustment in five-year inflation compensation, volatility, and credit spreads. In particular, the responses of the VIX and both investment-grade and high-yield credit spreads are more clearly driven by the low-frequency component over short-to-medium horizons, while the response of five-year inflation compensation also appears more persistent for the low-frequency component. This pattern suggests that trade-war news is transmitted through both short-run announcement effects and a slower-moving repricing of macro-financial conditions.

The frequency-domain results, therefore, complement the baseline evidence. They show that trade-war shocks are not purely one-day financial disturbances: while they arrive as high-frequency news innovations, part of their effect propagates through a lower-frequency component associated with more persistent revisions to growth expectations and risk premia. Equity returns also exhibit this two-part structure, with a stronger immediate response to the high-frequency component but a more sustained adjustment linked to the low-frequency component. Overall, the evidence suggests that high-frequency variation dominates the raw identified shock, whereas low-frequency variation plays a more important role in shaping the persistence of macro-financial responses.

5.3 Political Regime Heterogeneity

This regime heterogeneity exercise uses the baseline 76-event trade-war calendar employed in the main identification exercise. It therefore asks whether the transmission of the same baseline identified trade-war shock differs across the Trump I, Biden, and Trump II periods. Because event days are not evenly distributed across political regimes, the regime-interaction results should be interpreted as suggestive evidence of heterogeneous transmission rather than as a separate identification design with balanced regime samples.

This regime heterogeneity exercise uses the baseline 76-event trade-war calendar employed in the main identification exercise. The event counts are uneven across political regimes: 65 baseline event days fall in the Trump I period, 6 fall in the Biden period, and 5 fall in the Trump II period. The regime-interaction estimates therefore compare responses across politically distinct periods using the same identified baseline

trade-war shock, but the Biden and Trump II estimates are based on relatively small event counts. For this reason, the regime results should be interpreted as suggestive evidence of heterogeneous transmission rather than as a separate identification design with balanced regime samples. Figure 8 and Table 3 examine whether the transmission of the identified trade-war shock differs across political regimes. We compare the Trump I, Biden, and Trump II periods using regime-interacted local projections. The evidence points to some regime-dependent transmission, but the heterogeneity is uneven across variables and cumulative response horizons.

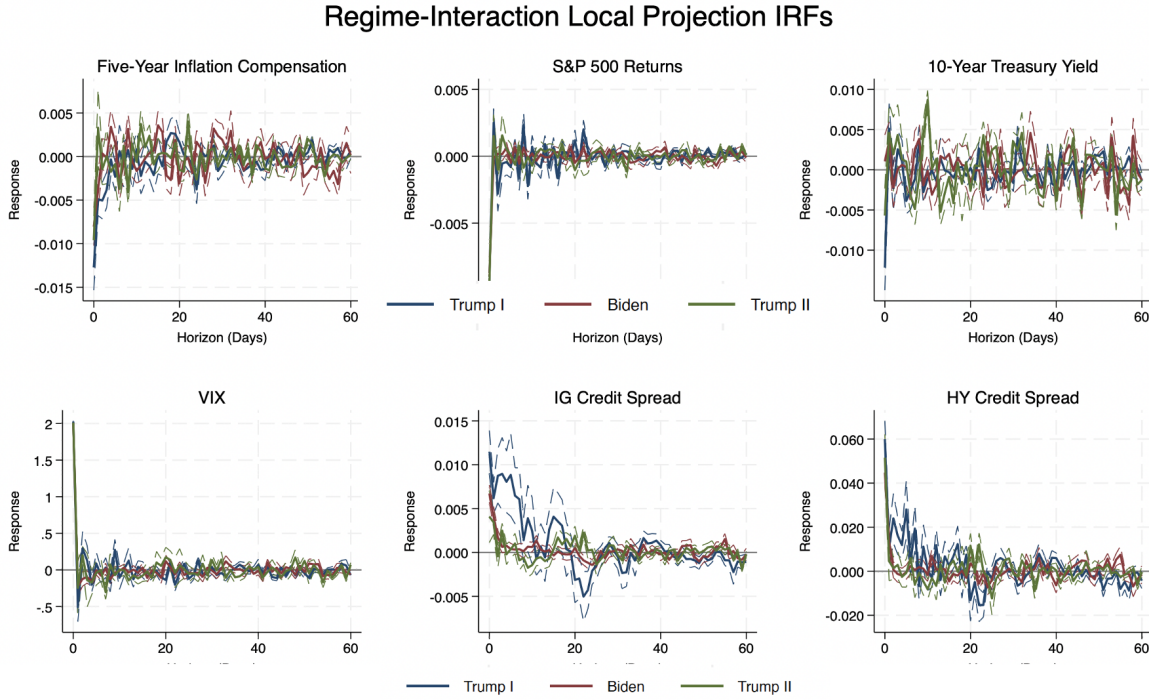


Figure 8: Regime-Interaction Local Projection IRFs

Table 3: Wald Tests of Regime Differences Across Cumulative Response Windows

Variable	Cumulative (0–5)		Cumulative (0–20)		Cumulative (0–60)	
	T1 = Biden	T1 = T2	T1 = Biden	T1 = T2	T1 = Biden	T1 = T2
Five-Year Inflation Compensation	0.089*	0.628	0.010***	0.073*	0.824	0.663
S&P 500 Returns	0.741	0.454	0.050*	0.027**	0.384	0.505
10-Year Treasury Yield	0.108	0.182	0.568	0.375	0.967	0.950
VIX	0.571	0.587	0.014**	0.039**	0.887	0.025**
IG Credit Spread	0.078*	0.091*	0.176	0.080*	0.442	0.385
HY Credit Spread	0.037**	0.052*	0.099*	0.028**	0.598	0.373

Notes: The table reports p-values from Wald tests of equality of regime-specific local-projection coefficients across cumulative response windows. The 0–5, 0–20, and 0–60 windows are defined as $y_{t+h} - y_{t-1}$ for $h = 5, 20$, and 60 , respectively. “T1 = Biden” tests the null hypothesis that the Trump I and Biden coefficients are equal. “T1 = T2” tests the null hypothesis that the Trump I and Trump II coefficients are equal. ***, **, and * denote significance at the 1%, 5%, and 10% levels, respectively.

The clearest evidence of regime heterogeneity appears over the cumulative 0–20 window. For five-year inflation compensation, the null hypothesis of equal responses between Trump I and Biden is rejected at the 1 percent level, while the Trump I–Trump II comparison is marginally significant at the 10 percent level. This suggests that the medium-horizon cumulative response of inflation compensation differs across political regimes, especially between Trump I and Biden. However, the differences are not persistent over the full 0–60 window, where neither comparison is statistically significant.

For equity returns, there is little evidence of regime differences over the 0–5 and 0–60 windows. The main exception is the cumulative 0–20 window, where both the Trump I–Biden and Trump I–Trump II comparisons are statistically significant or marginally significant. This points to some medium-window heterogeneity in equity-market transmission, but not to a stable difference across the full response horizon. For the ten-year Treasury yield, none of the Wald tests are statistically significant across the three cumulative windows, indicating that the yield response is comparatively invariant across political regimes.

The evidence of regime heterogeneity is stronger for volatility and credit spreads. For the VIX, the cumulative 0–20 window shows significant differences between Trump I and both Biden and Trump II, while the 0–60 window shows a significant Trump I–Trump II difference. This suggests that regime-dependent transmission is more visible in market volatility than in long-term yields. Investment-grade credit spreads show weaker evidence, with only marginal significance in the 0–5 and 0–20 windows. High-yield credit spreads display more systematic differences at shorter and medium cumulative horizons: the Trump I–Biden comparison is significant in the 0–5 window and marginally significant in the 0–20 window, while the Trump I–Trump II comparison is significant in the 0–20 window and marginally significant in the 0–5 window.

The results indicate that political regimes matter for some dimensions of the macro-financial transmission of trade-war shocks, but the differences are concentrated in specific variables and cumulative horizons. The evidence is strongest over the 0–20 window and is most visible in five-year inflation compensation, the VIX, and high-yield credit spreads. By contrast, the ten-year Treasury yield response is largely stable across regimes. Given the smaller number of event days in the Biden and Trump II periods, these findings should be interpreted as suggestive rather than definitive evidence of regime-dependent transmission.

5.4 Asymmetric effects of negative and positive trade-war news.

Unlike the regime heterogeneity exercise, which uses the baseline 76-event calendar, the asymmetry exercise uses the broader Trade War Event Database (TWDB). The complete TWDB event table is displayed in appendix. This choice is deliberate. The purpose of the asymmetry exercise is to compare adverse and favorable trade-war news, and the TWDB provides a more systematic event-level classification by sign, event type, event family, and policy regime. It also contains a larger number of de-escalation events than

the baseline calendar, which makes it better suited for a negative-positive split. In the TWDB sample, 63 event days are classified as negative, 39 are classified as positive, and one event day is classified as mixed. The mixed event is excluded from the two-way negative-positive comparison.

To preserve a common identification framework across event types, we do not identify separate shocks from the negative and positive subsamples. Instead, we first identify a single trade-war shock from the full TWDB event set and then split the resulting shock series according to the sign of the event classification. This avoids confounding directional asymmetry with differences in covariance structure across small event subsamples. The asymmetry exercise should therefore be interpreted as an extension of the baseline identification rather than as a replacement for the baseline 76-event specification. Figure 9 compares the local-projection responses to negative and positive trade-war news, where the underlying trade-war shock is identified from the broader TWDB event set and then split according to the sign of the event classification. The results suggest some evidence of asymmetry, although the differences are not sharp across all variables and horizons. Negative trade-war news tends to be associated with somewhat more adverse responses in five-year inflation compensation, long-term Treasury yields, and corporate credit spreads, while positive news is associated with partial reversals. At the same time, the responses to positive news are generally more volatile, and in several cases, the two paths remain relatively close to each other. Equity returns and the VIX show only limited separation between negative and positive news, whereas somewhat clearer differences appear for Treasury yields and for IG and HY credit spreads. Overall, the evidence suggests that negative trade-war news is somewhat more disruptive than positive trade-war news is supportive, but the asymmetry is not sufficiently strong to overturn the broader conclusion that trade-war shocks mainly operate through weaker growth expectations and tighter financial conditions.

This pattern is broadly consistent with the recent evidence in Gaies (2025), who shows that tariff hikes and tariff cuts have asymmetric effects on U.S. stock returns, with tariff hikes generating stronger and more persistent effects than tariff cuts.⁶ A similar interpretation applies here. Adverse trade-war news appears to have a larger and somewhat more persistent financial impact than favorable trade-war news has in the opposite direction. At the same time, the relatively noisy response of positive news suggests that de-escalation events are not always interpreted by financial markets as fully credible or durable reversals of trade policy risk. This helps explain why the positive-news responses do not generate a clean mirror-image recovery relative to the negative-news responses.

⁶Gaies (2025) studies the asymmetric effects of tariff hikes and tariff cuts on U.S. stock prices using a nonlinear ARDL framework and time-varying causality tests.

Asymmetry: Negative vs Positive Trade-War News (TWDB shock)

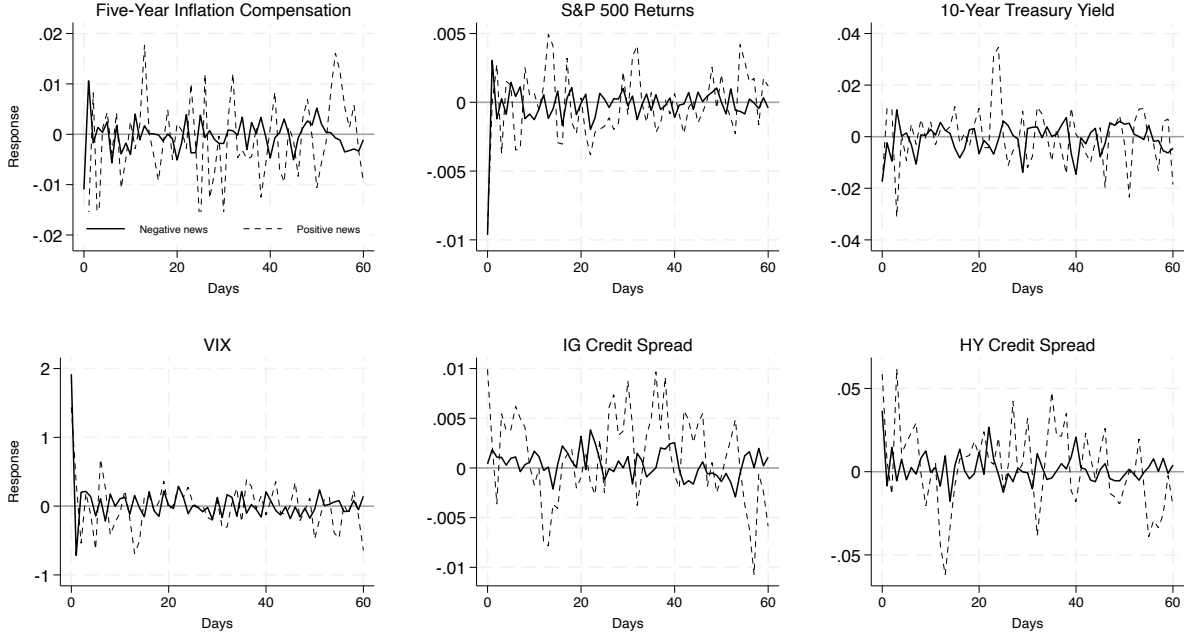


Figure 9: Baseline and asymmetric responses to trade-war shocks.

5.5 Decomposing the Inflation-Compensation Response

Having shown that the identified trade-war shock lowers five-year inflation compensation on impact, we now examine which component of inflation compensation drives this response. Figure 10 reports the responses of daily changes in the decomposed five-year inflation-compensation components to the same GMM-identified trade-war shock used in the baseline analysis. The decomposition separates inflation compensation into expected inflation, a TIPS liquidity premium, and a residual premium component.

Several features stand out. First, the decline in five-year inflation compensation is concentrated on impact. A one-standard-deviation trade-war shock lowers the daily change in five-year inflation compensation by about 1 basis point on the event day. The response then moves back toward zero over subsequent trading days, consistent with an immediate repricing of medium-horizon inflation compensation rather than a persistent upward inflationary response.

In addition, the decomposition suggests that the impact decline is not primarily a TIPS liquidity phenomenon. The estimated response of the TIPS liquidity premium is economically very small throughout the response horizon. Although the sign of the liquidity component is broadly consistent with the idea that weaker TIPS liquidity can lower observed inflation compensation, its magnitude is negligible relative to the responses of the other components.

Responses of Daily Changes in Five-Year Inflation-Compensation Components to the Identified Trade-War Shock

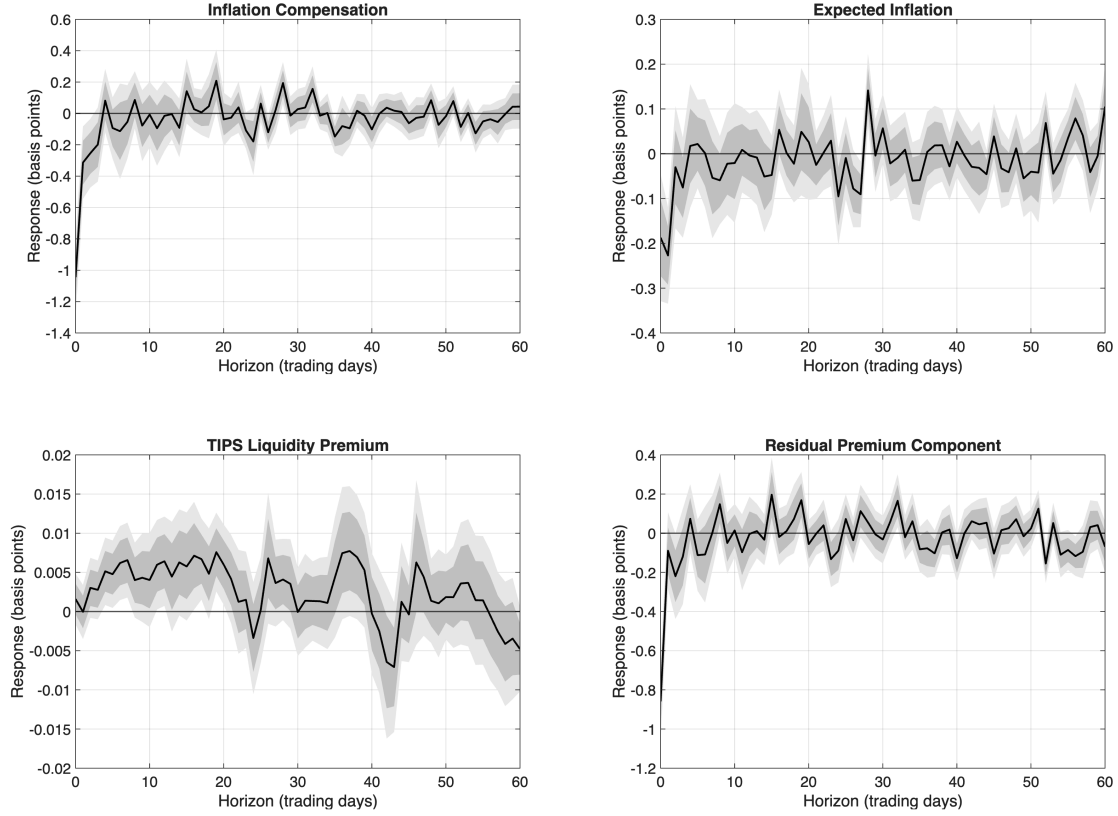


Figure 10: Responses of Daily Changes in Five-Year Inflation-Compensation Components

Notes: The figure reports local-projection responses of daily changes in decomposed five-year inflation-compensation components to the GMM-identified trade-war shock. The shaded areas denote 68% and 90% Newey–West confidence bands. Responses are expressed in basis points. The residual premium component is computed as inflation compensation plus the estimated TIPS liquidity premium minus expected inflation.

Then, the decline in inflation compensation is mainly associated with the expected-inflation and residual-premium components. On impact, expected inflation falls modestly, while the residual premium component accounts for the larger part of the decline. This pattern suggests that financial markets interpret trade-war shocks as weakening the medium-horizon inflation outlook and repricing non-expectational premia, rather than as generating sustained cost-push inflation pressure. The response of the residual premium component is also consistent with the broader macro-financial evidence: trade-war shocks lower equity returns, reduce long-term Treasury yields, raise volatility, and widen credit spreads, all of which point to weaker growth expectations and tighter financial conditions.

The decomposition reinforces the interpretation of the baseline results. The fall in five-year inflation compensation reflects a combination of lower expected inflation and a lower residual premium component, while the estimated TIPS liquidity premium plays only a minor role. Since this exercise applies the same GMM-identified shock as the baseline analysis, it should be interpreted as a mechanism exercise rather than as a separate identification strategy. The corresponding two-year and ten-year decomposition results

are reported in Appendix Figures 11 and 12. These appendix results show a broadly similar pattern, with stronger short-horizon repricing at the two-year maturity and more muted responses at the ten-year maturity.

6 Robustness and External Validation

6.1 Identification sensitivity and falsification tests

This section examines whether the baseline event-day heteroskedasticity design is sensitive to the construction of the trade-war event set, the treatment of overlapping macroeconomic announcements, the definition of the non-event regime, and the inclusion of daily macro-financial controls. The exercise is motivated by the high-frequency and event-based identification literature, which emphasizes the importance of narrow event definitions, placebo timing checks, and transparent event construction when extracting policy shocks from financial-market data (Bianchi et al., 2023; Poilly and Tripier, 2025). Because our design is implemented at the daily frequency, the following exercises should be interpreted as robustness and falsification checks rather than as substitutes for an intraday announcement-window design.

A remaining concern is that trade-war announcements may be partially anticipated or leaked before the recorded event day. This issue is difficult to eliminate completely in daily data. If financial markets react before the event date, the estimated event-day response may understate the full effect of the announcement. Conversely, if announcements occur after regular trading hours or on non-trading days, a mechanical calendar-date assignment may misalign the market reaction. We address these concerns in several ways. First, the baseline event calendar is aligned to the relevant trading day, so announcements outside regular market hours or on non-trading days are assigned to the next trading day on which financial markets can respond. Second, the event-buffer specification removes adjacent non-event observations around baseline event days from the non-event regime, reducing the risk that pre-event or post-event adjustment contaminates the covariance comparison. Third, the lead placebo tests examine whether the covariance shift appears before the event date. Fourth, the clean-event specification removes event days that overlap with major macroeconomic announcements. These checks reduce, but do not fully eliminate, concerns about anticipation and information leakage. The estimated responses should therefore be interpreted as market reactions around identifiable trade-war announcement dates rather than as the total effect of all information about trade policy.

We first re-estimate the baseline specification under several alternative event definitions. The clean-event specification excludes ambiguous trade-war events and event days that overlap with major macroeconomic news. The no-macro-news specification excludes all FOMC announcement days, CPI release days, and employment-report release days from the estimation sample. The event-buffer specification removes the

trading days immediately before and after trade-war events from the non-event regime, reducing the possibility that adjacent trading days contaminate the comparison group. We also estimate a VAR-control specification in which the reduced-form VAR is augmented with daily monetary and macro-financial controls, including FOMC, CPI, and employment-report indicators, changes in the effective federal funds rate, and oil-price changes. Finally, we repeat the analysis using the extended trade-war event database. The detailed dates affected by identification-sensitivity specifications could be seen in the appendix.

The five-year inflation-compensation response remains negative across all identification-sensitivity specifications. In the baseline specification, a one-standard-deviation adverse trade-war shock lowers five-year inflation compensation by about 1.08 basis points on impact. The corresponding impact responses are about -1.06 basis points for the clean-event specification, -1.16 basis points when major macro-news days are excluded, -1.09 basis points in the event-buffer specification, -0.91 basis points when monetary and macro-financial controls are included in the VAR, and -1.08 basis points when the extended event database is used. Thus, across the main identification-sensitivity exercises, the impact response remains negative and ranges from approximately -0.91 to -1.16 basis points. These estimates indicate that the baseline result is not driven by a particular event definition, by the inclusion of major macro-news days, by the definition of the non-event regime, or by the omission of daily monetary and macro-financial controls.

The covariance-shift diagnostics also support the event-day heteroskedasticity design. In the baseline specification, the covariance-difference matrix is dominated by a single positive component, and the variance of the projected trade-war shock is more than twice as large on event days as on non-event days. The baseline event-to-non-event variance ratio is 2.15. This pattern is consistent with the identifying assumption that trade-war event days reveal a stronger common shock component in financial-market residuals than non-event days. The same diagnostic logic is used across the alternative event sets and control specifications.

We also conduct placebo and timing checks. In the matched random placebo exercise, placebo dates are drawn from non-trade and non-macro-news days in the same calendar environment as the true event dates. The average placebo variance ratio is 1.07, compared with 2.15 for the actual trade-war event dates. The right-tail placebo p -value is 0.15, so we interpret the random-placebo evidence as supportive rather than definitive. To examine anticipation and information leakage, we also shift the event dates five and ten trading days before the actual announcements. The corresponding lead-placebo variance ratios are 0.92 and 0.59, respectively, compared with matched placebo benchmarks of 0.96 and 0.97. These lead-placebo results suggest that the covariance shift is not systematically present before trade-war announcements, although they do not rule out all forms of anticipation or gradual information diffusion.

Overall, the identification-sensitivity, placebo, and lead-timing exercises show that the baseline response of inflation compensation is robust to alternative event definitions, macro-news exclusions, non-event window

choices, and VAR controls. The falsification tests further suggest that the covariance shift is concentrated around actual trade-war announcement dates rather than being a generic feature of nearby non-event or pre-event trading days.

Table 4: Identification sensitivity: five-year inflation compensation

Specification	$h = 0$	$h = 5$	$h = 20$	$h = 60$
Baseline	-1.08	-0.08	-0.04	0.03
Clean events	-1.06	-0.09	-0.04	0.03
No macro-news days	-1.16	0.06	-0.03	0.04
Event-buffer specification	-1.09	-0.07	-0.04	0.03
VAR with macro and monetary controls	-0.91	-0.08	-0.04	0.03
Extended TWDB	-1.08	-0.08	-0.04	0.03

Notes: The table reports local-projection responses of five-year inflation compensation to a one-standard-deviation adverse trade-war shock. Responses are reported in basis points. The clean-event specification excludes ambiguous trade-war events and trade-war events overlapping with major macroeconomic news. The no-macro-news specification excludes FOMC announcement days, CPI release days, and employment-report release days. The event-buffer specification removes the trading days immediately before and after trade-war events from the non-event regime. The VAR-control specification augments the reduced-form VAR with FOMC, CPI, and employment-report indicators, changes in the effective federal funds rate, and oil-price changes.

Table 5: Placebo and timing diagnostics

Diagnostic	Test statistic	Placebo benchmark	Placebo p -value
Actual events vs. matched random placebo	2.15	1.07	0.15
Lead placebo: $t - 5$ trading days	0.92	0.96	0.25
Lead placebo: $t - 10$ trading days	0.59	0.97	0.56

Notes: The test statistic is the variance ratio of the projected trade-war shock on actual or placebo event days relative to non-event days. For the actual trade-war events, the placebo benchmark is the mean variance ratio from matched random placebo draws. For the lead-placebo tests, the placebo benchmark is the corresponding matched placebo mean for the shifted event dates. The p -values are right-tail placebo probabilities. A higher variance ratio indicates a stronger event-day covariance shift.

6.2 Media-Salience-Weighted Event Identification

We next examine whether the baseline event-day heteroskedasticity result is affected by treating all trade-war announcements as equally informative. Some announcements receive substantially more media attention than others, and more salient events may contain stronger information shocks. We therefore construct event-level media-salience weights using news coverage from the Global Database of Events, Language, and Tone (GDELT).

For each trade-war event, we use event-specific GDELT query blocks corresponding to its policy category. The query blocks cover U.S.–China tariffs, Phase One negotiations, Section 301 actions, Section 232 steel and aluminum tariffs, European retaliation, Mexico-related tariffs, Huawei export controls, and IEEPA or fentanyl-related tariff actions. The complete query design and keywords are reported in Appendix Table 20.

The preferred hybrid weight uses GDELT’s normalized TimelineVol measure when available, raw article counts when TimelineVol is unavailable, and a neutral weight of one when neither measure is available. The

coverage measures are log transformed, winsorized at the 95th percentile, and normalized to have a mean of one across event days. The weights are used to recompute the event-day residual covariance matrix, while the non-event covariance matrix remains unchanged. We then apply the same rank-one minimum-distance GMM estimator and sign normalization as in the baseline identification. The media weights affect the identification stage only and are not used as observation weights in the local-projection regressions. Full details of the weight construction and weighted covariance estimator are provided in Appendix Section A.7.

In the final hybrid-weight sample, 62 event days use TimelineVol coverage, 10 use raw article counts, and 4 receive a neutral weight. The effective number of event days is 71.8 out of 76, indicating that the weighting procedure does not concentrate identification on a small number of highly publicized announcements.

Table 6 compares the identification diagnostics from the unweighted baseline and the GDELT media-weighted specification. The event-to-non-event variance ratio remains approximately 2.15 in both specifications. The largest eigenvalue of the covariance-difference matrix increases from 4.18 to 5.52 under media weighting, indicating that the leading event-day covariance shift becomes somewhat more pronounced. The estimated impact response of five-year inflation compensation remains -1.08 basis points.

Table 6: Media-Weighted Identification Diagnostics

Specification	Event days	Effective event days	Largest eigenvalue	Variance ratio	Impact response
Unweighted baseline	76	76.00	4.18	2.15	-1.08
GDELT hybrid media-weighted	76	71.80	5.52	2.15	-1.08

Notes: The table reports identification diagnostics for the unweighted and GDELT media-weighted event-day covariance specifications. The GDELT hybrid weight uses TimelineVol coverage when available, raw article counts when TimelineVol is unavailable, and a neutral weight of one for the remaining missing observations. The weights are winsorized and normalized to have a mean of one across event days. The impact response is the response of five-year inflation compensation to a one-standard-deviation adverse trade-war shock and is reported in basis points.

Table 7 reports the responses of five-year inflation compensation at selected horizons. The media-weighted estimates are identical to the baseline estimates after rounding. Inflation compensation falls by 1.08 basis points on impact, while the response becomes small at longer horizons and is close to zero by 60 trading days.

Table 7: Inflation-Compensation Responses under Media-Weighted Identification

Specification	$h = 0$	$h = 5$	$h = 20$	$h = 60$
Unweighted baseline	-1.08	-0.08	-0.04	0.03
GDELT hybrid media-weighted	-1.08	-0.08	-0.04	0.03

Notes: The table reports local-projection responses of five-year inflation compensation to a one-standard-deviation adverse trade-war shock. Responses are reported in basis points. The GDELT hybrid media-weighted specification recomputes the event-day residual covariance matrix using event-level media-salience weights while leaving the non-event covariance matrix unchanged.

The similarity of the two sets of estimates shows that the baseline result is not driven by assigning equal importance to events with substantially different levels of media coverage. Media weighting strengthens the concentration of the event-day covariance shift in its leading component but leaves the magnitude and

persistence of the inflation-compensation response essentially unchanged.

6.3 VAR lag, maturity, and sample robustness

We next examine whether the baseline response depends on the lag length used to extract reduced-form VAR residuals, the inflation-compensation maturity, or the sample period. These checks address three related concerns. First, the identified shock could be sensitive to the short-run dynamics imposed in the residualization step. Second, the baseline five-year maturity may conceal different responses at shorter or longer horizons of the inflation-compensation curve. Third, the result could be driven mainly by the intensive 2018–2019 phase of the U.S.-China trade war rather than by the broader 2016–2026 sample.

To assess lag length sensitivity, we estimate the reduced-form VAR using alternative lag lengths $p \in \{1, 2, 3, 5\}$:

$$X_t = c_p + \sum_{\ell=1}^p A_{\ell,p} X_{t-\ell} + u_t^{(p)}, \quad (17)$$

where X_t contains daily changes in inflation compensation and the accompanying financial-market variables used in the baseline system. For each lag length, we use the resulting residuals $u_t^{(p)}$ to re-estimate the event-day covariance shift and recover a corresponding trade-war shock $s_t^{(p)}$. We then estimate the same local-projection specification:

$$y_{t+h} - y_{t-1} = \alpha_h^{(p)} + \beta_h^{(p)} s_t^{(p)} + \gamma_h^{(p)} s_{t-1}^{(p)} + \delta_h^{(p)} y_{t-1} + \varepsilon_{t+h,h}^{(p)}. \quad (18)$$

Table 8 shows that the impact response of five-year inflation compensation is essentially unchanged across lag choices. Although the unrounded estimates differ slightly, they all round to -1.08 basis points on impact. The event-day variance ratio remains above two in all specifications, ranging from 2.15 to 2.43. This stability indicates that the baseline result is not driven by a particular short-run dynamic specification of the reduced-form VAR.

We next examine whether the response differs across maturities. For maturity $m \in \{2, 5, 10\}$, we estimate

$$IC_{t+h}^{(m)} - IC_{t-1}^{(m)} = \alpha_h^{(m)} + \beta_h^{(m)} s_t^{(m)} + \gamma_h^{(m)} s_{t-1}^{(m)} + \delta_h^{(m)} IC_{t-1}^{(m)} + \varepsilon_{t+h,h}^{(m)}, \quad (19)$$

where $IC_t^{(m)}$ denotes inflation compensation at maturity m . Table 9 reports the baseline responses for two-year, five-year, and ten-year inflation compensation. The impact response is negative at all maturities, but the decline is larger at shorter horizons. A one-standard-deviation adverse trade-war shock lowers two-year inflation compensation by 1.25 basis points, five-year inflation compensation by 1.08 basis points, and ten-year inflation compensation by 0.90 basis points on impact. This downward-sloping maturity pattern

suggests that trade-war shocks are priced most strongly into near- and medium-term inflation compensation, while longer-horizon compensation responds less on impact. The negative response across maturities also shows that the baseline result is not specific to the five-year series.

Finally, we estimate the response over alternative sample periods. Let $\mathcal{S}$ denote a given subsample. For each subsample, the response is obtained from

$$IC_{t+h}^{(5)} - IC_{t-1}^{(5)} = \alpha_h^{(\mathcal{S})} + \beta_h^{(\mathcal{S})} s_t^{(\mathcal{S})} + \gamma_h^{(\mathcal{S})} s_{t-1}^{(\mathcal{S})} + \delta_h^{(\mathcal{S})} IC_{t-1}^{(5)} + \varepsilon_{t+h,h}^{(\mathcal{S})}, \quad t \in \mathcal{S}. \quad (20)$$

Table 10 reports split-sample estimates for the five-year inflation-compensation response. The pre-COVID and pre-2020 samples produce similar impact responses of -0.99 and -0.97 basis points, respectively. The post-COVID and post-2020 estimates are also negative, at -1.12 and -1.13 basis points, respectively. However, these later-sample estimates should be interpreted cautiously because they are based on a much smaller number of trade-war event days. Overall, the lag-length, maturity, and sample-split exercises indicate that the baseline impact response is not an artifact of the VAR lag choice, the selected inflation-compensation maturity, or the 2018–2019 trade-war period alone.

Table 8: VAR lag robustness: five-year inflation compensation

VAR lag	Event days	Variance ratio	$h = 0$	$h = 5$	$h = 20$	$h = 60$
1	76	2.15	-1.08	-0.08	-0.04	0.03
2	76	2.23	-1.08	-0.09	-0.04	0.01
3	76	2.43	-1.08	-0.08	-0.05	0.01
5	76	2.31	-1.08	-0.07	-0.07	0.01

Notes: The table reports local-projection responses of five-year inflation compensation to a one-standard-deviation adverse trade-war shock under alternative lag lengths in the reduced-form VAR. Responses are reported in basis points. The variance ratio is the variance of the projected trade-war shock on event days divided by its variance on non-event days.

Table 9: Maturity robustness

Maturity	Event days	Variance ratio	$h = 0$	$h = 5$	$h = 20$	$h = 60$
Two-year IC	76	2.16	-1.25	-0.12	-0.08	0.00
Five-year IC	76	2.15	-1.08	-0.08	-0.04	0.03
Ten-year IC	76	2.16	-0.90	-0.22	0.02	0.02

Notes: The table reports local-projection responses of two-year, five-year, and ten-year inflation compensation to a one-standard-deviation adverse trade-war shock. Responses are reported in basis points. The variance ratio is the variance of the projected trade-war shock on event days divided by its variance on non-event days.

6.4 Oil price benchmark, monetary policy shocks, joint controls, and economic significance

We compare the estimated response to trade war shocks with daily WTI price movements and the high frequency FOMC shocks developed by Jarociński and Karadi (2020). The purpose is to assess the economic

Table 10: Sample split robustness: five-year inflation compensation

Sample	Event days	Variance ratio	$h = 0$	$h = 5$	$h = 20$	$h = 60$
Pre-COVID sample	65	1.67	-0.99	-0.16	0.13	-0.08
Post-COVID sample	11	6.33	-1.12	-0.11	-0.12	0.08
Pre-2020 sample	63	1.87	-0.97	-0.08	0.09	-0.09
Post-2020 sample	13	5.35	-1.13	-0.08	-0.10	0.09

Notes: The table reports split-sample local-projection responses of five-year inflation compensation to a one-standard-deviation adverse trade-war shock. Responses are reported in basis points. The variance ratio is the variance of the projected trade-war shock on event days divided by its variance on non-event days. The post-COVID and post-2020 estimates are based on a smaller number of trade-war event days and should therefore be interpreted cautiously.

magnitude of the trade war shock and to examine whether the estimated response remains robust after controlling for other macroeconomic disturbances.

The WTI series is observed at the same daily frequency as the financial variables.⁷ The Jarociński and Karadi dataset distinguishes monetary policy shocks from central bank information shocks using financial market surprises around FOMC announcements.⁸ WTI price changes are treated as an oil price benchmark rather than as an identified structural oil supply shock.

For each separate shock or benchmark x_t^j , we estimate

$$IC_{t+h}^{5y} - IC_{t-1}^{5y} = \alpha_h^j + \beta_h^j x_t^j + \gamma_h^j x_{t-1}^j + \delta_h^j IC_{t-1}^{5y} + \varepsilon_{t+h,h}^j. \quad (21)$$

We also estimate a joint specification:

$$\begin{aligned} IC_{t+h}^{5y} - IC_{t-1}^{5y} = & \alpha_h + \beta_h^{TW} s_t + \beta_h^{Oil} o_t + \beta_h^{MP} m_t + \beta_h^{CBI} c_t \\ & + \gamma_h^{TW} s_{t-1} + \gamma_h^{Oil} o_{t-1} + \gamma_h^{MP} m_{t-1} + \gamma_h^{CBI} c_{t-1} + \delta_h IC_{t-1}^{5y} + \varepsilon_{t+h,h}, \end{aligned} \quad (22)$$

where s_t , o_t , m_t , and c_t denote the trade war shock, WTI price innovation, monetary policy shock, and central bank information shock, respectively. As an alternative, we replace the monetary policy and central bank information shocks with the policy indicator surprise.

Table 11 reports the results. In the separate full sample specifications, a one standard deviation adverse trade war shock lowers five year inflation compensation by 1.08 basis points on impact, while a one standard deviation WTI price innovation raises it by 1.36 basis points. In the sample matched to the shock series of Jarociński and Karadi (2020), our identified trade war shock lowers inflation compensation by 1.20 basis points, compared with a decline of 1.34 basis points following the Jarociński and Karadi monetary policy tightening shock.

⁷The WTI crude oil price series is Crude Oil Prices: West Texas Intermediate, Cushing, Oklahoma, FRED series DCOILWTICO. The original source is the U.S. Energy Information Administration.

⁸The shocks are merged with the daily dataset on FOMC announcement dates, set equal to zero on other trading days, and standardized using announcement day observations.

In the joint specification, the adverse trade war shock lowers inflation compensation by approximately 0.99 basis points on impact. Although the magnitude is smaller than the 1.20 basis point response in the corresponding separate specification, the estimated response remains economically similar. The alternative specification using the policy indicator surprise produces almost the same trade war response. These results indicate that the baseline effect is not primarily driven by coincident oil price movements or FOMC news.

Table 11: Oil price and monetary policy shock comparisons

Shock or benchmark	$h = 0$	$h = 5$	$h = 20$	$h = 60$
<i>Panel A: Separate oil price benchmark</i>				
GMM identified trade war shock	-1.08	-0.08	-0.04	0.03
WTI price innovation	1.36	-0.06	-0.01	0.02
<i>Panel B: Separate monetary policy shock comparison</i>				
GMM identified trade war shock	-1.20	-0.09	0.00	0.00
JK monetary policy shock	-1.34	0.89	0.34	1.08
JK central bank information shock	0.10	0.11	0.47	-1.13
JK policy indicator surprise	-1.28	0.92	0.55	0.56
<i>Panel C: Joint specification</i>				
GMM identified trade war shock	-0.99	-0.11	0.00	-0.03
WTI price innovation	1.24	-0.07	-0.02	-0.01
JK monetary policy shock	-1.00	0.89	0.45	0.91
JK central bank information shock	-0.02	0.26	0.56	-0.96
<i>Panel D: Joint specification using the policy indicator</i>				
GMM identified trade war shock	-0.99	-0.10	0.00	-0.02
WTI price innovation	1.23	-0.08	-0.02	-0.01
JK policy indicator surprise	-0.90	0.87	0.55	0.56

Notes: Responses are reported in basis points. Panels A and B report separate local projection estimates. Panels C and D report joint estimates in which the listed shocks and their first lags enter simultaneously. The WTI innovation is an oil price benchmark rather than an identified structural oil shock. The JK monetary policy shock, central bank information shock, and policy indicator surprise are based on the high frequency FOMC shock measures developed by Jarociński and Karadi (2020).

We also assess the economic importance of the trade war shock relative to the decline in five year inflation compensation during the main 2018 escalation period. For event group G , the typical impact effect is

$$\Delta IC_0^G = \hat{\beta}_0 \bar{s}_G, \quad (23)$$

where $\bar{s}_G$ is the average identified trade war shock for event group G . Its share of the 2018 decline is

$$\text{Share}^G = \frac{\Delta IC_0^G}{|\Delta IC^{2018}|} \times 100. \quad (24)$$

A one standard deviation adverse trade war shock is equivalent to approximately 3.5 percent of the 30.7 basis point decline in five year inflation compensation during the 2018 escalation period. A typical escalation event accounts for approximately 0.5 percent of the same decline. Individual announcements therefore have modest effects in isolation, but repeated adverse events may contribute to a broader decline in inflation compensation.

Table 12: Economic significance of trade war shocks

Measure	Effect on five year IC	Share of 2018 decline
One standard deviation adverse shock	-1.08 bps	3.51%
Typical trade war event	-0.05 bps	0.17%
Typical escalation event	-0.16 bps	0.54%
Typical de escalation event	0.34 bps	-1.10%

Notes: The table reports the economic magnitude of the five year inflation compensation response. The one standard deviation adverse shock effect is the baseline impact response. The typical event effects multiply the baseline impact response by the average identified shock on all event days, escalation event days, and de escalation event days. The 2018 decline is the difference between average five year inflation compensation in January and February 2018 and the average from November 2018 to January 2019. A positive effect for de escalation events means that these events raise inflation compensation, so their signed share is negative relative to the decline.

6.5 External inflation expectation benchmarks

We next compare the response of our daily market based decomposition with two external measures of inflation expectations. This exercise is useful because the expected inflation component in the daily decomposition is model implied rather than directly observed. If the interpretation is correct, then independent measures of expected inflation should also tend to decline after the identified trade war shock.

We use two external benchmarks. The first comes from the Federal Reserve Bank of Cleveland inflation expectations release, which provides model based expected inflation measures at different horizons and a long horizon inflation risk premium measure. The second comes from the Survey of Professional Forecasters, which provides survey based CPI and PCE inflation expectations. These series are not available at the daily frequency. The Cleveland Fed series is monthly, while the SPF series is quarterly. For this reason, we do not treat them as direct substitutes for the daily decomposition. Instead, we aggregate the daily GMM identified trade war shock to the frequency of each external benchmark and then estimate lower frequency local projections.

Let s_t denote the daily GMM identified trade war shock. For each month or quarter τ , we construct the lower frequency shock as the sum of daily shocks within that period:

$$S_\tau = \sum_{t \in \tau} s_t. \quad (25)$$

We then estimate local projections of the external benchmark series on the aggregated shock:

$$B_{\tau+h} - B_{\tau-1} = \alpha_h + \beta_h S_\tau + \Gamma'_h Z_{\tau-1} + \varepsilon_{\tau+h,h}, \quad (26)$$

where B_τ denotes the external inflation expectation or risk premium benchmark, and $Z_{\tau-1}$ contains lagged benchmark information. The coefficient β_h measures the response of the external benchmark to the aggregated trade war shock.

Table 13 reports the Cleveland Fed monthly responses. The expected inflation measures generally decline after the identified trade war shock. The response is largest at the shorter inflation horizon. The two year Cleveland Fed expected inflation measure falls by 2.87 basis points on impact. The five year and ten year measures fall by 1.16 and 0.59 basis points, respectively. The ten year inflation risk premium response is smaller and less stable across horizons. Its impact response is negative, but the subsequent responses change sign. This suggests that the external evidence is clearer for expected inflation than for the inflation risk premium.

Table 14 reports the SPF quarterly responses. The survey based five year CPI and PCE inflation forecasts both decline on impact, by 1.34 and 0.94 basis points, respectively. The ten year SPF responses are smaller, especially for CPI inflation, but they are still negative on impact. This pattern is consistent with the maturity results in the market based data, where shorter and medium horizon inflation compensation reacts more strongly than longer horizon compensation.

Overall, these external benchmark exercises support the view that adverse trade war shocks are associated with weaker inflation expectations. They should not be read as a direct validation of the level of the daily expected inflation component, because the external series differ in frequency, construction, and data source. The Cleveland Fed measures are monthly and model based, while the SPF measures are quarterly and survey based. The main value of this exercise is therefore directional. Independent measures of inflation expectations move in the same direction as the expected inflation channel in the daily decomposition, while the evidence for the inflation risk premium is more mixed.

Table 13: Cleveland Fed inflation expectation benchmark responses

Benchmark	$h = 0$	$h = 1$	$h = 3$	$h = 6$	$h = 12$
Expected inflation, 2 year	-2.87	-3.31	-2.78	-0.88	-13.40
Expected inflation, 5 year	-1.16	-1.53	-0.82	0.11	-9.98
Expected inflation, 10 year	-0.59	-0.91	-0.18	0.37	-8.11
Inflation risk premium, 10 year	-0.23	0.03	0.69	0.43	-4.83

Notes: The table reports monthly local projection responses of Cleveland Fed inflation expectation and inflation risk premium benchmarks to the aggregated GMM identified trade war shock. Responses are reported in basis points. Horizons are measured in months.

Table 14: SPF survey based inflation expectation responses

Benchmark	$h = 0$	$h = 1$	$h = 2$	$h = 4$
SPF CPI inflation forecast, 5 year	-1.34	-0.69	-0.19	-3.89
SPF PCE inflation forecast, 5 year	-0.94	-1.38	-0.56	-3.52
SPF CPI inflation forecast, 10 year	-0.11	-0.05	-0.62	-1.90
SPF PCE inflation forecast, 10 year	-0.61	-0.72	-0.50	-1.90

Notes: The table reports quarterly local projection responses of SPF survey based inflation expectation benchmarks to the aggregated GMM identified trade war shock. Responses are reported in basis points. Horizons are measured in quarters.

6.6 Maturity robustness of the decomposition

Following the TIPS inflation-compensation decomposition literature, we treat breakeven inflation as a combination of expected inflation, premium variation, and TIPS liquidity effects (Gürkaynak et al., 2010; Abrahams et al., 2016; D’Amico et al., 2018; Kim et al., 2019). We therefore examine whether the component-level interpretation of the trade-war response is robust across the two-year, five-year, and ten-year maturities. This exercise is based on the accounting identity

$$IC_t(m) = EI_t(m) + RPC_t(m) - LP_t(m), \quad (27)$$

where $IC_t(m)$ denotes inflation compensation at maturity m , $EI_t(m)$ denotes the model-implied expected-inflation component, $LP_t(m)$ denotes the estimated TIPS liquidity-premium component, and $RPC_t(m)$ denotes the residual premium component. The minus sign reflects the convention that a higher TIPS liquidity premium raises TIPS yields and therefore lowers observed inflation compensation. Equivalently, the residual premium component is defined as

$$RPC_t(m) = IC_t(m) + LP_t(m) - EI_t(m). \quad (28)$$

Because the residual premium component is defined from the accounting identity, the relevant robustness question is not whether the identity holds mechanically, but whether the component responses deliver a similar interpretation across maturities. This distinction is important because $RPC_t(m)$ should not be interpreted narrowly as a pure inflation risk premium. Instead, it captures the remaining premium component after accounting for model-implied expected inflation and the estimated transaction-volume-based TIPS liquidity component. It may therefore also contain other premium variation, tax effects, embedded option effects, measurement error, and wedges not separately identified by the liquidity proxy.

For each component $z_t(m) \in \{IC_t(m), EI_t(m), LP_t(m), RPC_t(m)\}$, we estimate the same local-projection specification:

$$z_{t+h}(m) - z_{t-1}(m) = \alpha_h^{z,m} + \beta_h^{z,m} s_t + \gamma_h^{z,m} s_{t-1} + \delta_h^{z,m} z_{t-1}(m) + \varepsilon_{t+h,h}^{z,m}, \quad (29)$$

where s_t is the baseline GMM-identified adverse trade-war shock. The coefficient $\beta_h^{z,m}$ measures the response of component z at maturity m to a one-standard-deviation adverse trade-war shock. Responses are reported in basis points.

Table 15 reports the impact responses of the decomposition components across maturities. At the five-year maturity, inflation compensation falls by about 1.04 basis points on impact. This decline is accounted for mainly by the expected-inflation and residual-premium components. Expected inflation falls by about

0.19 basis points, while the residual premium component falls by about 0.86 basis points. By contrast, the estimated TIPS liquidity-premium response is close to zero.

The same broad pattern is present across maturities. At the two-year horizon, the inflation-compensation response is larger and is split between expected inflation and the residual premium component. At the ten-year horizon, the response is smaller, but the estimated liquidity component remains quantitatively negligible. These results support the interpretation that, within this reduced-form decomposition, adverse trade-war shocks lower market-based inflation compensation mainly through expected-inflation and residual-premium channels rather than through the transaction-volume-based TIPS liquidity component. The conclusion should be interpreted with caution because the residual premium component is not a direct measure of inflation risk premia alone. Nevertheless, the negative response of inflation compensation is robust across maturities, and the component-level interpretation is not driven by the five-year specification alone.

Table 15: Impact responses of decomposition components

Maturity	IC	EI	LP	RPC
2-year	-1.25	-0.50	0.0004	-0.75
5-year	-1.04	-0.19	0.0016	-0.86
10-year	-0.85	-0.09	0.0027	-0.75

Notes: The table reports impact responses to a one-standard-deviation adverse trade-war shock. Responses are reported in basis points. IC denotes inflation compensation, EI denotes expected inflation, LP denotes the estimated TIPS liquidity-premium component, and RPC denotes the residual premium component. Since $IC = EI + RPC - LP$, a positive liquidity-premium response contributes negatively to inflation compensation, although its magnitude is quantitatively negligible in these estimates.

7 Discussion and Policy Implications

The central question of the paper is whether financial markets interpret trade-war announcements primarily as inflationary cost-push shocks or as broader macro-financial disturbances associated with weaker growth and higher risk premia. The evidence supports the latter interpretation. Following adverse trade-war news, inflation compensation, equity prices, and long-term Treasury yields fall, while volatility and corporate credit spreads rise. This joint response points to weaker expected activity, greater risk aversion, safe-haven demand, and tighter financial conditions rather than a dominant persistent cost-push inflation channel.

This does not imply that tariffs have no inflationary effects. Tariffs can raise import prices and production costs through direct pass-through channels (Amiti et al., 2019; Fajgelbaum et al., 2020). At daily horizons, however, financial markets appear to place greater weight on the demand, uncertainty, and risk-premium effects of trade-war news. The decomposition results support this interpretation. The decline in five-year inflation compensation is associated mainly with the expected-inflation and residual-premium components, while the transaction-volume-based TIPS liquidity component is small. The residual premium should never-

theless be interpreted cautiously because it may contain inflation risk compensation, tax effects, embedded option effects, measurement error, and other unidentified wedges.

Implications for Monetary Policy

Trade-war shocks create a difficult diagnostic problem for monetary policymakers because different channels can move inflation and activity in opposite directions. Import-price pass-through may raise headline inflation, while weaker demand may reduce expected activity and inflation compensation. At the same time, lower equity prices, higher volatility, and wider credit spreads may tighten financial conditions without an immediate change in the policy rate.

The persistence of the tariff regime is therefore important. Alessandria et al. (2025) show that beliefs about the expected duration of U.S.–China tariffs changed during the trade-war period. A temporary tariff announcement may have limited effects on medium-term inflation, whereas a persistent escalation may have larger effects on investment, production networks, and price setting. Our estimates indicate that weaker-growth and risk-premium channels dominate the immediate market response, but they do not imply that import-price pass-through is absent at lower frequencies.

The findings also matter for central-bank communication. Treating trade-war shocks only as inflationary supply disturbances may understate their effects on expected growth and financial conditions. A balanced assessment should recognize both the direct effect of tariffs on import prices and the indirect effects operating through aggregate demand, uncertainty, risk sentiment, and credit conditions. High-frequency market responses are useful for identifying how investors initially interpret the shock, but they should not provide the sole basis for slower-moving monetary-policy decisions.

Inflation Anchoring and Measurement

The maturity pattern provides cautiously reassuring evidence about long-run inflation anchoring. Ten-year inflation compensation responds less than the two-year and five-year measures, suggesting that trade-war shocks affect short- and medium-horizon inflation pricing more strongly than long-run expectations. This conclusion should not be generalized mechanically, however, because the sample covers a period in which long-run U.S. inflation expectations remained relatively well anchored. Responses could be stronger during a recession, financial crisis, or period of already elevated inflation.

Repeated announcements may also have cumulative effects even when the response to an individual event is modest. The asymmetry results show that adverse trade-war news has stronger effects than favorable news. If markets regard de-escalation announcements as less credible or less durable, positive news may not fully

reverse the effects of earlier escalation.

Market-based inflation compensation should therefore be interpreted together with survey evidence and broader financial indicators. Breakeven inflation is available at daily frequency and responds immediately to trade-war announcements, but it also contains expected inflation, inflation risk compensation, TIPS liquidity effects, and other market-specific wedges (Gürkaynak et al., 2010). Survey expectations are less exposed to TIPS-market liquidity conditions, but they may respond slowly because of information frictions, rational inattention, and heterogeneous beliefs (Mankiw et al., 2004; Coibion et al., 2020).

Qualifications

The event-day heteroskedasticity design identifies market reactions around observable trade-war announcements, but it is less suited to measuring anticipation, information leakage, or the gradual accumulation of trade policy uncertainty between events. The state-dependent results should also be interpreted as conditional transmission exercises rather than as estimates of separate uncertainty shocks. Political-regime differences are suggestive because event days are unevenly distributed across administrations and relatively few events occur during the Biden and Trump II periods.

The main policy implication is that trade-war shocks should be viewed as macro-financial disturbances rather than solely as tariff-induced relative-price shocks. Their effects on price stability depend on the balance between import-price pass-through, expected activity, uncertainty, and risk-premium channels. In the daily financial-market responses examined here, weaker-growth expectations and higher risk premia dominate persistent cost-push inflation pressure.

8 Conclusion

This paper examines how trade-war shocks affect market-based inflation compensation and broader financial conditions. Using daily U.S. data and identification through event-day heteroskedasticity, we recover a structural trade-war shock and estimate its dynamic effects on inflation compensation, equity returns, Treasury yields, volatility, and corporate credit spreads.

Adverse trade-war shocks lower inflation compensation, particularly at the two-year and five-year maturities. Equity prices and long-term Treasury yields also fall, while volatility and credit spreads rise. This pattern indicates that financial markets interpret trade-war escalations mainly as shocks to expected activity, risk sentiment, and financial conditions rather than as persistent cost-push inflation disturbances. The smaller response at the ten-year maturity suggests that long-run inflation anchoring remains relatively stable.

The extension results reinforce this interpretation. The baseline response is broadly robust across uncertainty environments and is driven mainly by the higher-frequency component of the identified shock. Negative trade-war news produces stronger responses than positive news, while the political-regime results indicate some heterogeneity across administrations but should be interpreted cautiously because event days are unevenly distributed across regimes. The inflation-compensation decomposition further shows that the decline is associated mainly with expected inflation and the residual premium component, while the transaction-volume-based TIPS liquidity component is small.

The estimated effect is economically meaningful. A one-standard-deviation adverse trade-war shock lowers five-year inflation compensation by about 1.08 basis points on impact. A typical escalation event has a smaller effect, but repeated announcements can produce a meaningful cumulative decline. The estimated impact is also comparable in magnitude to the response to an externally identified high-frequency monetary-policy tightening shock.

The results imply that trade-war shocks should be treated as macro-financial disturbances rather than solely as tariff-induced price shocks. Central banks should interpret movements in breakeven inflation together with survey expectations, Treasury yields, equity prices, volatility, credit spreads, and decomposition evidence. Although the event-day design cannot fully capture anticipation, information leakage, or the gradual accumulation of trade policy uncertainty, the evidence shows that trade-war shocks affect short-run price stability mainly through weaker-growth and risk-premium channels without materially undermining longer-run nominal anchoring.

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Appendix

A Additional Methodological Details

This appendix provides the technical implementation of the state-dependent, frequency-domain, political-regime, directional-asymmetry, inflation-compensation decomposition, and identification-diagnostic exercises described in Section 3.

A.1 State-Dependent Local Projections

Let s_t denote the standardized trade-war shock and let U_{t-1} denote a lagged standardized uncertainty measure:

$$U_{t-1} \in \{zTPU_{t-1}, zEPU_{t-1}\}. \quad (30)$$

For each variable y_t^j and horizon $h = 0, 1, \dots, H$, we estimate

$$\begin{aligned} y_{t+h}^j - y_{t-1}^j = & \alpha_h^j + \beta_h^j s_t + \theta_h^j (s_t \times U_{t-1}) + \gamma_h^j s_{t-1} \\ & + \kappa_h^j (s_{t-1} \times U_{t-2}) + \rho_h^j U_{t-1} + \delta_h^j y_{t-1}^j + \varepsilon_{t+h}^j. \end{aligned} \quad (31)$$

The coefficient β_h^j measures the response when lagged uncertainty is at its sample mean, while θ_h^j measures how transmission varies with the pre-existing uncertainty environment. A positive value of θ_h^j implies that the response becomes more positive when uncertainty is higher, while a negative value implies that it becomes more negative.

The low- and high-uncertainty responses are evaluated at one standard deviation below and above the mean:

$$IRF_{h,\text{low}}^j = \beta_h^j - \theta_h^j, \quad IRF_{h,\text{high}}^j = \beta_h^j + \theta_h^j. \quad (32)$$

The TPU and EPU specifications are estimated separately. Confidence intervals are constructed from the relevant linear combinations of coefficients and account for the covariance between $\hat{\beta}_h^j$ and $\hat{\theta}_h^j$. The exercise is interpreted as a test of conditional transmission rather than as the identification of a separate uncertainty shock.

A.2 Frequency-Domain Construction

Let $\{w_t\}_{t=1}^T$ denote the identified trade-war shock series. Its discrete Fourier transform is

$$W(\omega_j) = \sum_{t=0}^{T-1} w_t e^{-i\omega_j t}, \quad \omega_j = \frac{2\pi j}{T}, \quad j = 0, \dots, T-1. \quad (33)$$

For cutoff frequency ω_c , the lower-frequency spectrum is

$$W^{LF}(\omega_j) = W(\omega_j) \mathbf{1}\{|\omega_j| \leq \omega_c\}, \quad (34)$$

and the higher-frequency spectrum is

$$W^{HF}(\omega_j) = W(\omega_j) \mathbf{1}\{|\omega_j| > \omega_c\}. \quad (35)$$

The inverse transformations are

$$w_t^{LF} = \frac{1}{T} \sum_{j=0}^{T-1} W^{LF}(\omega_j) e^{i\omega_j t}, \quad w_t^{HF} = \frac{1}{T} \sum_{j=0}^{T-1} W^{HF}(\omega_j) e^{i\omega_j t}. \quad (36)$$

By construction,

$$w_t = w_t^{LF} + w_t^{HF}. \quad (37)$$

The baseline cutoff corresponds to a 20-trading-day period. The higher-frequency component therefore captures fluctuations operating at horizons shorter than approximately one trading month, while the lower-frequency component captures movements at longer horizons.

The two standardized components enter the same local projection:

$$\begin{aligned} y_{t+h} - y_{t-1} = & \alpha_h + \beta_h^{LF} w_t^{LF} + \beta_h^{HF} w_t^{HF} + \gamma_h^{LF} w_{t-1}^{LF} \\ & + \gamma_h^{HF} w_{t-1}^{HF} + \rho_h y_{t-1} + u_{t+h}. \end{aligned} \quad (38)$$

The coefficients β_h^{LF} and β_h^{HF} measure the responses to the lower- and higher-frequency components of the identified shock.

This is a frequency-based decomposition rather than a Beveridge–Nelson permanent–transitory decomposition. The Beveridge–Nelson approach defines the permanent component from the limiting forecast of an integrated level series (Beveridge and Nelson, 1981; Oh et al., 2008). Our object is a standardized daily financial shock rather than an $I(1)$ macroeconomic level variable. The spectral approach is therefore better suited to separating short-lived announcement effects from more persistent repricing dynamics.

A.3 Political-Regime Heterogeneity

The full-sample trade-war shock is interacted with mutually exclusive regime indicators:

$$y_{t+h} - y_{t-1} = \alpha_h + \beta_{1,h} \left(s_t \times D_t^{\text{Trump I}} \right) + \beta_{2,h} \left(s_t \times D_t^{\text{Biden}} \right) + \beta_{3,h} \left(s_t \times D_t^{\text{Trump II}} \right) + \Gamma_h X_t + u_{t+h}, \quad (39)$$

where X_t contains the lagged dependent variable, lagged shock interactions, and any additional controls.

The coefficients $\beta_{1,h}$, $\beta_{2,h}$, and $\beta_{3,h}$ measure the responses during the Trump I, Biden, and Trump II regimes. Regime differences are evaluated using horizon-specific and response-window Wald tests. Because the numbers of Biden and Trump II events are small, the estimates are interpreted as suggestive evidence rather than as equally precise regime-specific structural responses.

A.4 Directional Asymmetry

The broader Trade War Event Database is used because it provides a systematic classification of adverse and favorable trade-war developments. Negative events include tariff announcements, tariff implementations, retaliatory actions, and other escalations. Positive events include tariff postponements, temporary truces, exemptions, rollback announcements, and progress toward negotiated settlements.

We first identify a common shock from the full event set and then define

$$s_t^- = s_t \mathbf{1} \{ \text{negative event at } t \}, \quad s_t^+ = s_t \mathbf{1} \{ \text{positive event at } t \}. \quad (40)$$

The asymmetric local projection is

$$y_{t+h}^j - y_{t-1}^j = \alpha_h^j + \beta_h^{-,j} s_t^- + \beta_h^{+,j} s_t^+ + \gamma_h^{-,j} s_{t-1}^- + \gamma_h^{+,j} s_{t-1}^+ + \delta_h^j y_{t-1}^j + \varepsilon_{t+h}^j. \quad (41)$$

The coefficients $\beta_h^{-,j}$ and $\beta_h^{+,j}$ measure the responses to adverse and favorable news within the same specification. The setup does not impose that the two responses must be equal in magnitude or opposite in sign.

A.5 Construction of the Inflation-Compensation Components

Inflation compensation is measured as

$$IC_t(h) = y_t^N(h) - y_t^{TIPS}(h), \quad (42)$$

where $y_t^N(h)$ and $y_t^{TIPS}(h)$ denote nominal Treasury and TIPS yields at maturity h .

Monthly annualized CPI inflation is

$$\pi_m = 12 [\log(CPI_m) - \log(CPI_{m-1})]. \quad (43)$$

Information in the nominal and real yield curves is summarized by

$$X_m = (PC_{1,m}^N, PC_{2,m}^N, PC_{3,m}^N, PC_{1,m}^{R\perp}, PC_{2,m}^{R\perp}, L_m)'. \quad (44)$$

The first three elements are principal components of the nominal yield curve. The real-yield factors are extracted from TIPS yield variation after controlling for the nominal factors and the liquidity factor, following the joint nominal–real yield-curve approach in Abrahams et al. (2016).

The relative transaction-volume ratio is

$$R_t = \frac{\text{Treasury ex-TIPS transactions}_t}{\text{TIPS transactions}_t}. \quad (45)$$

A higher value of R_t indicates weaker trading activity in TIPS relative to other Treasury securities. The standardized liquidity factor is

$$L_t = \frac{MA_{13}(R_t) - \overline{MA_{13}(R)}}{sd(MA_{13}(R))}. \quad (46)$$

The monthly version L_m is used in the estimation step, while the daily version L_t is used to construct the daily liquidity component.

Expected inflation is estimated from

$$\pi_{m+1} = a_\pi + b'_\pi X_m + u_{m+1}, \quad (47)$$

together with the state transition equation

$$X_{m+1} = c + \Phi X_m + v_{m+1}. \quad (48)$$

For maturity h , the monthly expected-inflation projection is

$$\widehat{EI}_m(h) = \frac{1}{H_h} \sum_{j=1}^{H_h} E_m[\pi_{m+j}], \quad (49)$$

where $H_h = 24, 60, 120$ for the two-year, five-year, and ten-year maturities.

After estimation at the monthly frequency, the forecasting system is evaluated at the daily state vector X_t^d :

$$\widehat{EI}_t(h) = \frac{1}{H_h} \sum_{j=1}^{H_h} E_t [\pi_{t+j} \mid X_t^d]. \quad (50)$$

The resulting series is a model-implied daily projection rather than a directly observed daily measure of expected inflation.

For each maturity, we estimate

$$IC_m(h) - \widehat{EI}_m(h) = \alpha_h + \lambda_h L_m + \xi_m(h). \quad (51)$$

Since a higher L_m indicates weaker relative TIPS liquidity, the daily TIPS liquidity-premium component is defined as

$$\widehat{LP}_t(h) = -\hat{\lambda}_h L_t. \quad (52)$$

Under this convention, a higher liquidity-premium component raises TIPS yields and lowers observed inflation compensation.

The residual premium component is

$$\widehat{RPC}_t(h) = IC_t(h) + \widehat{LP}_t(h) - \widehat{EI}_t(h). \quad (53)$$

For each component

$$z_t(h) \in \left\{ IC_t(h), \widehat{EI}_t(h), \widehat{LP}_t(h), \widehat{RPC}_t(h) \right\}, \quad (54)$$

we define

$$\Delta z_t(h) = z_t(h) - z_{t-1}(h). \quad (55)$$

We then estimate

$$\begin{aligned} \Delta z_{t+k}(h) - \Delta z_{t-1}(h) &= \alpha_{h,k} + \beta_{h,k} \widehat{w}_t + \gamma_{h,k} \widehat{w}_{t-1} \\ &\quad + \delta_{h,k} \Delta z_{t-1}(h) + \varepsilon_{t+k}(h), \end{aligned} \quad (56)$$

for $k = 0, 1, \dots, 60$. Comparing the responses of the expected-inflation, liquidity, and residual-premium components helps identify which parts of inflation compensation account for the baseline response.

A.6 Identification Diagnostics

Let λ_j denote the j -th eigenvalue of the covariance-difference matrix $\widehat{R}_0 - \widehat{R}_1$. The share of positive covariance-difference variation explained by the leading component is

$$\text{Share}_1 = \frac{\lambda_1}{\sum_{j=1}^K \max(\lambda_j, 0)}. \quad (57)$$

A large first eigenvalue, a small second eigenvalue, and a value of Share_1 close to one support the single-shock approximation. We also report the ratio of the variance of the projected trade-war shock on event days to its variance on non-event days:

$$VR = \frac{\text{Var}(\widehat{w}_t \mid t \in \mathcal{T}_0)}{\text{Var}(\widehat{w}_t \mid t \in \mathcal{T}_1)}. \quad (58)$$

Values of VR greater than one indicate that the identified shock is more volatile on trade-war event days. Randomization tests compare the observed covariance shift, leading eigenvalue, and variance ratio with distributions obtained by repeatedly reshuffling the event-day indicator while preserving the number of event observations.

A.7 Media-Saliency-Weighted Identification

This appendix describes the construction of the GDELT media-saliency weights and their incorporation into the event-day residual covariance matrix.

Let E denote the set of trade-war event days and let $\mathcal{B}_e$ denote the set of GDELT query blocks assigned to event day e . For each event-query pair, we collect GDELT's normalized TimelineVol measure, denoted $C_{e,b}^{TV}$, and the raw article count, denoted $C_{e,b}^{Raw}$.

When more than one query block is relevant to an event, we use the maximum coverage across the selected blocks:

$$M_e^{TV} = \max_{b \in \mathcal{B}_e} C_{e,b}^{TV}, \quad M_e^{Raw} = \max_{b \in \mathcal{B}_e} C_{e,b}^{Raw}. \quad (59)$$

Using the maximum rather than the sum reduces the risk of double-counting articles that match more than one query block.

The coverage measures are log transformed:

$$r_e^{TV} = \log(1 + M_e^{TV}), \quad r_e^{Raw} = \log(1 + M_e^{Raw}). \quad (60)$$

Each transformed measure is winsorized at the 95th percentile across event days. Let $\bar{r}_e^{TV}$ and $\bar{r}_e^{Raw}$

denote the winsorized values. The corresponding weights are normalized to have a mean of one:

$$w_e^{TV} = \frac{\bar{r}_e^{TV}}{|E|^{-1} \sum_{j \in E} \bar{r}_j^{TV}}, \quad w_e^{Raw} = \frac{\bar{r}_e^{Raw}}{|E|^{-1} \sum_{j \in E} \bar{r}_j^{Raw}}. \quad (61)$$

The preliminary hybrid weight is defined as

$$\tilde{w}_e^H = \begin{cases} w_e^{TV}, & \text{if TimelineVol is available,} \\ w_e^{Raw}, & \text{if TimelineVol is unavailable but the raw article count is available,} \\ 1, & \text{if neither measure is available.} \end{cases} \quad (62)$$

The hybrid weight is then re-normalized to have a mean of one:

$$w_e^H = \frac{\tilde{w}_e^H}{|E|^{-1} \sum_{j \in E} \tilde{w}_j^H}. \quad (63)$$

The effective number of event days is

$$N_{\text{eff}} = \frac{(\sum_{e \in E} w_e^H)^2}{\sum_{e \in E} (w_e^H)^2}. \quad (64)$$

Let $\hat{u}_t$ denote the reduced-form VAR residuals. The unweighted non-event residual mean is

$$\bar{u}_N = \frac{1}{|N|} \sum_{t \in N} \hat{u}_t, \quad (65)$$

and the corresponding covariance matrix is

$$\hat{\Sigma}_N = \frac{1}{|N|} \sum_{t \in N} (\hat{u}_t - \bar{u}_N) (\hat{u}_t - \bar{u}_N)'. \quad (66)$$

For the media-weighted event regime, the weighted residual mean is

$$\bar{u}_E^H = \frac{\sum_{t \in E} w_t^H \hat{u}_t}{\sum_{t \in E} w_t^H}. \quad (67)$$

The weighted event covariance matrix is

$$\hat{\Sigma}_E^H = \frac{\sum_{t \in E} w_t^H (\hat{u}_t - \bar{u}_E^H) (\hat{u}_t - \bar{u}_E^H)'}{\sum_{t \in E} w_t^H}. \quad (68)$$

The media-weighted covariance difference is

$$\hat{D}^H = \hat{\Sigma}_E^H - \hat{\Sigma}_N. \quad (69)$$

The media-weighted impact vector is estimated by fitting the rank-one restriction

$$\hat{D}^H \approx a^H (a^H)' \quad (70)$$

using the same minimum-distance GMM estimator and sign normalization as in the baseline specification.

After recovering the media-weighted shock s_t^H , we estimate

$$y_{t+h} - y_{t-1} = \alpha_h + \beta_h s_t^H + \gamma_h s_{t-1}^H + \delta_h y_{t-1} + \varepsilon_{t+h,h}, \quad (71)$$

where y_t denotes five-year inflation compensation. Newey–West standard errors use a lag length equal to the projection horizon. The coefficient β_h measures the response to a one-standard-deviation adverse trade-war shock.

B Trade War Event Days

The paper uses two trade-war event indicators. The baseline indicator is the main event set used for identification. It is deliberately narrow and aligned with the trading day on which financial markets are most likely to respond to the announcement. When an announcement occurs after market close, on a non-trading day, or when the relevant market reaction is more naturally assigned to the next trading day, the baseline event date is shifted to the corresponding trading day in the daily financial data.

The extended trade-war event database uses a broader definition of trade-war news. It includes additional trade-policy announcements beyond the baseline set and is closer to the original source or announcement date recorded in the event database. As a result, the two indicators need not flag the same calendar day even when they refer to closely related events. Some baseline event days appear only in the baseline indicator because they reflect trading-day alignment, while the extended database may record the same or related announcement on the original source date. Conversely, the extended database includes additional dates that are not part of the narrower baseline event set.

For this reason, the extended event specification should be interpreted as a robustness check rather than a replacement for the baseline event set. The baseline indicator is designed to match trade-war announcements to daily financial-market reactions as closely as possible. The extended indicator asks whether the results

survive when the event set is broadened to include a wider range of trade-war and trade-policy news. The fact that the estimated response remains close to the baseline suggests that the main result is not driven by one particular event-coding rule.

Table 16 lists the strict set of trade-war event days used in the identification strategy.

Table 16: Baseline Trade War Event Days Used in the Empirical Analysis

Date	Event Description	News Sign
2017-04-20	U.S. opens a Section 232 national security investigation	Negative
2017-04-27	U.S. launches further Section 232 investigations	Negative
2017-08-18	USTR begins a Section 301 investigation into China	Negative
2017-10-31	USITC recommends remedies on solar panel imports	Negative
2017-11-21	USITC recommends remedies on washing machine imports	Negative
2018-01-22	U.S. imposes safeguard tariffs on solar panels and washing machines	Negative
2018-02-16	U.S. Commerce publishes the Section 232 reports on steel and aluminum	Negative
2018-03-01	Trump announces planned tariffs on steel and aluminum	Negative
2018-03-09	Trump signs formal proclamations imposing steel and aluminum tariffs	Negative
2018-03-22	The administration releases the Section 301 findings and proposes tariffs on up to \$60 billion of Chinese imports	Negative
2018-04-02	China imposes retaliatory tariffs on \$3 billion of U.S. goods and responds to U.S. steel and aluminum tariffs	Negative
2018-04-03	The United States publishes a proposed tariff list covering Chinese imports	Negative
2018-04-04	China releases a retaliation list covering products including autos and aircraft	Negative
2018-04-06	Trump asks officials to examine an additional \$100 billion of tariffs on Chinese goods	Negative
2018-05-21	The United States and China announce a temporary pause in the trade conflict	Positive

Date	Event Description	News Sign
2018-05-29	The United States revives tariff plans after the temporary pause	Negative
2018-05-31	The United States proceeds with steel and aluminum tariffs on Canada, Mexico, and the European Union	Negative
2018-06-15	USTR confirms a two stage tariff plan on roughly \$50 billion of Chinese imports	Negative
2018-06-18	Trump directs USTR to prepare an additional \$200 billion tariff list	Negative
2018-06-19	China signals retaliation following the U.S. threat of an additional \$200 billion tariff list	Negative
2018-06-22	The European Union implements retaliation against selected U.S. products	Negative
2018-07-06	The first round of U.S. and Chinese tariffs takes effect	Negative
2018-07-10	USTR releases the proposed \$200 billion tariff schedule	Negative
2018-07-11	The proposed \$200 billion tariff schedule is absorbed in markets following its release	Negative
2018-07-20	Trump threatens tariffs on all remaining Chinese imports	Negative
2018-07-25	The United States and the European Union signal efforts to reduce tariffs	Positive
2018-08-02	USTR states that higher tariff rates are being considered for the proposed \$200 billion list	Negative
2018-08-07	USTR finalizes the second tranche of the \$50 billion tariff package	Negative
2018-08-23	The second stage of the \$50 billion U.S.-China tariff package takes effect	Negative
2018-08-27	The United States and Mexico reach a preliminary agreement on trade	Positive
2018-09-17	The United States finalizes 10 percent tariffs on \$200 billion of Chinese goods	Negative
2018-09-18	China announces tariffs on \$60 billion of U.S. imports in retaliation	Negative

Date	Event Description	News Sign
2018-09-24	U.S. tariffs on \$200 billion of Chinese goods and Chinese retaliation on \$60 billion of U.S. goods take effect	Negative
2018-10-01	The United States and Canada reach a new trade agreement	Positive
2018-11-30	The United States-Mexico-Canada Agreement is signed	Positive
2018-12-03	The United States and China agree to a temporary truce after the G20 meeting	Positive
2018-12-06	Canada announces the arrest of Huawei's CFO at the request of the United States	Negative
2018-12-31	Trump reports progress after speaking with Xi	Positive
2019-01-07	U.S.-China trade negotiations resume	Positive
2019-01-22	The United States cancels preparatory trade talks with China	Negative
2019-01-31	Trump expresses optimism about ongoing trade negotiations	Positive
2019-02-07	Trump says he will not meet Xi before the March deadline	Negative
2019-02-15	Xi meets U.S. officials following trade talks in Beijing	Positive
2019-02-22	Trump postpones the planned tariff increase on \$200 billion of Chinese imports	Positive
2019-03-28	Reports suggest China is offering broader concessions on technology transfer	Positive
2019-04-03	Negotiators appear closer to a final trade agreement	Positive
2019-05-06	Trump announces a higher tariff rate on \$200 billion of Chinese imports and Lighthizer confirms that the tariff increase will proceed on May 10	Negative
2019-05-07	Lighthizer confirms that the tariff increase on Chinese imports will proceed on May 10	Negative
2019-05-10	The United States raises tariffs on the existing \$200 billion list	Negative
2019-05-13	China responds with its own tariff increases	Negative
2019-05-20	Canada removes retaliatory countermeasures	Positive
2019-05-31	Trump announces proposed tariffs on Mexico	Negative
2019-06-10	Trump withdraws the proposed tariffs on Mexico	Positive
2019-07-01	Trump and Xi agree to restart trade negotiations at the G20 summit	Positive

Date	Event Description	News Sign
2019-07-16	Trump says a deal with China is still distant and threatens more tariffs	Negative
2019-08-01	Trump announces tariffs on most of the remaining Chinese imports	Negative
2019-08-05	The Chinese currency weakens sharply as trade tensions intensify	Negative
2019-08-13	USTR postpones part of the new tariff package on Chinese goods	Positive
2019-08-23	Trump escalates tariffs further after China announces retaliation on \$75 billion of U.S. goods	Negative
2019-09-11	China removes some tariffs and Trump postpones a planned tariff increase	Positive
2019-09-20	A Chinese delegation cancels a planned visit to U.S. farms	Negative
2019-10-11	Trump cancels the October 15 tariff increase and refers to Phase One of a deal	Positive
2019-12-12	Trump signals progress toward a deal and cancels the tariff increase scheduled for December 15	Positive
2020-01-15	The United States and China sign the Phase One agreement	Positive
2020-02-14	The Phase One agreement comes into force	Positive
2021-02-08	China is reported to have fallen short of its Phase One purchase commitments	Negative
2021-10-04	USTR sets out the Biden administration's China trade policy	Ambiguous
2022-03-23	USTR restores selected Section 301 exclusions	Positive
2024-05-14	The United States announces higher tariffs on selected Chinese products	Negative
2024-09-13	USTR finalizes the 2024 tariff increases	Negative
2024-09-27	The 2024 Section 301 tariff tranche becomes effective	Negative
2025-02-04	The United States implements a new tariff increase on Chinese imports	Negative
2025-03-04	A further tariff increase on Chinese imports takes effect	Negative
2025-04-09	A sharp additional tariff increase is implemented	Negative
2025-04-10	Another tariff increase affecting selected goods is introduced	Negative
2025-04-11	Sector specific carveout adjustments are implemented	Positive

Table 17: Extended trade-war event database

Date	Event description	Sign
2017-04-20	U.S. opens a Section 232 national security investigation	Negative
2017-04-27	U.S. launches further Section 232 investigations	Negative
2017-08-18	USTR begins a Section 301 investigation into China	Negative
2017-10-31	USITC recommends remedies on solar panel imports	Negative
2017-11-21	USITC recommends remedies on washing machine imports	Negative
2018-01-22	U.S. imposes safeguard tariffs on solar panels and washing machines	Negative
2018-02-16	U.S. Commerce publishes the Section 232 reports on steel and aluminum	Negative
2018-03-01	Trump announces planned tariffs on steel and aluminum	Negative
2018-03-07	The European Union announces a planned retaliatory response to prospective U.S. steel and aluminum tariffs	Negative
2018-03-08	Trump signs formal proclamations imposing steel and aluminum tariffs	Negative
2018-03-22	The administration releases the Section 301 findings and proposes tariffs on up to \$60 billion of Chinese imports; Trump issues revised steel and aluminum tariff proclamations temporarily exempting the EU, South Korea, Brazil, Argentina, and Australia	Negative; Positive
2018-03-23	U.S. steel and aluminum tariffs go into effect with exemptions for selected countries	Negative
2018-03-28	South Korea receives a permanent exemption for steel but accepts a quota	Positive
2018-04-02	China imposes retaliatory tariffs on \$3 billion of U.S. goods; China responds to U.S. steel and aluminum tariffs	Negative
2018-04-03	The United States publishes a proposed tariff list covering Chinese imports	Negative
2018-04-04	China releases a retaliation list covering products including autos and aircraft	Negative
2018-04-05	Trump asks officials to examine an additional \$100 billion of tariffs on Chinese goods	Negative
2018-04-30	The U.S. extends steel and aluminum tariff exemptions for the EU, Canada, and Mexico until June 1, 2018	Positive

Date	Event description	Sign
2018-05-21	The United States and China announce a temporary pause in the trade conflict	Positive
2018-05-23	The Commerce Department initiates a national security investigation into imported autos and parts	Negative
2018-05-29	The United States revives tariff plans after the temporary pause	Negative
2018-05-31	The United States proceeds with steel and aluminum tariffs on Canada, Mexico, and the European Union	Negative
2018-06-15	USTR confirms a two-stage tariff plan on roughly \$50 billion of Chinese imports	Negative
2018-06-18	Trump directs USTR to prepare an additional \$200 billion tariff list	Negative
2018-06-22	The European Union implements retaliation against selected U.S. products	Negative
2018-07-02	Canada imposes retaliatory tariffs on U.S. products after steel and aluminum tariffs	Negative
2018-07-06	The first round of U.S. and Chinese tariffs takes effect	Negative
2018-07-10	USTR releases the proposed \$200 billion tariff schedule	Negative
2018-07-16	The U.S. files WTO challenges against Canada, China, the EU, Mexico, and Turkey over retaliatory tariffs	Negative
2018-07-20	Trump threatens tariffs on all remaining Chinese imports	Negative
2018-07-25	The United States and the European Union signal efforts to reduce tariffs	Positive
2018-08-01	USTR states that higher tariff rates are being considered for the proposed \$200 billion list	Negative
2018-08-07	USTR finalizes the second tranche of the \$50 billion tariff package	Negative
2018-08-10	Trump raises steel and aluminum tariff rates on Turkey	Negative
2018-08-14	Turkey announces retaliatory tariffs on U.S. imports	Negative
2018-08-23	The second stage of the \$50 billion U.S.-China tariff package takes effect	Negative
2018-08-27	The United States and Mexico reach a preliminary agreement on trade	Positive
2018-09-17	The United States finalizes 10 percent tariffs on \$200 billion of Chinese goods	Negative
2018-09-18	China announces tariffs on \$60 billion of U.S. imports in retaliation	Negative

Date	Event description	Sign
2018-09-24	U.S. tariffs on \$200 billion of Chinese goods and Chinese retaliation on \$60 billion of U.S. goods take effect	Negative
2018-10-01	The United States and Canada reach a new trade agreement	Positive
2018-11-30	The United States-Mexico-Canada Agreement is signed	Positive
2018-12-03	The United States and China agree to a temporary truce after the G20 meeting	Positive
2018-12-06	Canada announces the arrest of Huawei's CFO at the request of the United States	Negative
2018-12-31	Trump reports progress after speaking with Xi	Positive
2019-01-07	U.S.-China trade negotiations resume	Positive
2019-01-22	The United States cancels preparatory trade talks with China	Negative
2019-01-31	Trump expresses optimism about ongoing trade negotiations	Positive
2019-02-07	Trump says he will not meet Xi before the March deadline	Negative
2019-02-15	Xi meets U.S. officials following trade talks in Beijing	Positive
2019-02-22	Trump postpones the planned tariff increase on \$200 billion of Chinese imports	Positive
2019-03-28	Reports suggest China is offering broader concessions on technology transfer	Positive
2019-04-03	Negotiators appear closer to a final trade agreement	Positive
2019-05-06	Trump announces a higher tariff rate on \$200 billion of Chinese imports; Lighthizer confirms that the tariff increase will proceed on May 10	Negative
2019-05-10	The United States raises tariffs on the existing \$200 billion list	Negative
2019-05-13	China responds with its own tariff increases	Negative
2019-05-17	The United States agrees to remove steel and aluminum tariffs on Canada and Mexico; Trump delays a decision on whether to impose auto tariffs and opens negotiations with the EU, Japan, and others	Positive
2019-05-20	Canada removes retaliatory countermeasures	Positive
2019-05-30	Trump announces proposed tariffs on Mexico	Negative
2019-06-07	Trump withdraws the proposed tariffs on Mexico	Positive
2019-06-17	India implements retaliatory tariffs against U.S. exports in response to Trump's metal tariffs	Negative

Date	Event description	Sign
2019-07-01	Trump and Xi agree to restart trade negotiations at the G20 summit	Positive
2019-07-16	Trump says a deal with China is still distant and threatens more tariffs	Negative
2019-08-01	Trump announces tariffs on most of the remaining Chinese imports	Negative
2019-08-05	The Chinese currency weakens sharply as trade tensions intensify	Negative
2019-08-13	USTR postpones part of the new tariff package on Chinese goods	Positive
2019-08-23	Trump escalates tariffs further after China announces retaliation on \$75 billion of U.S. goods	Negative
2019-09-11	China removes some tariffs and Trump postpones a planned tariff increase	Positive
2019-09-20	A Chinese delegation cancels a planned visit to U.S. farms	Negative
2019-10-11	Trump cancels the October 15 tariff increase and refers to Phase One of a deal	Positive
2019-12-12	Trump signals progress toward a deal and cancels the tariff increase scheduled for December 15	Positive
2019-12-19	The U.S. House approves the new trade agreement with Canada and Mexico	Positive
2020-01-15	The United States and China sign the Phase One agreement	Positive
2020-01-24	Trump broadens steel and aluminum tariffs to additional derivative products	Negative
2020-02-14	The Phase One agreement comes into force	Positive
2020-08-06	Trump reimposes tariffs on some Canadian aluminum and Canada plans retaliation	Negative
2020-09-15	The United States ends duties on Canadian aluminum retroactively to September 1	Positive
2021-02-01	Biden reinstates duties on aluminum from the UAE	Negative
2021-02-08	China is reported to have fallen short of its Phase One purchase commitments	Negative
2021-05-17	The U.S. and EU agree to discuss excess steel and aluminum capacity and the EU suspends scheduled retaliation	Positive
2021-11-01	The U.S. lifts tariffs on some steel and aluminum imports from the EU under a quota arrangement	Positive

Date	Event description	Sign
2022-02-04	Biden extends Trump-era solar tariffs	Negative
2022-02-07	The U.S. rolls back some tariffs on Japanese steel	Positive
2022-03-22	The U.S. eases tariffs on U.K. steel and aluminum	Positive
2022-03-23	USTR restores selected Section 301 exclusions	Positive
2022-05-09	The United States suspends Section 232 tariffs on Ukrainian steel for one year	Positive
2022-06-06	Biden temporarily suspends tariffs on Southeast Asian solar panels	Positive
2022-12-09	WTO rules against Trump's steel and aluminum tariffs	Positive
2023-01-27	The U.S. appeals the WTO ruling on steel and aluminum tariffs	Negative
2023-02-07	Safeguard tariffs on washing machines end	Positive
2023-02-24	The U.S. raises Russian steel and aluminum tariffs	Negative
2023-12-19	The EU prolongs tariff suspension for U.S. products related to the steel and aluminum dispute	Positive
2023-12-28	The U.S. extends the steel and aluminum tariff-rate quota policy toward the EU for two years	Positive
2024-05-14	The United States announces higher tariffs on selected Chinese products	Negative
2024-05-16	Biden further modifies Trump-era solar tariffs by ending the waiver and expanding coverage	Negative
2024-09-13	USTR finalizes the 2024 tariff increases	Negative
2024-09-27	The 2024 Section 301 tariff tranche becomes effective	Negative
2024-12-11	USTR imposes more new tariffs on China following the four-year review, including tungsten, wafers, and polysilicon	Negative
2025-02-04	The United States implements a new tariff increase on Chinese imports	Negative
2025-03-04	A further tariff increase on Chinese imports takes effect	Negative
2025-04-09	A sharp additional tariff increase is implemented	Negative
2025-04-10	Another tariff increase affecting selected goods is introduced	Negative
2025-04-11	Sector-specific carveout adjustments are implemented	Positive

Notes: The table reports the extended trade-war event database used in the robustness analysis. The date is the trading-day-aligned event date matched to the daily financial-market data. The sign classifies whether the event is expected to intensify trade-war risk, reduce trade-war risk, or contain mixed information. The extended database is broader than the baseline event set and is used as a robustness check.

C Identification Diagnostics for 2 year and 10 year inflation compensation

Table 18: Identification Diagnostics: Event-Day Heteroskedasticity (Two-Year Specification)

Statistic	Value
Event days (residual sample)	76
Non-event days (residual sample)	2105
Largest eigenvalue of $R_0 - R_1$	3.9897
Second eigenvalue of $R_0 - R_1$	-0.0000
Eigenvalue ratio $\lambda_1/ \lambda_2 $	408185.40
Share of first positive eigenvalue	1.0000
Frobenius norm $\ R_0 - R_1\ _F$	3.9897
Variance ratio of identified shock	2.0154

Notes: R_0 and R_1 denote covariance matrices of reduced-form residuals on trade-war event and non-event days, respectively. The two-year specification replaces five-year inflation compensation with two-year inflation compensation while keeping the remaining variables unchanged.

Table 19: Identification Diagnostics: Event-Day Heteroskedasticity (Ten-Year Specification)

Statistic	Value
Event days (residual sample)	76
Non-event days (residual sample)	2105
Largest eigenvalue of $R_0 - R_1$	4.7123
Second eigenvalue of $R_0 - R_1$	-0.0000
Eigenvalue ratio $\lambda_1/ \lambda_2 $	357263.47
Share of first positive eigenvalue	1.0000
Frobenius norm $\ R_0 - R_1\ _F$	4.7123
Variance ratio of identified shock	2.1895

Notes: R_0 and R_1 denote covariance matrices of reduced-form residuals on trade-war event and non-event days, respectively. The ten-year specification replaces five-year inflation compensation with ten-year inflation compensation while keeping the remaining variables unchanged.

D Additional Decomposition Results

Responses of Daily Changes in Two-Year Inflation-Compensation Components to the Identified Trade-War Shock

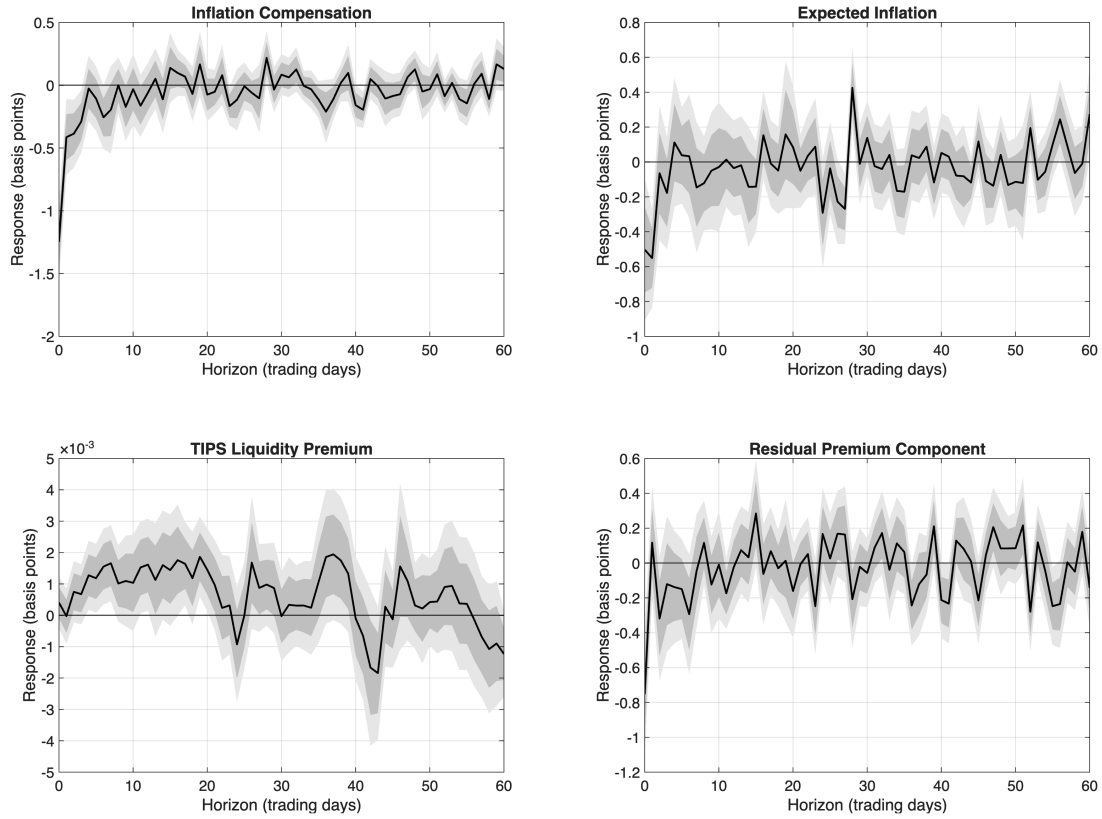


Figure 11: Responses of Daily Changes in Two-Year Inflation-Compensation Components

Responses of Daily Changes in Ten-Year Inflation-Compensation Components to the Identified Trade-War Shock

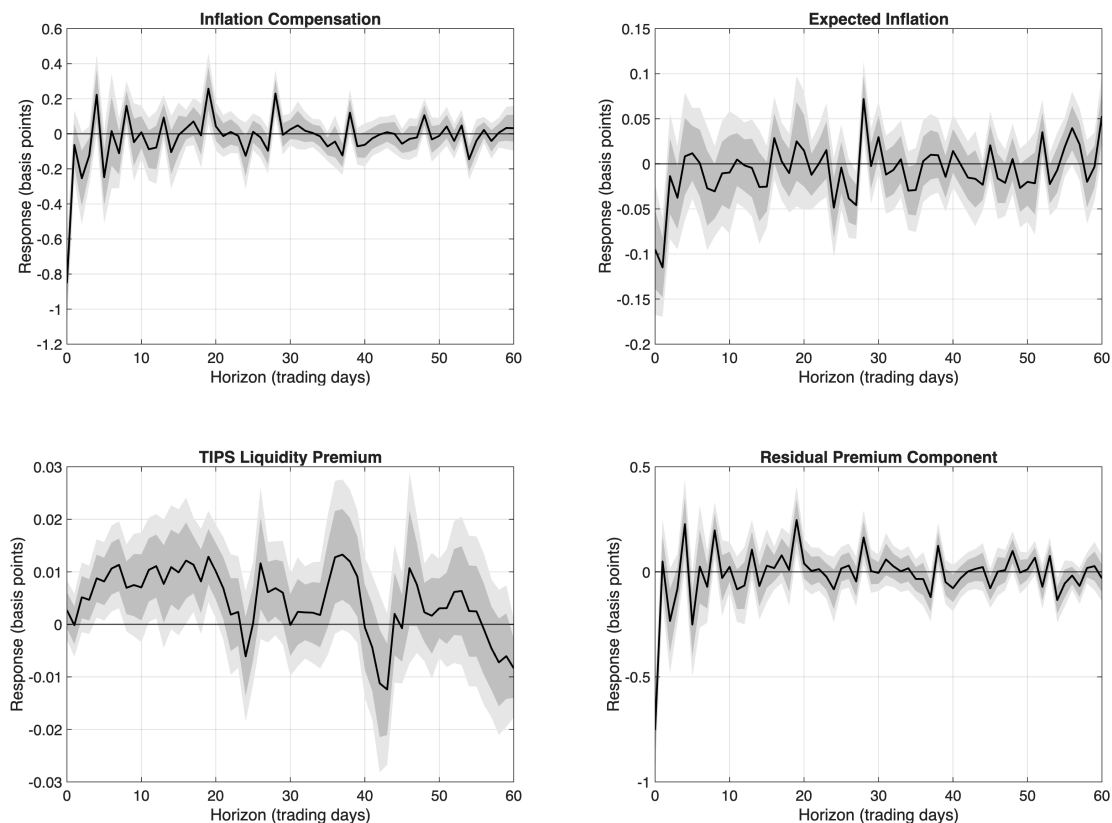


Figure 12: Responses of Daily Changes in Ten-Year Inflation-Compensation Components

E Construction of GDELT Media-Salience Weights

Table 20: GDELT Query Blocks Used to Construct Media-Salience Weights

No.	Query block	Keywords included
1	China tariff / trade war	U.S. China; US China; United States China; tariff; tariffs; trade war; trade dispute
2	Phase One / trade agreement	U.S. China; US China; United States China; phase one; phase 1; trade deal; trade agreement; economic and trade agreement
3	Section 301 / USTR / China	China; Chinese; Section 301; USTR; Chinese imports; Chinese goods; tariff increase; tariff hike; retaliatory tariffs
4	Section 232 / steel and aluminum	Section 232; steel tariffs; aluminum tariffs; steel and aluminum tariffs; national security tariffs
5	EU trade retaliation	European Union; EU; Juncker; tariff; tariffs; retaliatory tariffs; trade retaliation; trade countermeasures; trade truce
6	Mexico / NAFTA / USMCA	Mexico; Mexican; NAFTA; USMCA; tariff; tariffs; Mexico tariffs; trade agreement; trade deal; trade dispute

No.	Query block	Keywords included
7	Huawei / export controls	Huawei; entity list; export controls; China; Chinese; trade war; tariff; tariffs; trade restrictions; national security
8	IEEPA / fentanyl tariffs	IEEPA; fentanyl; tariff; tariffs; additional tariff; tariff increase

Notes: The table reports the keyword groups used to construct event-specific GDELT media-salience weights. For each trade-war event day, the query block is selected according to the event description and event category. When more than one query block is relevant to an event, we use the maximum coverage across selected blocks to avoid double-counting articles that match multiple keyword groups. The preferred hybrid media weight uses TimelineVol coverage when available, raw article counts when TimelineVol is unavailable, and a neutral weight of one for the remaining missing observations. The resulting weights are winsorized and normalized to have mean one across event days.

E.1 Date changes in identification-sensitivity specifications

This appendix reports the calendar-date changes underlying the identification-sensitivity specifications in Section 6.1. The baseline specification uses the main set of 76 trade-war event days and does not remove additional observations. The clean-event specification removes trade-war event days that either overlap with major macroeconomic announcements or are classified as ambiguous. The no-macro-news specification removes all FOMC announcement days, CPI release days, and employment-report release days from the estimation sample. The event-buffer specification does not remove trade-war event days; instead, it removes the adjacent non-event trading days immediately before and after trade-war events from the non-event regime. The VAR-control specification does not remove dates, but instead augments the reduced-form VAR with macroeconomic and monetary controls. The extended-TWDB specification changes the trade-war event indicator by using the broader TWDB event set.

Table 21: Summary of date changes in identification-sensitivity specifications

Specification	Date-change rule	Number of dates affected
Baseline	No additional dates removed	0
Clean events	Remove ambiguous trade-war events and trade-war events overlapping with major macro-news days	8
No macro-news days	Remove all FOMC announcement days, CPI release days, and employment-report release days from the estimation sample	322
Event-buffer specification	Remove one trading day before and after each trade-war event from the non-event regime	124
VAR with macro and monetary controls	No dates removed; macro and monetary variables are included as VAR controls	0
Extended TWDB	Replace the baseline event indicator with the extended TWDB event indicator	36 added; 9 baseline-only

Notes: The table summarizes how each identification-sensitivity specification changes the estimation sample or event set. The event-buffer specification removes adjacent non-event observations from the non-event covariance regime but keeps the trade-war event days themselves. The extended-TWDB specification is not a deletion rule; it uses a broader event indicator, which adds some event days relative to the baseline and excludes some baseline trading-day-aligned dates because of differences in event-date coding.

Table 22: Dates affected by identification-sensitivity specifications

Specification	Category	Dates
Clean events	NFP overlap	2018-03-09; 2018-04-06; 2018-07-06
Clean events	CPI overlap	2019-05-10; 2019-08-13; 2021-02-08; 2025-04-10
Clean events	Ambiguous event	2021-10-04
No macro-news days	CPI, 2016	2016-01-20; 2016-02-19; 2016-04-14; 2016-05-17; 2016-06-16; 2016-07-15; 2016-08-16; 2016-09-16; 2016-10-18; 2016-11-17; 2016-12-15
No macro-news days	CPI, 2017	2017-01-18; 2017-02-13; 2017-02-15; 2017-05-12; 2017-07-14; 2017-08-11; 2017-09-14; 2017-10-13; 2017-11-15
No macro-news days	CPI, 2018	2018-01-12; 2018-02-14; 2018-03-13; 2018-04-11; 2018-05-10; 2018-06-12; 2018-07-12; 2018-08-10; 2018-09-13; 2018-10-11; 2018-11-14; 2018-12-12
No macro-news days	CPI, 2019	2019-01-11; 2019-02-11; 2019-02-13; 2019-03-12; 2019-04-10; 2019-05-10; 2019-06-12; 2019-07-11; 2019-08-13; 2019-09-12; 2019-10-10; 2019-11-13
No macro-news days	CPI, 2020	2020-01-14; 2020-02-11; 2020-02-13; 2020-03-11; 2020-05-12; 2020-07-14; 2020-08-12; 2020-09-11; 2020-10-13; 2020-11-12; 2020-12-10
No macro-news days	CPI, 2021	2021-01-13; 2021-02-08; 2021-02-10; 2021-03-10; 2021-04-13; 2021-05-12; 2021-06-10; 2021-07-13; 2021-08-11; 2021-09-14; 2021-10-13; 2021-11-10; 2021-12-10
No macro-news days	CPI, 2022	2022-01-12; 2022-02-08; 2022-02-10; 2022-03-10; 2022-04-12; 2022-05-11; 2022-06-10; 2022-07-13; 2022-08-10; 2022-09-13; 2022-10-13; 2022-11-10; 2022-12-13
No macro-news days	CPI, 2023	2023-01-12; 2023-02-10; 2023-02-14; 2023-03-14; 2023-04-12; 2023-05-10; 2023-06-13; 2023-07-12; 2023-08-10; 2023-09-13; 2023-10-12; 2023-11-14; 2023-12-12
No macro-news days	CPI, 2024	2024-01-11; 2024-02-09; 2024-02-13; 2024-03-12; 2024-04-10; 2024-05-15; 2024-07-11; 2024-08-14; 2024-09-11; 2024-10-10; 2024-11-13; 2024-12-11
No macro-news days	CPI, 2025	2025-01-15; 2025-02-12; 2025-03-12; 2025-04-10; 2025-05-13; 2025-06-11; 2025-07-15; 2025-08-12; 2025-09-11; 2025-10-24; 2025-12-18
No macro-news days	CPI, 2026	2026-01-13
No macro-news days	FOMC, 2016	2016-01-27; 2016-04-27; 2016-06-15; 2016-07-27; 2016-09-21; 2016-11-02; 2016-12-14
No macro-news days	FOMC, 2017	2017-02-01; 2017-05-03; 2017-07-26; 2017-09-20; 2017-11-01
No macro-news days	FOMC, 2018	2018-01-31; 2018-05-02; 2018-06-13; 2018-08-01; 2018-09-26; 2018-11-08; 2018-12-19

Specification	Category	Dates
No macro-news days	FOMC, 2019	2019-01-30; 2019-03-20; 2019-05-01; 2019-06-19; 2019-07-31; 2019-09-18; 2019-10-30
No macro-news days	FOMC, 2020	2020-01-29; 2020-03-03; 2020-03-23; 2020-03-31; 2020-04-29; 2020-07-29; 2020-08-27; 2020-09-16; 2020-11-05; 2020-12-16
No macro-news days	FOMC, 2021	2021-01-27; 2021-03-17; 2021-04-28; 2021-06-16; 2021-07-28; 2021-09-22; 2021-11-03; 2021-12-15
No macro-news days	FOMC, 2022	2022-01-26; 2022-03-16; 2022-05-04; 2022-06-15; 2022-07-27; 2022-09-21; 2022-11-02; 2022-12-14
No macro-news days	FOMC, 2023	2023-02-01; 2023-03-22; 2023-05-03; 2023-06-14; 2023-07-26; 2023-09-20; 2023-11-01; 2023-12-13
No macro-news days	FOMC, 2024	2024-01-31; 2024-03-20; 2024-05-01; 2024-07-31; 2024-09-18; 2024-11-07; 2024-12-18
No macro-news days	FOMC, 2025	2025-01-29; 2025-03-19; 2025-05-07; 2025-06-18; 2025-07-30; 2025-09-17; 2025-10-29; 2025-12-10
No macro-news days	FOMC, 2026	2026-01-28
No macro-news days	FOMC and CPI overlap	2016-03-16; 2017-03-15; 2017-06-14; 2017-12-13; 2019-12-11; 2020-06-10; 2024-06-12
No macro-news days	NFP, 2016	2016-01-08; 2016-02-05; 2016-03-04; 2016-04-01; 2016-05-06; 2016-06-03; 2016-07-08; 2016-08-05; 2016-09-02; 2016-10-07; 2016-11-04; 2016-12-02
No macro-news days	NFP, 2017	2017-01-06; 2017-02-03; 2017-03-10; 2017-04-07; 2017-05-05; 2017-06-02; 2017-07-07; 2017-08-04; 2017-09-01; 2017-10-06; 2017-11-03; 2017-12-08
No macro-news days	NFP, 2018	2018-01-05; 2018-02-02; 2018-03-09; 2018-04-06; 2018-05-04; 2018-06-01; 2018-07-06; 2018-08-03; 2018-09-07; 2018-10-05; 2018-11-02; 2018-12-07
No macro-news days	NFP, 2019	2019-01-04; 2019-02-01; 2019-03-08; 2019-04-05; 2019-05-03; 2019-06-07; 2019-07-05; 2019-08-02; 2019-09-06; 2019-10-04; 2019-11-01; 2019-12-06
No macro-news days	NFP, 2020	2020-01-10; 2020-02-07; 2020-03-06; 2020-04-03; 2020-05-08; 2020-05-11; 2020-06-05; 2020-07-02; 2020-08-07; 2020-09-04; 2020-10-02; 2020-11-06; 2020-12-04
No macro-news days	NFP, 2021	2021-01-08; 2021-02-05; 2021-03-05; 2021-05-07; 2021-06-04; 2021-07-02; 2021-08-06; 2021-09-03; 2021-10-08; 2021-11-05; 2021-12-03
No macro-news days	NFP, 2022	2022-01-07; 2022-02-04; 2022-03-04; 2022-04-01; 2022-05-06; 2022-06-03; 2022-07-08; 2022-08-05; 2022-09-02; 2022-10-07; 2022-11-04; 2022-12-02

Specification	Category		Dates
No macro-news days	NFP, 2023		2023-01-06; 2023-02-03; 2023-03-10; 2023-05-05; 2023-06-02; 2023-07-07; 2023-08-04; 2023-09-01; 2023-10-06; 2023-11-03; 2023-12-08
No macro-news days	NFP, 2024		2024-01-05; 2024-01-10; 2024-02-02; 2024-03-08; 2024-04-05; 2024-05-03; 2024-06-07; 2024-07-05; 2024-08-02; 2024-08-21; 2024-09-06; 2024-10-04; 2024-11-01; 2024-12-06
No macro-news days	NFP, 2025		2025-01-10; 2025-02-07; 2025-03-07; 2025-04-04; 2025-05-02; 2025-06-06; 2025-07-03; 2025-08-01; 2025-09-05; 2025-11-20; 2025-12-16
No macro-news days	NFP, 2026		2026-01-09
Event-buffer specification	Adjacent days, 2017	non-event	2017-04-19; 2017-04-21; 2017-04-26; 2017-04-28; 2017-08-17; 2017-08-21; 2017-10-30; 2017-11-01; 2017-11-20; 2017-11-22
Event-buffer specification	Adjacent days, 2018	non-event	2018-01-19; 2018-01-23; 2018-02-15; 2018-02-20; 2018-02-28; 2018-03-05; 2018-03-08; 2018-03-12; 2018-03-20; 2018-03-23; 2018-03-29; 2018-04-05; 2018-04-09; 2018-05-18; 2018-05-22; 2018-05-25; 2018-05-30; 2018-06-01; 2018-06-14; 2018-06-20; 2018-06-21; 2018-06-25; 2018-07-05; 2018-07-09; 2018-07-12; 2018-07-19; 2018-07-23; 2018-07-24; 2018-07-26; 2018-08-01; 2018-08-03; 2018-08-06; 2018-08-08; 2018-08-22; 2018-08-24; 2018-08-28; 2018-09-14; 2018-09-19; 2018-09-21; 2018-09-25; 2018-09-28; 2018-10-02; 2018-11-29; 2018-12-04; 2018-12-07; 2018-12-28
Event-buffer specification	Adjacent days, 2019	non-event	2019-01-02; 2019-01-04; 2019-01-08; 2019-01-18; 2019-01-23; 2019-01-30; 2019-02-01; 2019-02-06; 2019-02-08; 2019-02-14; 2019-02-19; 2019-02-21; 2019-02-25; 2019-03-27; 2019-03-29; 2019-04-02; 2019-04-04; 2019-05-03; 2019-05-08; 2019-05-09; 2019-05-14; 2019-05-17; 2019-05-21; 2019-05-30; 2019-06-03; 2019-06-07; 2019-06-11; 2019-06-28; 2019-07-02; 2019-07-15; 2019-07-17; 2019-07-31; 2019-08-02; 2019-08-06; 2019-08-12; 2019-08-14; 2019-08-22; 2019-08-26; 2019-09-10; 2019-09-12; 2019-09-19; 2019-09-23; 2019-10-10; 2019-10-15; 2019-12-11; 2019-12-13
Event-buffer specification	Adjacent days, 2020	non-event	2020-01-14; 2020-01-16; 2020-02-13; 2020-02-18
Event-buffer specification	Adjacent days, 2021	non-event	2021-02-05; 2021-02-09; 2021-10-01; 2021-10-05
Event-buffer specification	Adjacent days, 2022	non-event	2022-03-22; 2022-03-24

Specification	Category	Dates
Event-buffer specification	Adjacent non-event days, 2024	2024-05-13; 2024-05-15; 2024-09-12; 2024-09-16; 2024-09-26; 2024-09-30
Event-buffer specification	Adjacent non-event days, 2025	2025-02-03; 2025-02-05; 2025-03-03; 2025-03-05; 2025-04-08; 2025-04-14
Extended TWDB	Added event days, negative	2018-03-07; 2018-03-08; 2018-03-23; 2018-04-05; 2018-05-23; 2018-07-02; 2018-07-16; 2018-08-01; 2018-08-10; 2018-08-14; 2019-05-30; 2019-06-17; 2020-01-24; 2020-08-06; 2021-02-01; 2022-02-04; 2023-01-27; 2023-02-24; 2024-05-16; 2024-12-11
Extended TWDB	Added event days, positive	2018-03-28; 2018-04-30; 2019-05-17; 2019-06-07; 2019-12-19; 2020-09-15; 2021-05-17; 2021-11-01; 2022-02-07; 2022-03-22; 2022-05-09; 2022-06-06; 2022-12-09; 2023-02-07; 2023-12-19; 2023-12-28

Notes: The table reports the dates affected by each identification-sensitivity specification. “Clean events” removes event days from the trade-war event set and the estimation sample. “No macro-news days” removes all FOMC, CPI, and employment-report release dates from the estimation sample. “Event-buffer specification” removes only adjacent non-event observations from the non-event covariance regime; it does not remove the trade-war event days themselves. “Extended TWDB” reports event-set changes relative to the baseline event indicator rather than a sample-exclusion rule.