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Are Tainted Elections Inflationary? Evidence from Haiti (2004–2018)

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Abstract

This paper studies whether elections generate short-run inflation bursts in Haiti, a fragile democracy marked by violence, disruptions, and widespread vote-buying. Using monthly CPI data from 2004–2018, we combine linear autoregressive regressions, (individual and stacked) event studies, and symmetric local projections. While linear specifications reveal limited statistical evidence of an electoral effect, dynamic methods consistently show that inflation rises in the months surrounding elections. The local projections imply a cumulative increase of roughly 0.54 percentage points in the consumer price index over the first three post-election months, corresponding to about 454 Haitian gourdes (HTG) - approximately 1.6% of a typical monthly salary in 2025. This short-lived inflationary impulse is economically meaningful in a context of widespread poverty. A series of placebo-based randomization test and robustness checks confirms that the pattern is unlikely to arise solely from calendar noise. Our results highlight the economic costs of electoral instability in fragile democracies.

Keywords: Vote buying, Electoral cycle, Time series analysis, Local projection.
JEL: D72, E31, O10

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1 Introduction

Elections in Haiti are rarely routine democratic exercises. They are frequently associated with heightened social tensions, disruptions to commerce, and episodes of violence, to the point that commentators regularly question whether the country would be better served by appointed political authorities rather than recurrent electoral crises ([Étienne, 2019]). Presidential, legislative, and municipal elections alike have repeatedly been marred by strikes, blockades, and wide-ranging disturbances.[1] These episodes coincide with noticeable increases in prices, raising the question of whether Haiti experiences an electoral cycle in inflation.

Several mechanisms could generate such patterns. One possibility is that inflation rises because of pre-electoral hoarding in an environment where violence and fraud are expected. Another is that governments deliberately engineer short-term political business cycles by distributing cash or goods to influence voters. A third mechanism is mechanical: protests, blockades, and transport disruptions frequently reduce the supply of goods reaching local markets. Distinguishing among these channels is challenging, particularly in an economy characterized by fragile institutions and recurrent shocks.

Haiti is a particularly relevant case for investigating these dynamics. The country exhibits widespread vote-buying ([Justesen and Manzetti, 2023]; [Woller et al., 2023]), consistent with evidence from many low-income democracies in which campaigns rely heavily on direct transfers of food, cash, or goods ([Aidt et al., 2020]). Such practices can leave a monetary footprint and influence prices during election months. More broadly, disruptions associated with electoral violence and mobilization are routinely described as contributing to market tightness and short-term inflationary pressures.

This paper examines whether elections in Haiti generate systematic and distinct movements in inflation. We use a monthly time series covering 2004–2018 from the comprehensive Consumer Price Index (CPI) published by the Haitian Institute of Statistics and Informatics (IHSI).

To identify the dynamic behavior of inflation around elections, we combine three empirical approaches. First, autoregressive linear regressions provide a parsimonious benchmark but impose the strong assumption that all electoral

[1] See, for instance, the New York Times article from January 2016: "Haiti postpones presidential runoff vote amid escalating violence."https://www.nytimes.com/2016/01/23/world/americas/haiti-postpones-presidential-runoff-vote.html

episodes generate identical effects. Second, we estimate both individual and stacked event-study profiles. These visualizations highlight substantial heterogeneity across elections but reveal a recurrent pattern of rising inflation in the months immediately surrounding the vote. Third, we employ the local projections method of [Jordà, 2005],[Jordà and Taylor, 2016] to recover impulse responses to elections. The symmetric LP specification confirms a mild but coherent acceleration in inflation in the one to two months following the election, with effects fading out after several months. Because Haiti experiences only five election sequences over the sample period, we complement these results with a placebo-based randomization test. While low-powered, the test indicates that the observed inflation response is larger than what is typically produced by randomly assigned pseudo-election dates.

Our findings suggest that elections in Haiti generate a short-lived increase in monthly inflation that is economically nontrivial but transitory, consistent with mechanisms such as short-run market disruptions or temporary liquidity injections associated with campaign practices. These results position Haiti as an instructive case for understanding the economic consequences of electoral processes in fragile democracies, where political competition may amplify rather than stabilize macroeconomic volatility.

Beyond statistical significance, we also assess the economic magnitude of the estimated effects. Translating the local-projection responses into a cumulative price-level change over the first three post-election months yields an inflationary impact of approximately 0.54 percentage points. Given a typical monthly salary of 28,000 Haitian gourdes (HTG) in Haiti in 2025, this corresponds to roughly 454 HTG of additional household expenditures during the quarter surrounding an election - i.e., about 1.6% of monthly income. In a setting where many households operate at or below subsistence levels, even short-lived inflationary bursts can impose meaningful welfare costs. Or, in magnitude, the 0.54pp cumulative inflationary effect of an election corresponds to roughly the inflation usually observed in a single month, although slightly below the sample average monthly rate (0.65%).

This paper contributes to several strands of research. First, it extends the literature on political and electoral cycles by examining inflation dynamics in a fragile democracy, a setting that has received little empirical attention relative to the large body of work on advanced and middle-income economies. However, these economies have lessons to deliver, as they are examples of countries where democratic processes exist, but may be used to benefit an oligarchy, and not the general population ([Rahman et al., 2022], for example, speak of "storm autocracies" about island countries). Haiti in particular has been defined as a "phantom state", as the political elite and the criminal groups converge to run the country ([Niño and González, 2022]).

Second, this paper contributes to the study of inflation in low-income and small open economies. It does so by exploiting a high-frequency CPI dataset for Haiti and by showing that electoral instability can generate short-lived but economically meaningful price movements. While the literature on vote-buying has become relatively large and has documented the cost of political campaigns ([Bekkouche et al., 2022], [François et al., 2022]), the consequences of vote-buying for consumption around elections are also now well established ([Aidt et al., 2020], [Mitra and Mitra, 2025]). Providing food or cash to voters, however, is only one of several ways to rig an election ([Cheeseman and Klaas, 2018]), even if it is a particularly prevalent strategy in weak democracies, often combined with violence and fraud ([Collier and Vicente, 2012]). By contrast, the impact of elections on inflation - operating through vote-buying practices - has so far remained unexplored. This paper fills this gap by analyzing how elections in a fragile state affect price dynamics.

Third, this paper contributes to the literature on vote-buying and clientelism by explicitly linking campaign-related transfers, electoral protests, and political disruptions to inflationary pressures in weak institutional environments. In doing so, it extends existing work on clientelism and political distortions ([Lundstedt and Edgell, 2022], [Hicken, 2011]) by providing quantitative evidence on a channel - price dynamics - that has so far received little attention.

Finally, by combining event studies, stacked designs, local projections, and randomization inference, the paper speaks to the methodological literature on identifying short-run shocks in small samples. To our knowledge, this is the first study to produce systematic monthly evidence on election-induced inflation cycles in a fragile state.

The rest of the paper proceeds as follows. Section 2 presents the electoral context of Haiti. Section 3 describes the data and the construction of the variables and outlines the empirical methodology. Section 4 reports the results from linear regressions, event studies, and local projections. Section 5 discusses the robustness of the results through placebo-based randomization inference and other tests. Section 6 concludes.

2 Electoral Context

The Republic of Haiti has been a sovereign and autonomous nation since January 1, 1804. The country has experienced various types of political regimes before adopting a democratic system. The history of elections in Haiti is complex and marked by periods of political instability.

Haiti experienced its first elections that could plausibly be described as democratic in 1946, following the collapse of the Lescot regime. However, this democratic opening was short-lived. In 1957, François Duvalier won the presidential election and rapidly dismantled democratic institutions, establishing a brutal dictatorship and banning political parties. After his death in 1971, his son Jean-Claude Duvalier assumed power and maintained the authoritarian regime until popular protests forced his departure in 1986. Since then, the country has experienced several electoral cycles, often tainted by violence and electoral fraud. In 1990, for what is generally considered as the first really democratic election, Jean-Bertrand Aristide was elected president but was overthrown by a military coup in 1991. He returned to power in 1994 but was again ousted in 2004.

Our analysis thus begins in 2004. Since then, the country has organized several elections, often contested and marred by violence and electoral fraud. The presidential and legislative elections of 2006 in Haiti were held after the fall of President Jean-Bertrand Aristide in February 2004, following violent protests. The electoral campaign lasted four months for the first round and two months for the second round (from October 2005 to February 2006 and from February 2006 to April 2006). The campaign was marked by active participation of candidates and their supporters, as well as violent clashes between different political factions. Candidates organized public gatherings and televised debates to promote their programs and convince voters to support them.

The sequence of elections in Haiti in 2009-2011 included elections to elect one-third of the Senate (2009), a new president (first round in 2010, second round in 2011), and all members of the Chamber of Deputies (2010). They were organized by the Provisional Electoral Council (CEP) with the support of the United Nations Stabilization Mission in Haiti (MINUSTAH) and the international community. The electoral campaign for the 2009 senatorial elections in Haiti was tense and marked by accusations of bias and favoritism toward certain candidates by the Provisional Electoral Council (CEP). It was characterized by bouts of violence, intimidation, and threats against certain candidates and their supporters.

The 2010 presidential elections were delayed by the devastating earthquake that struck the country in January of that year. The Haitian general elections of 2010-2011 took place on November 28, 2010 (first round) and March 20, 2011 (second round). The elections included presidential, legislative, and senatorial contests. The electoral process was marked by delays, accusations of fraud, and violence. The electoral campaign was tense, with more than 50 candidates vying for the presidency. The electoral process was complicated by disagreements on how the elections should be organized, as well as logistical and financial issues.

The Haitian general elections of 2015-2016 were held for the first round on August 9, 2015, and for the second round on October 25, 2015, to elect 119 deputies to the Chamber and 20 of the 30 Senate seats. These elections were canceled due to massive fraud. An additional second round took place on November 20, 2016, for six senatorial seats and 24 deputy seats. These elections were also marked by numerous delays, accusations of fraud, and violent protests. The electoral campaign was highly animated, with over 50 candidates running for the presidency. Again, the electoral process was complicated by disagreements on how the elections should be organized, as well as logistical and financial issues.

The elections finally took place in October 2015 but were marred by accusations of fraud and manipulation. The results were contested by several candidates, leading to protests and violence in the streets of the capital, Port-au-Prince, and other cities in the country.

The report of the Electoral Observation Mission of the Organization of American States (2017) provides a view on the electoral context. It for example states that, during the pre-election phase: "The violence and fear that it would spread on election day, as well as the improper use of public funds in favor of one party, marked this stage in the process." On election day, the observers note that 16 candidates from 10 parties have "resorted to violence or attempted to derail the process" and that 17 political parties "had committed acts of violence."

In January 2016, an agreement was reached among major candidates to organize new presidential elections, which took place in November 2016. Candidate Jovenel Moïse won the election and became president of Haiti in February 2017. The electoral campaign was marked by high citizen participation, numerous candidates, and increasing political and social tensions in the country. In 2016, Jovenel Moïse was elected president, but his tenure was tainted by allegations of corruption and mismanagement. During his mandate, legislative, local, and municipal elections took place, to elect members of the National Assembly, mayors, and municipal councilors. Legislative elections were held in October 2016, while municipal and local elections took place in January 2017. These elections were organized by the Senate, and then by the newly installed government, after several years of delays, postponements, and electoral disputes.

In 2020, protests erupted to demand the resignation of Jovenel Moïse, but he refused to step down and was assassinated in July 2021. Since then, the country has been plunged into a political crisis, and the holding of elections is uncertain.

3 Data and Methodology

3.1 Data sources and definitions

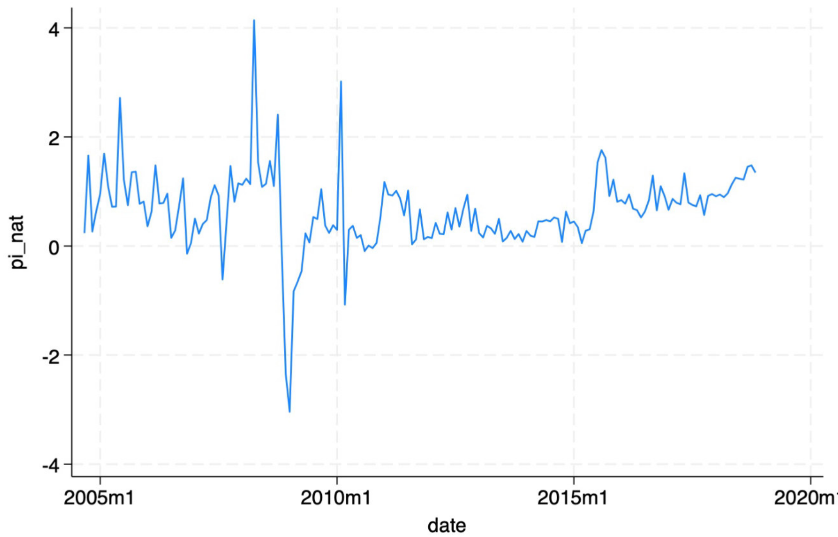
This paper uses monthly data from the Haitian Institute of Statistics and Informatics (IHSI) covering the period from August 2004 to June 2018. The IHSI computes and publishes three inflation series derived from its nationwide Consumer Price Index (CPI) survey: national CPI inflation, π_{nat}, inflation for locally produced goods, π_{local}, inflation for imported goods, π_{import}.

These series are directly provided by the IHSI and are based on a harmonized, nationally representative monthly price survey covering major consumption categories (food, housing, transportation, energy, communication, health, etc.). Figures 1–3 document the behavior of the three inflation rates over the sample period.

To isolate the effect of political disturbances, we include a set of exogenous variables capturing major natural, sanitary, and external shocks that can affect prices independently of political events. These *shock controls* are: the 2010 earthquake (binary, three-month window), the 2015 drought (binary, April–July 2015), the 2015–2016 ban on 23 Dominican imports (binary), and the monthly exchange-rate depreciation ($\Delta \ln$ Gourde/USD).[2] The variables

[2]From a descriptive point of view, exchange-rate movements do not appear to be systematically aligned with elections. Figures 4 and 5 in the Appendix C.1 overlays monthly nominal depreciation on election windows. Large depreciations mainly occur outside electoral periods, and average depreciation in a $\pm$6-month window around elections is very close to the sample mean. Our election indicators are thus not proxying for exchange-rate crises.

Figure 1: National inflation, monthly, 2004–2018



definitions are explained in Table 1.

3.2 Descriptive Statistics

Table 2 reports descriptive statistics for the three inflation series and for the exchange rate. National inflation averages 0.65% per month, with substantial month-to-month volatility (standard deviation of 0.72). Local inflation displays a similar mean and a slightly lower standard deviation. Imported inflation has the lowest mean (0.62%) but exhibits markedly higher volatility (standard deviation of 1.09) and the widest range of monthly changes.

The three inflation indices move closely together around major shocks, reflecting the fact that national inflation mechanically aggregates the local and imported components. At the same time, the two underlying series display distinct short-run sensitivities. Imported inflation responds more sharply to external conditions - most notably to exchange-rate fluctuations. As shown in Figure 4, the pronounced depreciations of the gourde in 2015–2016 coincide with clear spikes in π_{import}. Local inflation, in contrast, tends to vary with domestic supply conditions and internal distribution frictions. Because national inflation combines both components, it displays smoother month-to-month movements while still capturing the influence of both domestic and

Table 1: Control variables

Variables	**Type/modalities**	**Definition**
Earthquake	Yes = 1, No = 0	Binary variable that takes the value 1 between one and three months after the 2010 earthquake and 0 otherwise. The earthquake occurred in January 2010.
Drought	Yes = 1, No = 0	Binary variable that takes the value 1 between one and three months after the 2015 drought and 0 otherwise. The drought occurred between April and July 2015.
Exchange rate	Continuous	Monthly variation in the exchange rate.
Boycott	Yes = 1, No = 0	Binary variable that takes the value 1 during the interval between the enactment and the abolition of the Haitian government's decision to ban the transportation and sale of 23 Dominican products in Haiti in 2015, and 0 otherwise. This ban took place between September 2015 and September 2016.

Figure 2: Local inflation, 2004–2018

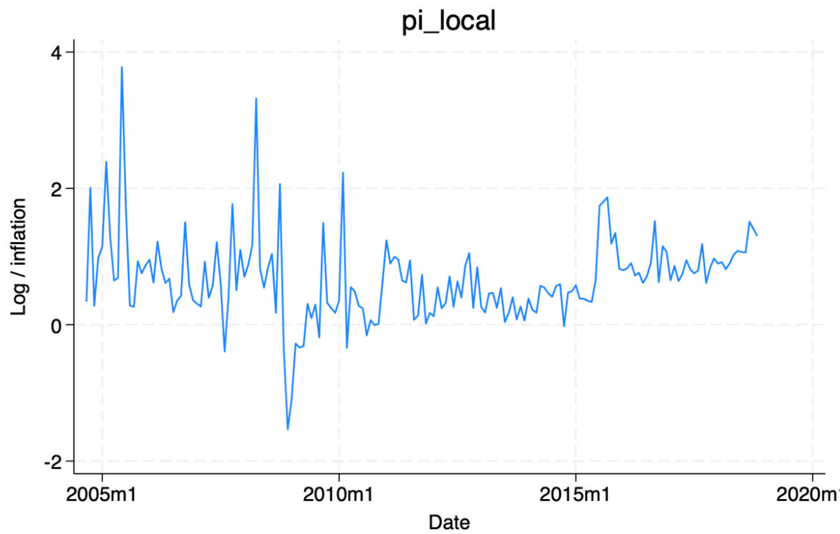


Table 2: Descriptive statistics for inflation and exchange rate series

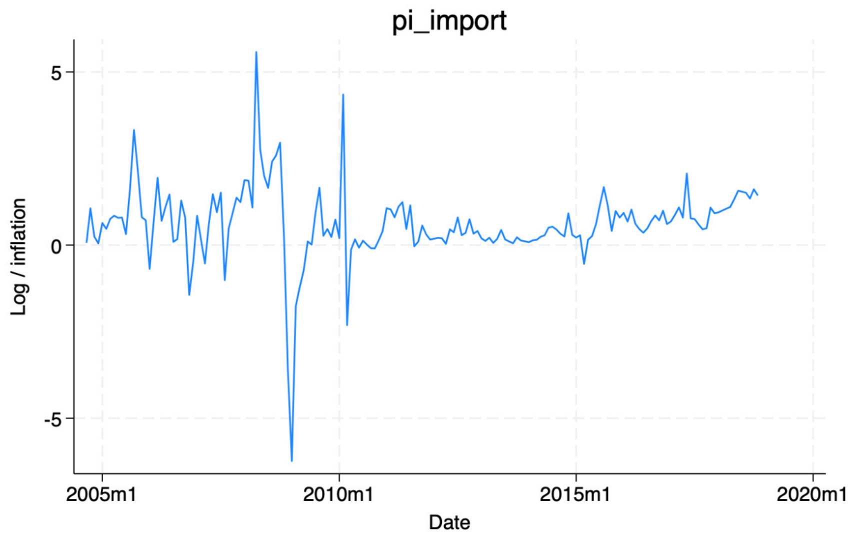
	Mean	SD	Min	Max	Obs
National inflation (π_{nat})	0.6510	0.7227	-3.0396	4.1400	171
Local inflation (π_{local})	0.6701	0.6322	-1.5347	3.7785	171
Imported inflation (π_{import})	0.6171	1.0918	-6.2408	5.5760	171
Exchange rate depreciation ($\Delta \ln FX$)	0.1847	0.8502	-4.7033	5.5362	166

external shocks.

A key modeling decision is to concentrate the baseline analysis on national CPI inflation rather than on local and imported subcomponents. This choice is justified along three dimensions.

First, from an economic perspective, national inflation is the conceptually relevant measure for household welfare, macroeconomic stability, and policy credibility. Political unrest can influence both domestic production conditions and external financial pressures, and the national CPI is the appropriate aggregate to capture these channels jointly. Second, from a data perspective, the three inflation series are mechanically linked and move closely together. National inflation is simply the weighted aggregation of local and imported components, and descriptive evidence shows that the main shocks - political

Figure 3: Imported inflation, 2004–2018



or otherwise - are reflected in all of them. Working with the aggregate therefore avoids treating the components as independent phenomena. Third, from an econometric perspective, imported inflation displays larger short-run fluctuations due to exchange-rate pass-through and external price movements. These high-frequency shocks can react to political uncertainty, but they also reflect global conditions unrelated to domestic events. Using headline CPI reduces this idiosyncratic variability, yielding a more stable series for identifying the inflationary impact of political disturbances.

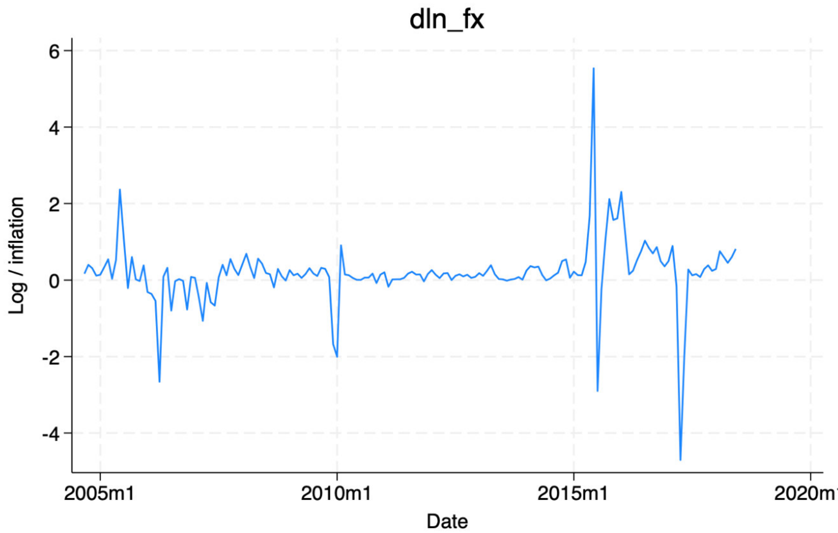
3.3 Methodology

This section presents the empirical strategy used to quantify the dynamic effects of political disturbances on inflation in Haiti. We proceed in three steps.

First, we estimate linear autoregressive models of monthly inflation, augmented with political-event indicators and a set of macroeconomic and environmental shocks. These regressions provide baseline estimates of the contemporaneous and near-term correlations between political events and inflation, while controlling for inflation persistence and external cost factors.

Second, we implement event-study specifications that trace the average path

Figure 4: Exchange rate (Gourde/USD), 2004–2018



of inflation before and after each political episode. We estimate both standard event-studies centered on individual events and a "stacked" event-study pooling all identified episodes into a single framework, following the recent literature on staggered or irregularly timed shocks ([Cengiz et al., 2019].

Third, we estimate impulse responses using local projections [Jordà, 2005], which allow us to analyze medium-run inflation dynamics without imposing strong parametric restrictions on the underlying data-generating process. These three approaches are complementary: linear regressions capture contemporaneous correlations, event-studies highlight pre-trends and short-run deviations, and local projections quantify persistent effects.

3.4 Equations and baseline specification

A potential concern is that our specifications may omit relevant macroeconomic controls, in particular a measure of real activity or the output gap. Monthly indicators of real activity are not available for Haiti over our sample, and annual GDP cannot be meaningfully combined with monthly inflation in a short time-series context. We therefore adopt a different strategy.

First, the baseline model already accounts for the main determinants of short-run inflation in a small open economy: lagged inflation captures intrinsic persistence, nominal exchange-rate depreciation controls for imported inflation, and we include explicit dummies for natural disasters, riots and fuel price shocks. Diagnostic results (provided in Appendix A) show no evidence of cointegration between CPI components, indicating that inflation can be modeled in a purely short-run autoregressive framework.

Second, we augment the specification with month-of-year fixed effects to capture systematic seasonal variation in inflation that may correlate with unobserved movements in real activity. Concretely, our initial baseline linear regressions take the form:

$$\pi_t = \alpha + \rho\pi_{t-1} + \beta\,\mathrm{Event}_t + \gamma' X_t + \sum_{m=2}^{12} \mu_m D_{m,t} + \varepsilon_t,$$

where Event_t is the electoral-disturbance indicator, X_t collects exchange-rate depreciation and the various shock controls, and $D_{m,t}$ denotes a dummy equal to one when month t falls in calendar month m. January is omitted as the reference month to avoid perfect multicollinearity. The month-of-year fixed effects are included purely as reduced-form controls for recurrent seasonal patterns, not to separately identify demand-side versus supply-side inflation channels. These dummies absorb stable seasonal patterns in demand and supply conditions - such as harvest cycles, school-fee payments, religious holidays or seasonal labor demand - that could otherwise be confounded with political dynamics.

Still, the inclusion of a relatively large set of controls - including seasonal fixed effects and multiple shock indicators - raises the concern that overfitting could attenuate or mask any electoral effect. To assess this possibility, we conduct several diagnostic exercises. First, we used an information criteria (AIC/BIC): comparing a parsimonious AR(1) specification to an augmented version with month dummies shows that both AIC and BIC strongly favor the simpler model. Adding month-of-year fixed effects increases the AIC from 163 to 179 and the BIC from 168 to 214, indicating over-parameterization without explanatory gains. Second, a joint F-test fails to reject the null that all 11 month dummies equal zero (F(11,93)=1.66, $p = 0.095$). Seasonal effects are small and statistically negligible. Third, we used a Frisch–Waugh–Lovell decomposition: residualizing inflation with respect to month dummies shows that seasonality explains only 5% of the variance ($R^2 = 0.049$). When the residual series is regressed on political-event indicators and shock controls, estimated coefficients remain essentially unchanged. Fourth, pairwise correlations between month-of-year dummies and political-event dummies are uniformly small (absolute values < 0.17), confirming that the two sets of indicators are empirically orthogonal. Taken together, these results lead us to retain the parsimonious specification as the baseline.

Finally, given the small open nature of the Haitian economy, we include monthly changes in the nominal exchange rate to control for external cost shocks. We use only the contemporaneous depreciation term rather than a distributed-lag structure for two reasons. First, the short sample (fewer than 200 monthly observations) makes specifications with many lagged regressors impractical. Second, the contemporaneous depreciation term is never statistically significant, and adding one or two lags of $\Delta \ln FX_t$ does not materially affect the estimated election coefficients; the FX terms remain jointly insignificant (see Appendix Table C.2). A parsimonious specification with a single contemporaneous FX regressor is therefore preferred.

3.5 Event-study Specifications

To analyze the dynamic behavior of inflation around political disturbances, we estimate standard event-study regressions centered on the onset month of each episode. Election sequences differ in duration across events, but the event-study design aligns them by their starting date. For event k beginning at time T_k, we construct a set of relative-time indicators:

$$D_{h,t}^{(k)} = \mathbf{1}\{t - T_k = h\}, \qquad h \in [-H^-, H^+],$$

where H^- and H^+ define the pre- and post-event windows used for all events. Inflation is then regressed on these dummies, together with lagged inflation and shock controls:

$$\pi_t = \alpha + \rho\, \pi_{t-1} + \sum_{h=-H^-}^{H^+} \theta_h D_{h,t}^{(k)} + \gamma' X_t + \varepsilon_t.$$

This specification traces the average deviation of inflation at each horizon h relative to the beginning of the political episode, allowing us to detect pre-trends and short-run inflation spikes.

3.6 Stacked Event-study

Electoral sequences in Haiti occur irregularly over time, and their number is small over the period. To maximize statistical power and obtain a unified estimate of the average dynamic response, we adopt a "stacked" event-study approach. All event windows are re-centered on their respective event dates and stacked into a single pseudo-panel:

$$\pi_{i,h} = \alpha + \rho\, \pi_{i,h-1} + \sum_{h=-H^-}^{H^+} \theta_h\, \mathbf{1}\{h\} + \gamma' X_{i,h} + \varepsilon_{i,h},$$

where i indexes events and h denotes relative time. This method, now standard in the treatment–effects literature, improves precision and avoids mechanically weighting large events more heavily than smaller ones.

3.7 Local Projections

Finally, we estimate impulse-response functions using local projections [Jordà, 2005]. For each horizon $h = 0, 1, \dots, H$, we estimate

$$\pi_{t+h} = \alpha_h + \rho_h \pi_{t-1} + \beta_h\, \text{Event}_t + \gamma_h' X_t + \varepsilon_{t,h},$$

where Event_t captures the occurrence of a political episode and X_t denotes the vector of control variables. The sequence $\{\beta_h\}_{h=0}^{H}$ traces the dynamic adjustment of inflation following a political disturbance. Local projections are robust to misspecification of the underlying autoregressive structure and are well suited to short time series with structural breaks or irregular event timing.

In combination, the linear regressions, event-studies, stacked event-studies and local projections provide a comprehensive picture of how political instability affects inflation dynamics in Haiti. The next section presents the empirical results.

But, before that, let us establish that our empirical strategy is reduced-form: the estimated responses should be interpreted as composite effects of multiple political and economic mechanisms rather than as structurally identified causal parameters. The event-study and stacked specifications assume that, conditional on lagged inflation and observed controls, no other shocks are systematically aligned with election dates. Local projections provide a flexible framework for tracing dynamic responses without imposing a full structural model, and they remain valid under broad forms of misspecification ([Montiel Olea and Plagborg-Møller, 2021]). However, LPs may also exhibit finite-sample distortions due to serial correlation, smoothing bias, and time-variation in the underlying data-generating process ([Kilian and Kim, 2011]; [Li et al., 2024]). For these reasons, we focus on the qualitative timing and shape of the responses rather than precise point estimates, and we complement our analysis with a placebo-based randomization test (detailed in Appendix B).[3]

4 Empirical Results

4.1 Linear regressions

Tables 3 and 4 report the baseline autoregressive regressions of monthly national inflation on political-event indicators and macroeconomic controls. Across all specifications, inflation exhibits substantial persistence, with autoregressive coefficients ranging from 0.25 to 0.70 depending on the lag structure. This confirms the well-documented inertia of consumer prices in small open economies.

Turning to electoral effects, the estimates reveal no statistically significant increase in inflation during election months. Across nine specifications, the coefficient on the contemporaneous election-month dummy is consistently small (between –0.01 and +0.14 percentage points) and never statistically distinguishable from zero. This result holds whether the model includes one or two lags of inflation, with or without additional controls.

In contrast, inflation is markedly higher during the post-2015 period: the dummy for the structural break after July 2015 enters with a positive and robust coefficient in all specifications (between +0.21 and +0.29 percentage points, significant at the 1% level). This shift reflects the deterioration of macroeconomic conditions after 2015, characterized by exchange-rate pressures, political instability, and disruptions in supply chains (see Figure 1).

Table 4 decomposes the political sequence into individual electoral episodes. The results display substantial heterogeneity across episodes. The 2006

[3] Recent theoretical work further shows that LPs possess a double-robustness property and outperform VAR-based inference under many forms of misspecification ([Olea-Montiel et al., 2024]).

election is associated with a strong and significant increase in inflation (between +0.44 and +0.55 p.p.). The 2009 and 2017 episodes appear mildly deflationary, although significance is not systematic. The 2010–11 sequence yields positive and significant coefficients, while the 2015–16 elections are associated with sizeable and negative effects in some specifications.

It is worth noting that electoral episodes differ considerably in duration, ranging from short three-month sequences to extended multi-month periods. Longer episodes may mechanically dilute the within-episode variation captured by the election dummy, while shorter episodes produce more concentrated shocks. Such heterogeneity in duration may therefore contribute to the non-significant or contrasting coefficients observed across episodes.

Regarding macroeconomic shocks, none of the disaster or riot indicators is robustly significant, and the contemporaneous exchange-rate depreciation does not materially affect inflation coefficients. This confirms that the baseline specification successfully isolates the political component of inflation dynamics.

Taken together, the linear regressions reveal slight evidence of an immediate impact of elections on the dynamics of the inflation rate in Haiti, once persistence and macroeconomic shocks are controlled for. This motivates a more flexible analysis of inflation dynamics around political events, which we implement next using event-study methods and local projections.

Table 3: Baseline regressions: inflation components on election month and controls

	π_{nat} (national CPI inflation)			π_{local} (local CPI inflation)			π_{import} (imported CPI inflation)		
	(1) AR(1)	(2) AR(1)+Break	(3) AR(2)+Break	(4) AR(1)	(5) AR(1)+Break	(6) AR(2)+Break	(7) AR(1)	(8) AR(1)+Break	(9) AR(2)+Break
$L1.\pi_{nat}$	0.452***	0.423***	0.386**	0.303***	0.266***	0.247***	0.699***	0.683**	0.613*
	(0.130)	(0.137)	(0.159)	(0.074)	(0.071)	(0.086)	(0.266)	(0.283)	(0.317)
$L2.\pi_{nat}$			0.101			0.061			0.176
			(0.092)			(0.068)			(0.154)
Election month (t)	0.093	0.050	0.089	0.137	0.082	0.110	0.013	-0.010	0.049
	(0.119)	(0.102)	(0.092)	(0.154)	(0.132)	(0.123)	(0.101)	(0.094)	(0.085)
Post-2015m7 (break)		0.227***	0.209***		0.289***	0.283***		0.122	0.084
		(0.084)	(0.073)		(0.095)	(0.086)		(0.146)	(0.133)
Constant	0.336***	0.311***	0.261***	0.445***	0.412***	0.374***	0.154	0.141	0.071
	(0.112)	(0.114)	(0.082)	(0.079)	(0.083)	(0.071)	(0.190)	(0.186)	(0.129)
Observations	165	165	164	165	165	164	165	165	164
R^2	0.205	0.220	0.237	0.124	0.157	0.174	0.211	0.213	0.225

Notes: Columns (1)–(3), (4)–(6), and (7)–(9) use national, local, and imported CPI inflation as the dependent variable. Specifications differ by the dynamic structure (AR(1) vs. AR(2)) and by inclusion of a post-2015m7 break dummy. All dynamic controls are lags of national inflation. All regressions include shock controls and nominal exchange-rate depreciation. HAC standard errors (Bartlett kernel, 12 lags) are reported in parentheses. *, **, *** denote significance at the 10%, 5%, and 1% levels.

Table 4: Electoral episodes and inflation components

	Dependent variable		
	π_{nat}	π_{local}	π_{import}
	(1)	(2)	(3)
$L1.\,\pi_{nat}$	0.400***	0.252***	0.646**
	(0.153)	(0.078)	(0.304)
D2006	0.478***	0.439***	0.548**
	(0.146)	(0.137)	(0.213)
D2009	-0.261	-0.334***	-0.142
	(0.163)	(0.108)	(0.285)
D2010–11	0.208**	0.199**	0.223**
	(0.086)	(0.098)	(0.104)
D2015–16	-0.459**	-0.312	-0.729***
	(0.185)	(0.242)	(0.182)
D2017	-0.212*	-0.153	-0.331***
	(0.114)	(0.139)	(0.115)
shock_disaster	0.136	-0.228	0.712
	(0.275)	(0.205)	(0.467)
shock_riot	0.135	0.012	0.347**
	(0.164)	(0.236)	(0.151)
shock_fuel	-0.391	-0.517	-0.172
	(0.381)	(0.465)	(0.404)
Nominal depreciation (%)	0.123	0.145	0.082
	(0.096)	(0.120)	(0.097)
Post-2015m7	0.391***	0.410***	0.365*
	(0.136)	(0.146)	(0.189)
Constant	0.302**	0.412***	0.117
	(0.118)	(0.083)	(0.198)
Observations	170	170	170
R^2	0.264	0.214	0.246

Notes: Columns (1)–(3) report regressions of national, local, and imported CPI inflation, respectively, on electoral-sequence indicators. All specifications include an AR(1) term in national inflation ($L1.\,\pi_{nat}$), domestic shock controls (natural disasters, riots, fuel-price shocks), contemporaneous nominal exchange-rate depreciation, and a post-2015m7 structural-break dummy. HAC standard errors (Bartlett kernel, 12 lags) are reported in parentheses. *, **, *** denote significance at the 10%, 5%, and 1% levels.

4.2 Event-study evidence

4.2.1 Single-event specifications

We now turn to the dynamic evolution of inflation around each electoral episode. Unlike the linear regressions, which impose a common contemporaneous effect across heterogeneous episodes, the event-study framework allows us to visualize inflation trajectories before and after each election. Because each episode provides only a small number of observations at each relative-time horizon, single-event event studies cannot support meaningful confidence intervals; the figures should therefore be interpreted as descriptive patterns rather than formal statistical estimates. This approach nonetheless reveals dynamics that are not captured by the linear specification.

Figures 5–9 display the evolution of monthly inflation from several months before the election to several months after, with the election month normalized to $t = 0$. Because each figure corresponds to a single electoral episode observed over a short time window, we report point estimates only. Confidence intervals are omitted, as standard error estimates are not informative in a single-event setting.

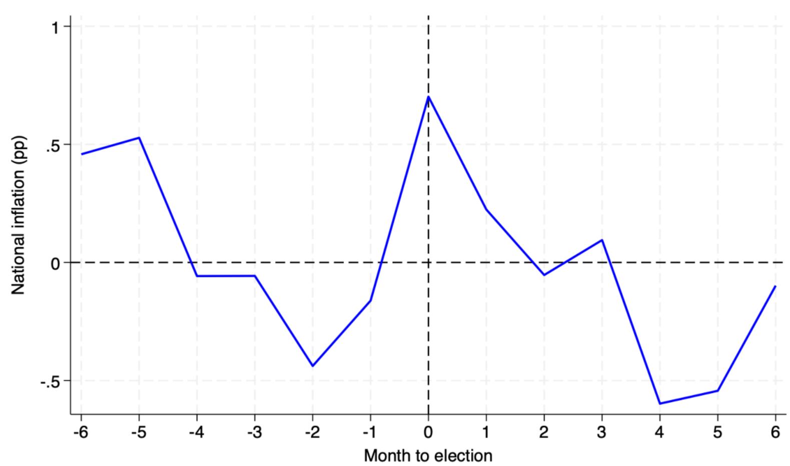


Figure 5: Event-study for the 2006 electoral sequence

Several episodes show clear short-run inflation spikes around the election date. The 2006 and 2010–2011 sequences exhibit pronounced increases in inflation during the months immediately surrounding $t = 0$, with visible

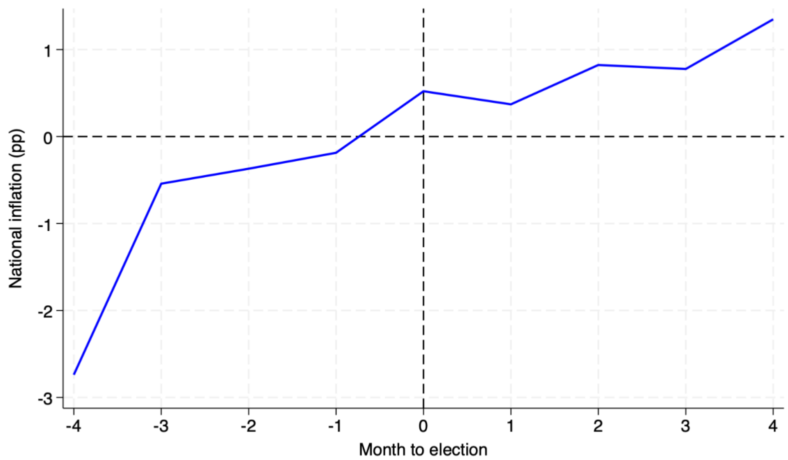


Figure 6: Event-study for the 2009 electoral sequence

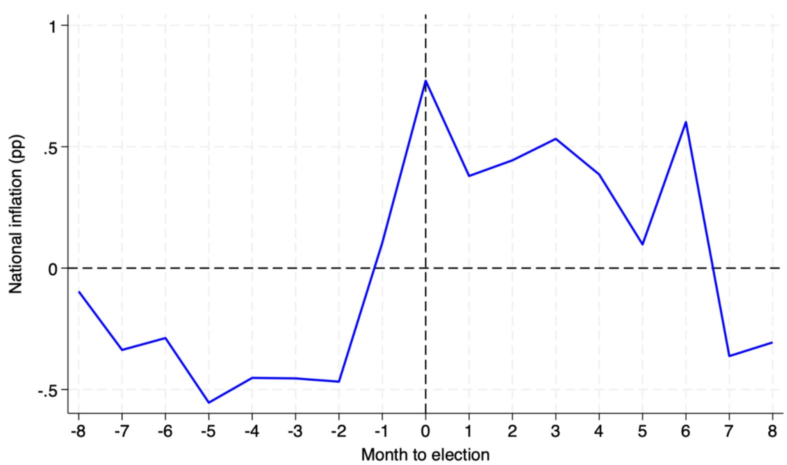


Figure 7: Event-study for the 2010–2011 electoral sequence

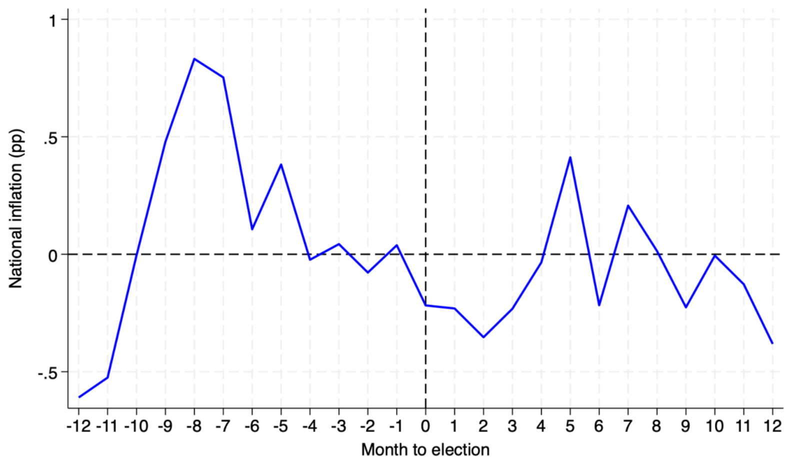


Figure 8: Event-study for the 2015–2016 electoral sequence

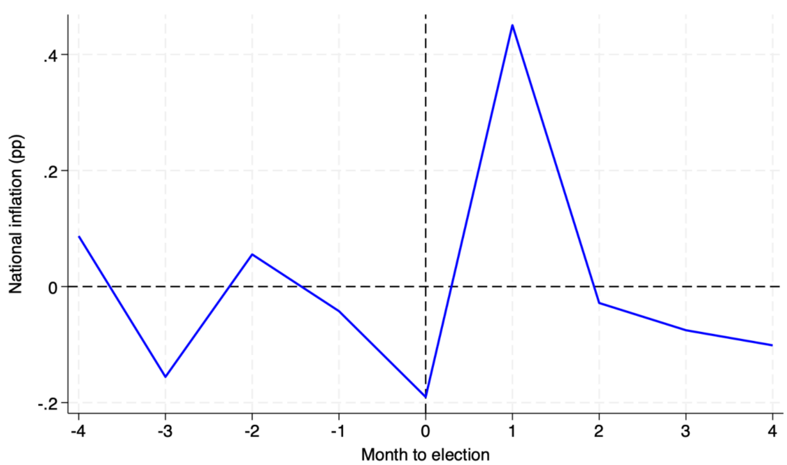


Figure 9: Event-study for the 2017 electoral sequence

upward shifts between $t = -1$ and $t = 1$. These patterns are consistent with pre-electoral spending pressures or short-run demand surges associated with electoral mobilization. The 2017 episode displays a similar but more muted pattern: inflation remains stable during the pre-election months but rises sharply in the month following the vote, again compatible with transitory demand or liquidity injections.

While some episodes are dominated by broader macroeconomic shocks - most notably 2009 and 2015–2016 - this heterogeneity is itself informative. It highlights the limitations of imposing a single average effect in a linear regression, and underscores the usefulness of the event-study approach for detecting short-run election-related movements even in the presence of background volatility.

Overall, the individual event-studies suggest that several electoral episodes are associated with short-lived increases in inflation, concentrated in a narrow window around the election month. These dynamic patterns would be those expected under a mechanism of electoral spending or vote-buying incentives, and they provide richer information than the baseline linear regressions.

4.2.2 Stacked event-study

To extract the average dynamic response across episodes, we pool all electoral sequences into a stacked event-study (Figure 10). Despite the heterogeneity observed in individual episodes, the pooled pattern is remarkably stable.

Inflation is mildly increasing in the months preceding the election, with point estimates rising from approximately 0.4 to nearly 0.9 percentage points between $t = -4$ and $t = -1$. This pre-election buildup is consistent with anticipatory price adjustments or increases in demand associated with the electoral cycle. Around the election month itself, inflation remains elevated relative to earlier months, and although confidence intervals are wide due to sample size, point estimates do not exhibit a decline at $t = 0$ or immediately afterward.

Following the election, inflation gradually returns toward its pre-event range, consistent with a transitory inflationary impulse rather than a persistent macroeconomic shock. The absence of a large post-election drop is also compatible with short-lived, demand-driven pressures such as public transfers or politically motivated spending.

Taken together, the stacked results suggest the presence of an economically

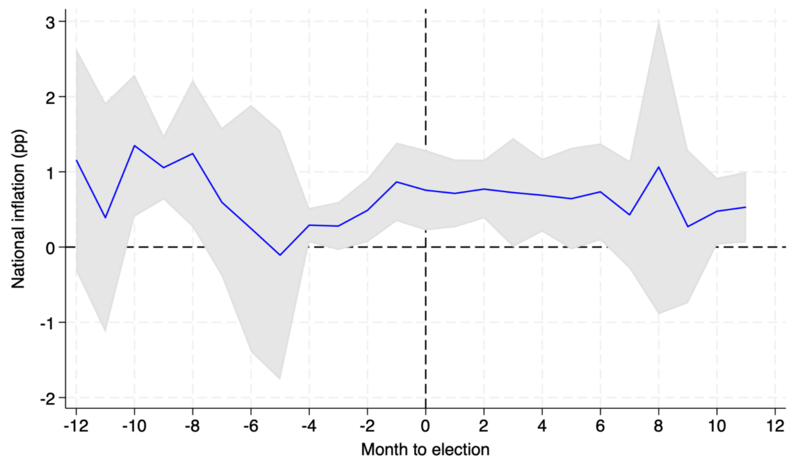


Figure 10: Stacked event-study pooling five electoral sequences

meaningful electoral pattern: inflation tends to rise in the months leading up to elections and stabilizes afterward. This average pattern aligns with the short-run increases observed in several individual sequences and provides complementary evidence to the linear regressions, which by construction cannot capture these localized but recurrent dynamics.[4]

4.3 Local projection impulse responses

To complement the event-study analysis, we estimate impulse-response functions using [Jordà, 2005] local projections. This approach traces the dynamic response of inflation to the onset of an electoral episode without imposing a specific autoregressive structure. Figure 11 displays the estimated response of monthly inflation from three months before the election to six months afterward.

The local projections reveal a smooth and intuitive dynamic pattern. Inflation shows a gradual upward movement in the months preceding the election: the estimated response increases from approximately –0.4 percentage points at $t = -3$ to values close to zero just before the election. This pre-election buildup is consistent with anticipatory pressures or politically motivated

[4]Robustness checks using local and imported inflation (see Figure 6 in Appendix C.3) show similar short-lived inflation spikes around elections, confirming that the headline results are not driven by a specific CPI subcomponent.

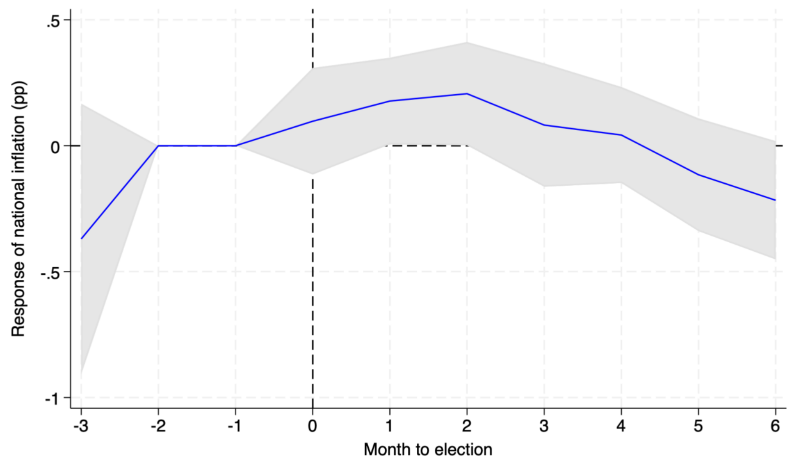


Figure 11: Local projection response of national inflation to an election shock (horizons $h = -3, \ldots, 6$)

spending, a mechanism that the linear regressions are not well suited to detect.

Around the election month itself ($t = 0$ to $t = 2$), inflation reaches its maximum response, with point estimates turning positive and peaking between $t = 1$ and $t = 2$. Although confidence intervals are wide due to the small sample ([Brugnolini, 2018],[Bruns and Lütkepohl, 2022], [Herbst and Johannsen, 2024]), and include zero at all horizons, the shape of the response mirrors the patterns observed in the event-study plots: a short-lived inflationary impulse concentrated in a narrow window around the election date.

Following this peak, the response gradually declines and converges back toward zero by $t = 5$ or $t = 6$, indicating that the electoral impulse is transitory and does not generate persistent inflation. This temporal profile is fully consistent with the interpretation of elections triggering short-run price pressures - through demand surges, liquidity injections, or pre-electoral transfers - rather than long-lasting macroeconomic imbalances.

Overall, the local projection results reinforce the evidence from the event-studies: several electoral episodes are associated with temporary increases in inflation, peaking around the election month and dissipating shortly afterward. These dynamics are precisely the type of short-run response that cannot be captured by a single contemporaneous coefficient in a linear regression, and they provide additional support for the hypothesis of short-lived, election-related spending pressures.

4.4 Discussion

A natural question is why the linear autoregressive regressions seem to provide weaker evidence of electoral effects than the event studies and the local projections. The explanation lies in what each method is designed to capture. In the first set of AR specifications (Table 3), a single election-month dummy aggregates all electoral episodes into one average contemporaneous effect. This construction mechanically averages together heterogeneous events.

The specifications with episode-specific dummies (Table 4) go one step further by allowing each electoral sequence to have its own coefficient. These regressions indeed reveal that some episodes (notably 2006 and 2010–11) are associated with significant inflation increases, while others are more neutral. Yet the AR framework still attributes the entire effect of each episode to a single month $t = 0$ and cannot capture the short-run build-up and subsequent decay documented in the dynamic analyses. The event studies and the local projections, by contrast, allow inflation to respond at horizons $h = -3$ to $h = +6$ around the election date. They show that the typical response peaks at $h = 1,2$ and then gradually returns to baseline, rather than being confined to the election month itself.

In this sense, the three approaches are not contradictory but complementary (see Table 5). The AR regressions highlight that there is no large, systematic jump in inflation exactly in the election month once persistence and shocks are controlled for, while the event studies and local projections show that elections are associated with short-lived inflation spikes in the broader window surrounding the vote. The dynamic methods are therefore better suited to detecting the kind of transitory, tightly timed responses that arise in a small sample with heterogeneous electoral episodes.

To gauge economic significance, we convert the local-projection estimates into a cumulative price-level effect. Focusing on the first three post-election months, where the inflation response is largest, the LPs imply an accumulated increase of approximately 0.54 percentage points in the consumer price index. Using the typical monthly salary in Haiti (28,000 HTG)[5], this corresponds

[5]According to recent labor-market statistics, the average monthly salary in Haiti is

Table 5: Summary of electoral effects across empirical methods

	Linear AR regressions	Event studies	Local projections
Treatment	Dummies	By-episode and pooled	Pooled shock
Timing of effect	$t = 0$	$h = -4$ to $h = +6$	$h = 0$ to $h = 6$
Main result	No systematic jump	Increases near elections	Increases post-election
Why different?	Averages across episodes, focuses on $t = 0$	Horizon-specific patterns	Dynamic responses

to about 454 HTG of additional household spending over the quarter surrounding an election. Although modest in percentage terms, this amount represents roughly 1.6% of a monthly salary, or the equivalent of several days of basic consumption for many households. The effect is therefore economically meaningful, representing 80% of an extra month of inflation.

Taken together, these results indicate that Haitian elections are associated with a short-lived but economically meaningful increase in monthly inflation. Because the number of electoral episodes is small, however, uncertainty remains about the statistical strength of these patterns. The next section therefore examines whether the detected post-election responses could arise from calendar noise or mechanical features of the data, using a placebo-based randomization test and other robustness checks.

5 Robustness tests and discussion

5.1 Randomization test

To assess whether the dynamic inflation response identified in the local projections could arise spuriously from calendar noise or recurring seasonal patterns, we implement a Fisher-style randomization inference (see, e.g., [Athey and Imbens, 2017], [Keele et al., 2012]). Our approach is also closely related to recent applications emphasizing the role of randomization tests in small samples and in the presence of complex heteroskedasticity [Young, 2019]. The procedure, detailed in Appendix B, proceeds as follows. First, election months are reassigned uniformly at random 500 times, subject to the restriction that no placebo election may fall within a ±6-month window of any true

about 28,000 HTG in 2025 (approximately USD 214).

election date. For each of these pseudo-election sequences, we re-estimate the symmetric local projection model.

As a summary measure of the post-election inflation response, we focus on the initial horizons where the baseline LPs display their strongest acceleration ($h = 1$ and $h = 2$). For each real or placebo sequence we compute the statistic

$$S = \frac{\beta(1) + \beta(2)}{2}.$$

The resulting placebo distribution of S_{plc} is centered around 0.073 with a standard deviation of 0.269. The true election sequence delivers $S_{\text{real}} = 0.223$, which lies well within the upper half of the placebo distribution but does not fall in its extreme tail. The implied randomization p-value is $p = 0.27$.

Figure 12 plots the distribution of placebo statistics together with the observed value from the actual election sequence.

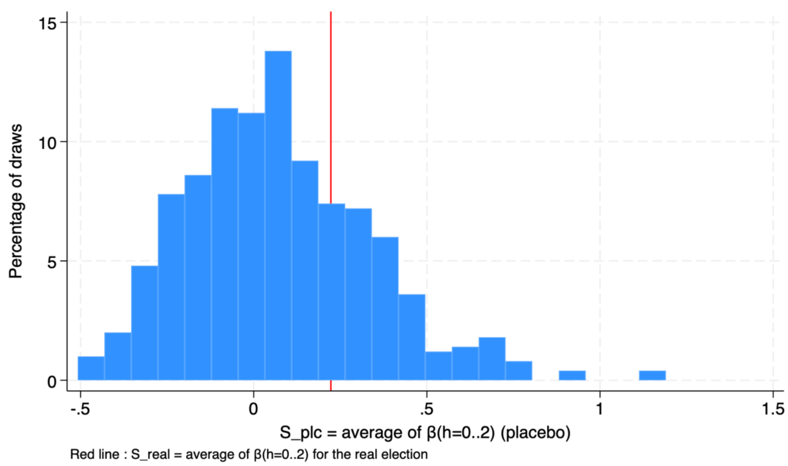


Figure 12: Randomization test for the effect of elections on inflation. The red line indicates the observed statistic S_{real}; the histogram shows the distribution of placebo statistics S_{plc} derived from 500 random reassigned election sequences.

Interpreting these results requires caution. Haiti's macroeconomic data contain only five electoral episodes over a sample of roughly 170 monthly observations. With such a small number of treated events, randomization inference is intrinsically low-powered, and the histogram confirms that placebo sequences can occasionally coincide with modest positive responses simply by chance. At the same time, the placebo distribution does not exhibit systematic or persistent inflationary patterns around randomly chosen dates, suggesting that the inflation dynamics uncovered in the baseline LPs do not arise from mechanical features of the data.

Taken together with (i) the individual event-study profiles showing systematic upward movements in inflation around the true electoral dates, (ii) the stacked event-study that smooths idiosyncratic election-specific variation, and (iii) the symmetric local projections revealing a short-lived post-election acceleration, the randomization exercise supports a coherent interpretation: Haitian elections are associated with a mild and temporary rise in inflation in the months immediately following the vote.

All in all, thus, across the three empirical approaches (baseline autoregressive regressions, individual and stacked event studies, and symmetric local projections), a consistent pattern emerges. Linear specifications, which impose a single parameter on all electoral episodes, tend to wash out heterogeneity and seldom deliver statistically significant coefficients. In contrast, the event studies make clear that inflation typically accelerates in the months surrounding an election, although with notable variation in timing and intensity across episodes. The stacked event study smooths this idiosyncratic variation and reveals a coherent upward drift in inflation before the vote and a modest decline thereafter. The local projections further confirm that inflation rises in the one to two months following an election and gradually returns to trend.

5.2 Further checks

As a further check, Table 5 in Appendix C.1 shows that adding a lagged depreciation term does not alter the estimated election effects: the FX coefficients remain small and statistically insignificant, and the magnitude and significance of the election coefficients are virtually unchanged.

A natural concern with only five elections is that the average local-projection response could be driven by a single episode. We therefore complement the baseline LPs with two robustness checks reported in Appendix C.4.1 and C.4.2. First, leave-one-out local projections show that removing any individual election sequence leaves the overall shape of the impulse response essentially unchanged: in all cases, inflation rises around $h = 1{,}2$ and gradually returns to baseline thereafter. Second, event-specific impulse responses - estimated separately for each electoral sequence - display some heterogeneity in magnitude, as expected given the different macroeconomic environments, but broadly share the same qualitative pattern of short-lived inflation spikes around the vote. Together, these robustness checks confirm that the average LP profile is not driven by a single outlying election and that the short-run inflationary impulse is a recurrent feature across episodes.

5.3 Interpretation and Mechanisms

The impulse responses estimated in this paper should be interpreted as reduced-form, composite effects of several channels that may operate simultaneously around elections in Haiti. The dynamic inflationary pattern we document—a short-lived increase in the months surrounding the vote followed by a rapid return to baseline—is consistent with multiple mechanisms identified in qualitative accounts of Haitian elections.

First, a substantial body of evidence, if only anecdotal or from media accounts, describes Haitian elections as periods of intensive vote-buying ([Justesen and Manzetti, 2023]; [Woller et al., 2023]). Campaigns frequently involve the distribution of cash, food, and goods, which temporarily increases households' liquidity and raises demand for basic items. Such demand surges naturally translate into short-run price pressures, a mechanism also highlighted in [Aidt et al., 2020] for other low-income democracies. The short-lived inflation spikes at $h = 1{,}2$ are consistent with these temporary injections of purchasing power.

Second, election periods in Haiti are routinely associated with road blockages, strikes, and localized violence. Reports from the [Electoral Observation Mission of the Organization of American States, 2017] document widespread unrest and logistical disruptions in the days surrounding the vote. These disturbances can restrict the flow of goods to urban markets, especially fresh food and locally supplied items, generating transitory upward pressure on prices. The rapid post-election return to trend in the IRFs aligns with this interpretation of temporary supply bottlenecks.

Third, retailers and households may adjust prices preemptively when anticipating electoral unrest. Anecdotal evidence suggests that merchants sometimes raise prices in anticipation of interruptions to supply chains or in response to expected increases in demand. This mechanism is consistent with the mild pre-election drift observed in the stacked event-study and the local projections.

Fourth, while Haiti's exchange rate shows strong medium-run comovement with imported inflation, FX depreciation does not systematically coincide with election months. The appendix shows that the gourde does not depreciate abnormally around elections and that adding lagged FX terms does not materially alter the IRFs. Hence, exchange-rate pass-through is unlikely to be the primary driver of the observed election-cycle dynamics.

Fifth, because the statistical design does not isolate any one of these mechanisms, the estimated IRFs capture the net effect of all channels that typically operate during election cycles: demand surges from vote-buying, temporary supply constraints due to unrest, anticipatory pricing, and associated liquidity pressures. What the methods reveal is not the dominance of one mechanism but their coincidence in time and their shared inflationary footprint. The transitory nature of the IRFs, peaking shortly after the vote and converging to zero by $h = 5{,}6$, is consistent with the inherently temporary character of these disturbances.

6 Conclusion

This paper investigates whether elections generate systematic movements in inflation in Haiti, a fragile democracy where electoral cycles are frequently associated with violence, market disruptions, and widespread vote-buying. Using monthly CPI data over 2004–2018, we combine linear autoregressive regressions, individual and stacked event studies, and symmetric local projections to uncover the dynamic behavior of inflation around electoral periods. Three main findings emerge.

First, linear regressions impose a single contemporaneous effect across heterogeneous electoral episodes and therefore produce limited evidence of an electoral cycle. Once inflation persistence and macroeconomic shocks are controlled for, election-month dummies generally do not reach statistical significance.

Second, when we allow for richer dynamics, both the individual event studies and the stacked event-study profiles reveal consistent short-run patterns. Several elections display clear increases in inflation within a narrow window surrounding the vote, with stacked estimates showing a mild but systematic buildup in the months preceding elections and a stabilization shortly afterward.

Third, local projections corroborate this interpretation. The estimated impulse-response functions show that inflation tends to accelerate modestly in the one to two months after an election before gradually returning to baseline. The effect is economically meaningful - comparable in magnitude to other major domestic shocks - but transitory. A placebo-based randomization test confirms that such patterns rarely occur around randomly assigned dates, although the small number of electoral episodes in the sample limits the statistical power of randomization inference.

Taken together, these results indicate that elections in Haiti generate a short-lived but coherent inflationary impulse.[6] In a context marked by vote-buying, campaign-related transfers, and recurrent electoral unrest, these dynamics may reflect temporary liquidity injections, disruptions to production and distribution networks, or anticipatory behavior by market participants.

Finally, translating the estimated impulse-response functions into economic terms shows that elections generate a cumulative price-level increase of roughly 0.54 percentage points over the first three post-election months. For a typical monthly salary of 28,000 HTG in 2025, this amounts to about 454 HTG in additional household spending - roughly 1.6% of monthly income, or several days of basic consumption for many households, or 80% of an average month of inflation. Even though the effect is transitory, it is far from negligible in a country where most families have limited financial buffers.

These findings suggest that electoral instability imposes tangible welfare costs on vulnerable populations, reinforcing the need for institutional reforms that reduce the economic volatility associated with Haiti's democratic cycle.

Beyond their empirical contribution, the findings speak to broader debates on political accountability and economic stability in weak democracies. In settings where electoral competition relies heavily on direct spending and mobilization strategies, even modest liquidity injections or temporary supply frictions can generate measurable price increases, disproportionately affecting low-income households. Policies that strengthen electoral oversight, regulate campaign expenditure, and protect key distribution channels during political transitions could therefore mitigate short-run welfare losses associated with elections. Future work might examine the micro-level mechanisms behind these dynamics - disentangling household demand responses, market expectations, and supply constraints - and extend the analysis to other fragile democracies. Haiti illustrates how uncertainty around electoral processes can spill over into economic volatility, underscoring the need for institutional reforms that make democratic cycles less destabilizing for households. From a policy perspective, three broad priorities emerge. First, strengthening

[6] The results are also robust across CPI components: local and imported inflation exhibit similar short-run spikes around elections, despite their different economic determinants.

electoral governance is central. Enhancing the autonomy, transparency, and operational reliability of the Provisional Electoral Council (CEP) would reduce uncertainty around the electoral calendar and limit the mobilization-related disruptions that feed into market instability. Second, regulating campaign finance and curbing vote-buying practices could dampen the liquidity surges that contribute to short-run price pressures. Clearer disclosure rules, spending ceilings, and systematic monitoring of in-kind transfers - supported by domestic institutions and international partners - would reduce the magnitude of election-driven monetary shocks. Third, safeguarding distribution networks during electoral periods is essential. Measures to maintain transport corridors, ensure basic market security, and monitor key commodity prices can prevent localized bottlenecks that amplify temporary price increases. Overall, the results suggest that modest, well-targeted reforms in electoral management, campaign regulation, and market logistics could reduce the macroeconomic volatility associated with Haiti's democratic cycle. They also point to the importance of further research on the channels - demand surges, liquidity movements, and supply frictions - through which electoral dynamics translate into short-lived inflation spikes in fragile democracies.

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Appendix to "Are Tainted Elections Inflationary? Evidence from Haiti (2004–2018)"

Diagnostics and Additional Results

Appendix A. Diagnostic Tests for CPI, Inflation and Exchange-Rate Series

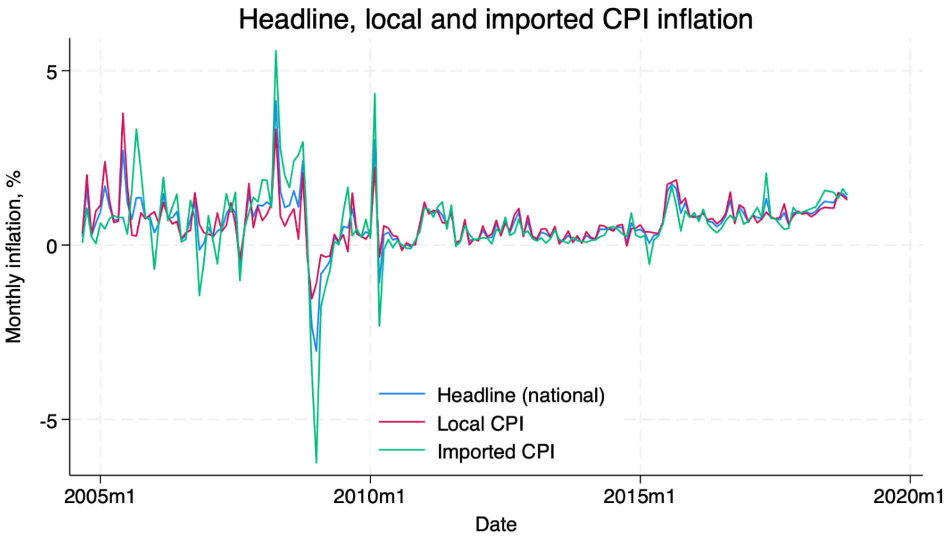
To avoid redundancy, diagnostics are presented only for the CPI aggregates, the main CPI sub-components included in the estimations, and the nominal exchange rate.

Figures 1–3 provide a graphical overview of the behaviour of the main inflation series and of the nominal exchange rate, while Tables 1–4 summarise the corresponding formal tests.

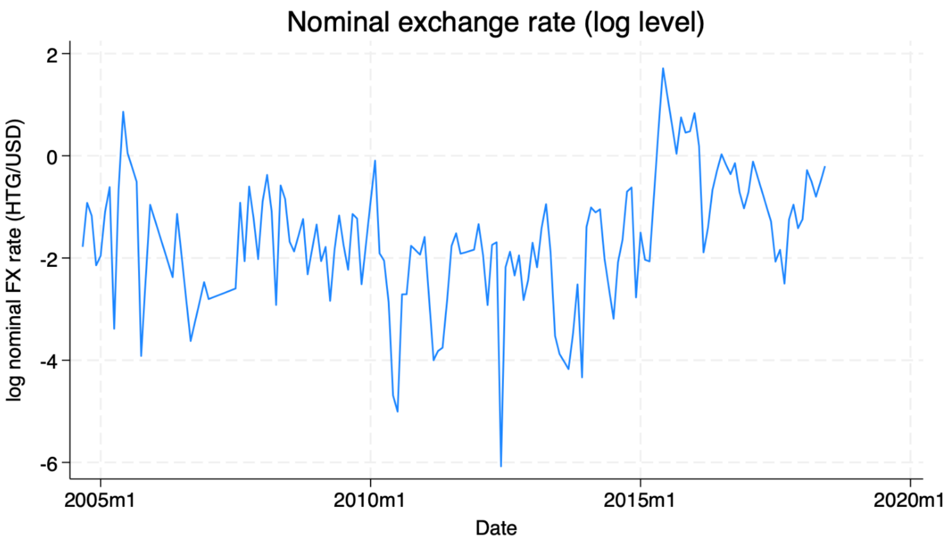
A.1 Stationarity Tests

Table 1 reports DF–GLS and augmented Dickey–Fuller (ADF) unit-root tests for all price series and for the nominal exchange rate.[1] All unit-root tests are conducted allowing for an intercept but no deterministic trend. This specification is consistent with the absence of clear linear trends in the inflation series and with our maintained specification for price levels. Critical values are selected accordingly. All CPI level series (in logs) behave as integrated processes: DF–GLS and ADF statistics are well above conventional critical values, with a few borderline rejections for specific sub-indices (transport, leisure and communication). In contrast, all inflation series (first differences of log CPI) are strongly stationary, with test statistics far below the 1% critical values. The monthly nominal depreciation of the gourde against the US dollar ($\Delta \ln(\text{EX})$, denoted *dln_fx*) also appears stationary.

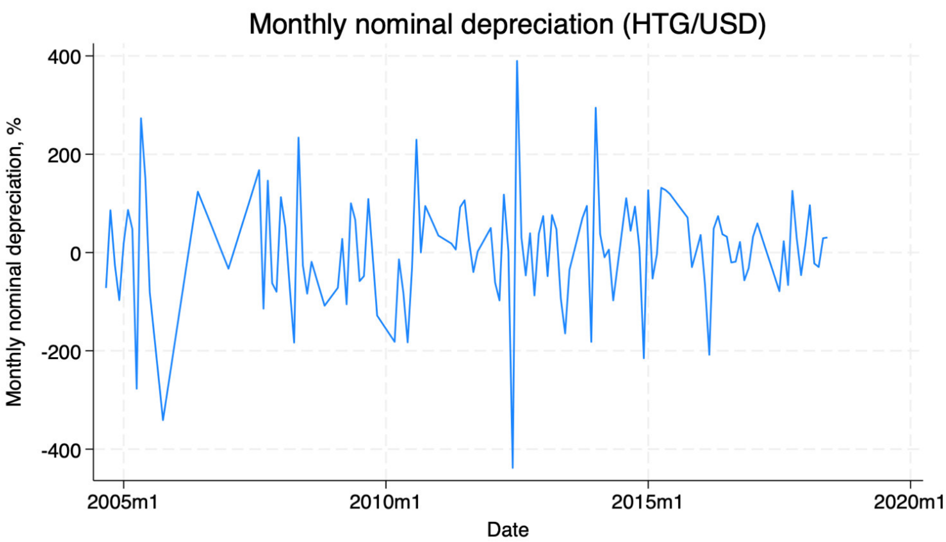
[1] Phillips–Perron tests, not reported to save space, yield very similar conclusions.



(a) Headline, local and imported CPI inflation



(b) Nominal HTG/USD exchange rate (level)



(c) Monthly nominal depreciation, *dln_fx*

Figure 1: Time-series of CPI inflation and nominal exchange rate

Table 1: Stationarity tests for CPI levels, inflation and exchange rate

Variable	DF–GLS	ADF
CPI levels (log series)		
ln_cpi_nat	0.89	−0.04
ln_cpi_local	1.26	−0.01
ln_cpi_import	0.35	−0.17
ln_cpi_food	0.80	0.23
ln_cpi_clothes	0.82	1.03
ln_cpi_catering	0.04	−0.35
ln_cpi_housing	−0.43	0.17
ln_cpi_communication	0.65	−4.86
ln_cpi_alcohol	2.19	2.20
ln_cpi_transport	−0.63	−2.58
ln_cpi_leisure	−2.82	−3.29
ln_cpi_imp_food	−0.31	−0.44
ln_cpi_imp_communicat	1.87	0.12
ln_cpi_imp_clothes	0.76	2.40
ln_cpi_imp_housing	−1.00	−1.79
ln_cpi_imp_transport	−1.74	−2.35
ln_cpi_imp_leisure	1.62	2.45
Inflation (first differences)		
pi_nat	−3.45	−8.23
pi_local	−2.56	−9.95
pi_import	−4.42	−7.41
pi_food	−2.96	−9.00
pi_clothes	−3.48	−6.88
pi_catering	−3.70	−8.38
pi_housing	−3.81	−11.06
pi_communicat	−10.22	−13.11
pi_alcohol	−2.68	−9.94
pi_transport	−4.84	−11.65
pi_leisure	−4.27	−9.05
pi_imp_food	−4.20	−5.91
pi_imp_communicat	−1.65	−10.55
pi_imp_clothes	−1.07	−9.48
pi_imp_housing	−4.31	−7.90
pi_imp_transport	−6.17	−11.07
pi_imp_leisure	−2.35	−9.87
Exchange rate		
dln_fx	−2.90	−10.31

Notes: DF–GLS and ADF statistics refer to tests run

The evidence confirms a pattern that is standard for price series in developing economies: CPI levels follow I(1) processes, while inflation and nominal depreciation are I(0). This motivates modelling inflation in differences, without error-correction terms.

A.2 Zivot–Andrews Structural Break Tests

Table 2 reports Zivot–Andrews tests allowing for a single endogenous break in the intercept. For each variable, we report the minimum t-statistic over possible break dates and the associated break index $t_{\min}$ (measured relative to the start of the monthly sample in 2004m1). For ease of interpretation, we also indicate an approximate calendar break date.

Table 2: Zivot–Andrews structural break tests

Variable	$t_{\min}$	10% crit.	5% crit.	1% crit.	Break index	Approx. date
CPI levels (log series)						
ln_cpi_nat	−2.81	−4.58	−4.80	−5.34	67	2009m7
ln_cpi_local	−2.92	−4.58	−4.80	−5.34	144	2015m12
ln_cpi_import	−2.68	−4.58	−4.80	−5.34	84	2010m12
ln_cpi_food	−2.90	−4.58	−4.80	−5.34	67	2009m7
ln_cpi_transport	−5.52	−4.58	−4.80	−5.34	34	2006m10
Inflation						
pi_nat	−9.88	−4.58	−4.80	−5.34	51	2008m3
pi_local	−5.38	−4.58	−4.80	−5.34	51	2008m3
pi_import	−9.11	−4.58	−4.80	−5.34	51	2008m3
Exchange rate						
dln_fx	−10.17	−4.58	−4.80	−5.34	128	2014m8
pi_nat (bis)	−9.88	−4.58	−4.80	−5.34	51	2008m3

Notes: $t_{\min}$ is the minimum t-statistic over all candidate break dates. The break index corresponds to the observation number in the time series, with index 1 = 2004m1. Critical values are from Zivot and Andrews (1992), model C (intercept break with trend).

For all CPI level series, the ZA statistics remain well above the 10% critical value, confirming the absence of statistically significant structural breaks in the price level, even though the estimated break dates cluster in the mid- to late-2000s and early 2010s. These level shifts are therefore not strong enough to justify introducing explicit break dummies in the CPI level equations, and they do not systematically line up with the election periods studied in the main text.

By contrast, the inflation series and the nominal depreciation display very large negative ZA statistics, indicating strong breaks in short-run dynamics. For headline inflation (π_{nat}), the break is located around 2008m3, which is consistent with the 2007–2008 international food and fuel price shock. The nominal exchange-rate depreciation exhibits a later and pronounced break around 2014m8, signalling a regime of heightened FX instability that predates the 2015 electoral cycle. Importantly, there is no evidence of additional breaks in CPI levels or inflation close to the 2010–2011 or 2015–2016 election episodes considered in the main analysis.

A.3 Autocorrelation Structure

All inflation series exhibit strong AR(1) dependence, with first-order ACF and PACF coefficients typically between 0.30 and 0.65. Some sub-series display mild AR(2) effects, but these are small relative to the dominant AR(1) component. The exchange-rate depreciation series (*dln_fx*) displays noticeably weaker persistence than inflation, with a modest positive first-order autocorrelation (ACF(1) $\approx$ 0.25) and little higher-order structure.

Table 3: Autocorrelation and partial autocorrelation (lag 1)

Variable	ACF(1)	PACF(1)
pi_nat	0.463	0.466
pi_local	0.353	0.355
pi_import	0.488	0.490
pi_food	0.393	0.395
pi_clothes	0.538	0.540
pi_imp_food	0.636	0.638
pi_imp_clothes	0.435	0.436
dln_fx	0.253	0.253

Figures 2 and 3 display the full ACF and PACF for headline inflation and the nominal depreciation, respectively. The patterns support the use of an AR(1) specification as the preferred dynamic structure in the baseline model, with AR(2) models reported as robustness checks, and suggest that a contemporaneous FX term is sufficient to capture the short-run impact of exchange-rate movements on inflation, without additional FX lags.

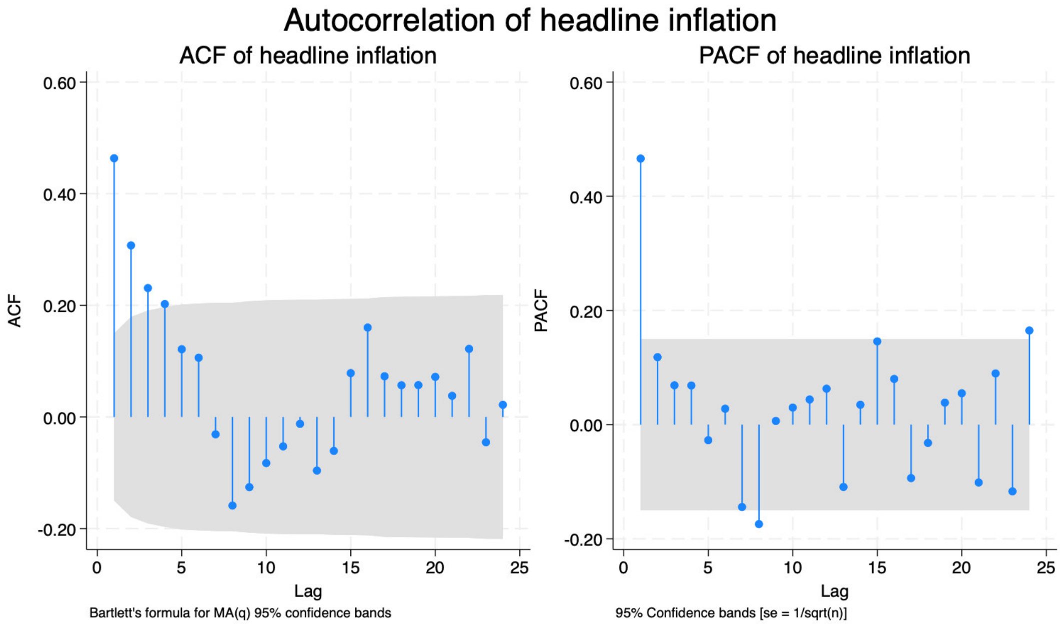


Figure 2: ACF and PACF of headline inflation, π_{nat}

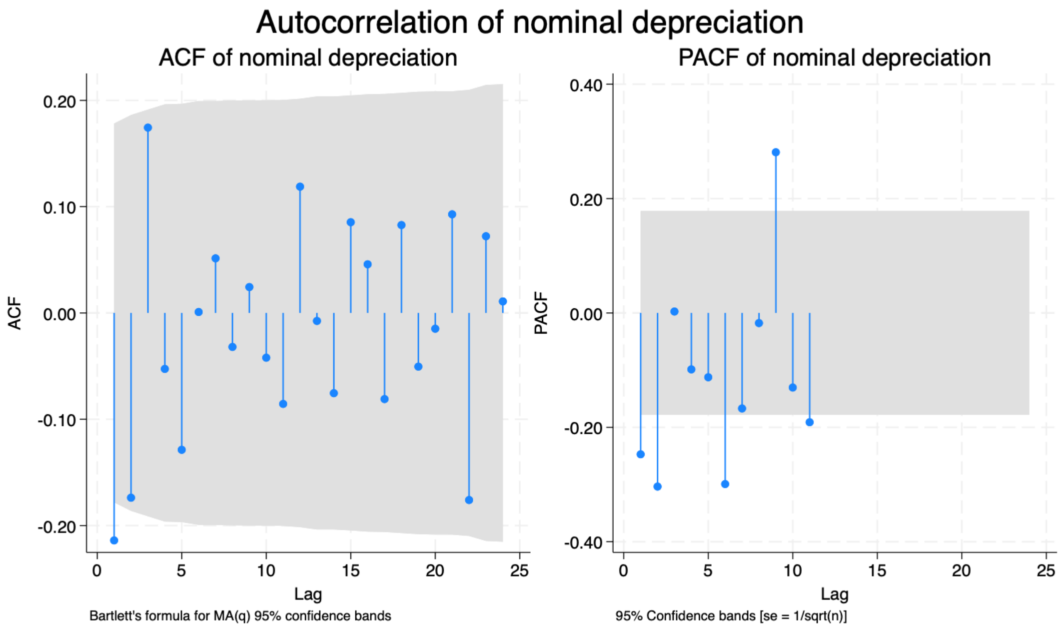


Figure 3: ACF and PACF of nominal depreciation, *dln_fx*

A.4 Johansen Cointegration Tests

We find no evidence of cointegration among the main CPI components. The Johansen tests on log-levels of national, local, imported, food and clothing sub-indices yield a cointegration rank of zero, in line with inflation being an I(0) process and with the absence of long-run price relationships in the sample.

Table 4: Johansen cointegration tests (log CPI levels)

Model	Rank r	Trace stat	Max-eigen stat	SBIC	HQIC	AIC
Nat + local + food & clothes	0	203.74	65.13	−58.92	−59.54	−59.96

A.5 Implications for Econometric Modeling

The diagnostic tests indicate that CPI levels are non-stationary, whereas all inflation series and the nominal depreciation rate are stationary. Inflation shows strong AR(1) persistence across all components and, for some sub-indices, mild but negligible AR(2) dependence. The exchange-rate depreciation is stationary and only moderately autocorrelated, which justifies treating it as a short-run control rather than part of a long-run equilibrium.

The structural-break analysis reveals no significant breaks in CPI levels and no breaks near the 2010–2011 or 2015–2016 election episodes. The economically relevant breaks in short-run dynamics occur in the mid- to late-2000s and are consistent with the international food and fuel price shock, as well as a change in the behaviour of the nominal exchange rate around 2014m8. Johansen tests further show no evidence of cointegration among the main CPI components.

Taken together, these results justify modelling inflation in a purely short-run autoregressive framework without deterministic trends or error-correction terms, and with the nominal exchange-rate depreciation included as an exogenous determinant of short-run inflation:

$$\pi_t = \alpha + \rho\pi_{t-1} + \beta \cdot \text{Election}_t + \gamma' X_t + \varepsilon_t,$$

where X_t includes disaster, riot and fuel shocks, as well as the monthly nominal depreciation of the gourde against the US dollar. Estimation in the main text uses HAC standard errors to account for possible residual autocorrelation and heteroskedasticity.

Appendix B. Randomization Inference

B.1 Methodology

We implement a Fisher-style randomization inference (RI) test to evaluate whether the local-projection election effects could arise from arbitrary calendar reassignment. Let $E = \{t_1, \ldots, t_5\}$ denote the months of the five observed elections. A placebo sequence $E^{(r)}$ is generated by randomly drawing five distinct months from the sample period, excluding a ± 6-month window around each real election in order to avoid trivial overlap. For each draw $r = 1, \ldots, 500$, we construct a placebo indicator $\text{elec0}^{(r)}$, re-estimate the local projection model for horizons $h = 0, \ldots, 6$, and extract the coefficients $\beta^{(r)}(h)$.

The short-run post-election response is summarised by

$$S^{(r)} = \frac{\beta^{(r)}(1) + \beta^{(r)}(2)}{2},$$

which focuses on the horizons exhibiting the clearest dynamics in the baseline results. The statistic computed on the true election sequence is

$$S_{\text{real}} = \frac{\beta_{\text{real}}(1) + \beta_{\text{real}}(2)}{2}.$$

The randomization p-value is given by

$$p = \frac{1 + \#\left\{S^{(r)} \geq S_{\text{real}}\right\}}{R + 1}, \qquad R = 500.$$

B.2 Results

The placebo distribution of $S^{(r)}$ has mean 0.073 and standard deviation 0.269. The real election sequence yields $S_{\text{real}} = 0.223$. Among the 500 placebo draws, 133 produce a value $S^{(r)} \geq S_{\text{real}}$, yielding a randomization p-value of $p = 0.27$.

The figure provided in the main text displays the histogram of placebo statistics along with the location of S_{real}. The placebo sequences do not replicate the coherent post-election acceleration of inflation observed in the real data, although the small number of observed elections naturally limits the inferential power of the test.

B.3 Discussion

Randomization inference is intentionally non-parametric and does not rely on functional-form assumptions, but its power depends heavily on the number of treatment events. With only five elections in our sample, the RI procedure can detect only exceptionally large and noise-free treatment effects. In this setting, the randomization results should therefore be interpreted as a transparency exercise rather than a definitive inferential test.

Importantly, the placebo sequences show no tendency to generate spurious “election cycles”, and the real-data statistic lies toward the right tail of the placebo distribution. Combined with the autoregressive evidence, the individual event studies, and the symmetric local projections, the RI exercise is consistent with—and does not contradict—our interpretation of a short-lived post-election inflation acceleration.

Appendix C. Additional Robustness Checks

C.1 Exchange-rate dynamics around elections

A natural concern is that the short-run inflation response identified in the event-study and local-projection exercises might simply reflect episodes of sharp exchange-rate depreciation that happen to coincide with elections. To assess this, we construct ±6-month event windows around each electoral sequence and plot the corresponding monthly nominal depreciation series *dln_fx*.

Figure 4 overlays the exchange-rate paths for the five electoral sequences. While the gourde experiences sizable depreciations in some periods—most notably around the 2015–16 sequence—there is no systematic pattern of large FX movements centered on election months. Several episodes display essentially flat exchange-rate dynamics throughout the window, and others feature depreciation episodes that precede or follow elections by several months.

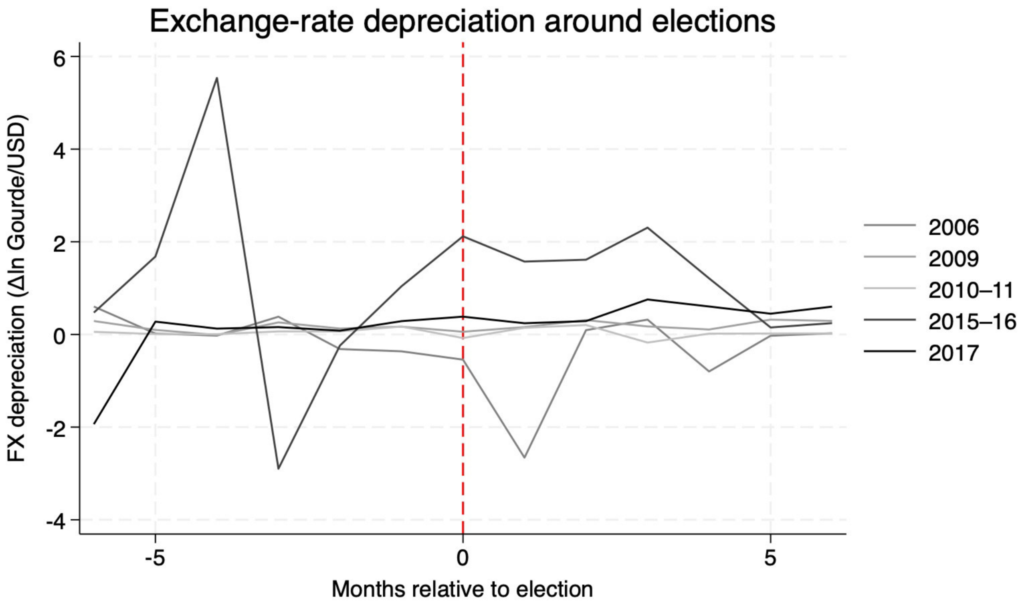


Figure 4: Exchange-rate depreciation in a ±6-month window around five electoral sequences

To summarise these patterns, Figure 5 reports the stacked average of *dln_fx* across all electoral windows. The average depreciation remains modest and shows no clear spike at $t = 0$, the election month. FX pressures tend to build up gradually over the broader period surrounding elections and occasionally intensify in specific episodes, but they do not exhibit a tight synchronization with the electoral calendar.

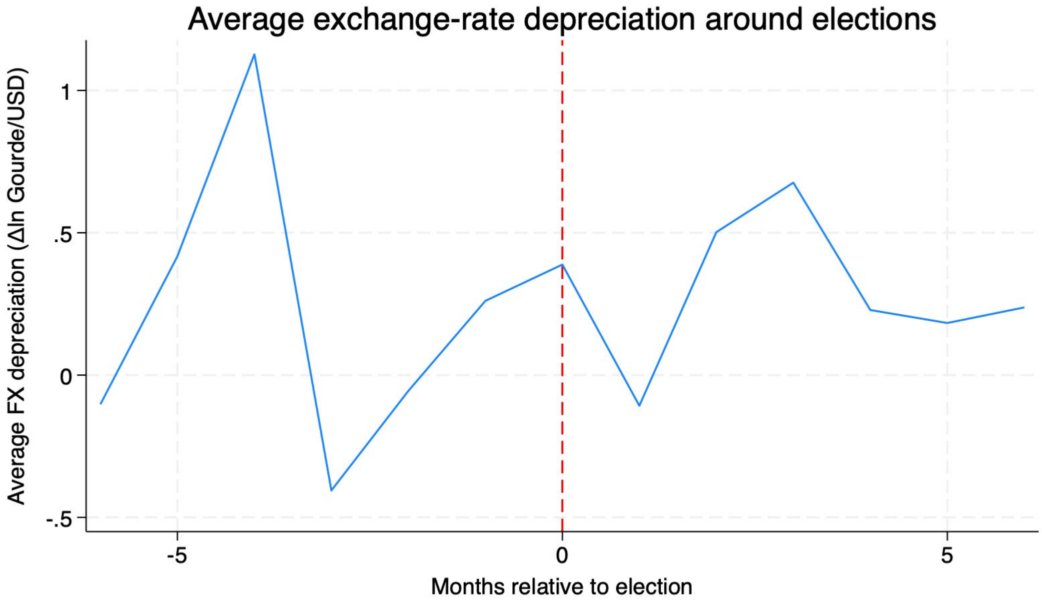


Figure 5: Average exchange-rate depreciation in a ±6-month window around elections

Taken together, these results suggest that the inflation responses documented in the main text are unlikely to be driven mechanically by exchange-rate movements that are perfectly aligned with election dates. Instead, FX depreciation appears as a broader macroeconomic background shock, which is already controlled for in the regressions through contemporaneous (and, in robustness checks, lagged) values of $\Delta \ln(\text{EX})$.

C.2 Lagged exchange-rate controls

As an additional robustness check, we augment the baseline AR(1) regression for national inflation with a lag of the monthly nominal depreciation. Table 5 compares the baseline specification, which includes contemporaneous depreciation only, with a version that also includes $L.\Delta \ln(\text{HTG/USD})$.

Table 5: Robustness to lagged exchange-rate depreciation

	(1) Baseline	(2) + lagged FX
L. π_{nat}	0.411***	0.417***
	(0.149)	(0.144)
Election month (t)	0.049	0.067
	(0.106)	(0.094)
Disaster shock	0.114	0.109
	(0.272)	(0.270)
Riot shock	−0.167	−0.114
	(0.113)	(0.120)
Fuel-price shock	−0.238	−0.209
	(0.334)	(0.331)
Monthly nominal depreciation	0.078	0.092
(HTG/USD, %)	(0.080)	(0.098)
Post-2015m7	0.261***	0.267**
	(0.098)	(0.108)
L.Monthly nominal depreciation		−0.076
(HTG/USD, %)		(0.093)
Constant	0.307***	0.307***
	(0.117)	(0.115)
R^2	0.228	0.235
Observations	165	165

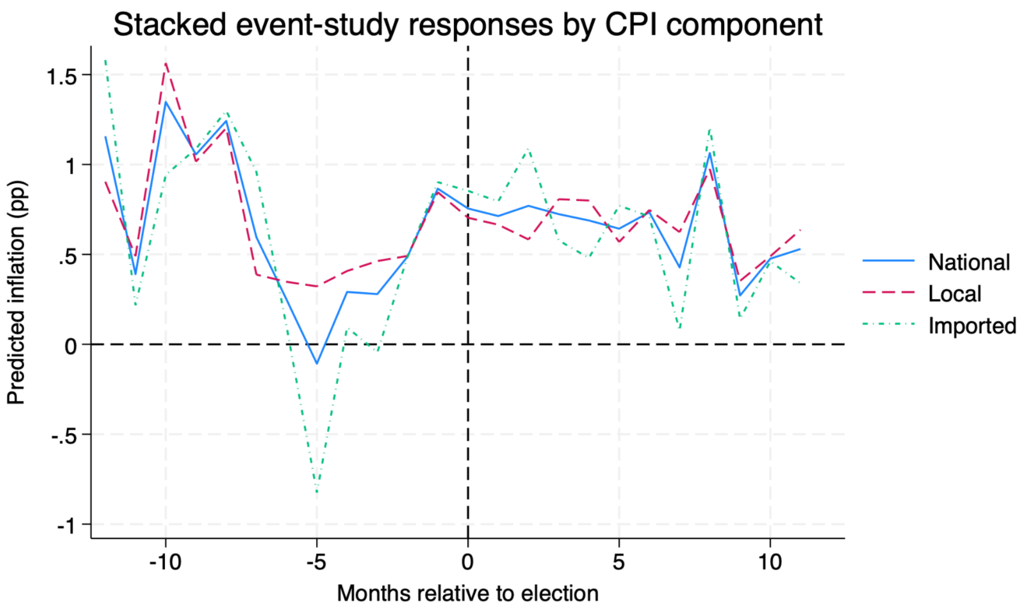
Notes: Dependent variable is monthly national inflation, π_{nat}. All specifications include the election-month dummy, indicators for natural disasters, riots and fuel-price shocks, contemporaneous exchange-rate depreciation and a post-2015m7 dummy. Column (2) additionally includes the first lag of depreciation. HAC standard errors (Bartlett kernel, 12 lags) in parentheses. * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$.

Including the lagged depreciation term leaves the coefficients on the election dummy and on the post-2015 dummy essentially unchanged, and the lag itself is small and statistically insignificant. The R^2 rises only marginally (from 0.228 to 0.235). These results suggest that adding further exchange-rate lags does not alter the estimated electoral effects and that a parsimonious specification with contemporaneous depreciation is sufficient for the baseline analysis.

C.3 Headline versus CPI components

To verify that the election-related inflation dynamics are not specific to headline CPI, we estimate the stacked event-study specification separately for national, local and imported inflation. The specification is identical to the main ES model - AR(1) persistence, episode fixed effects, and controls for domestic shocks and monthly nominal depreciation. Figure 6 presents the average predicted inflation around elections for each CPI component. All three series exhibit a similar short-run acceleration in the months immediately surrounding the vote. Although imported inflation is mechanically more sensitive to exchange-rate movements and local inflation is more closely tied to domestic supply conditions, the election-related pattern appears consistently across components. The magnitude of the spikes varies slightly, with imported prices displaying somewhat sharper movements, but the temporal profile is broadly aligned. These results confirm that the short-lived inflationary response to elections is not driven exclusively by imported goods or by local supply disruptions. Instead, the election cycle affects a wide set of consumer prices, consistent with a generalized short-run demand or liquidity impulse and with temporary market disruptions around electoral periods.

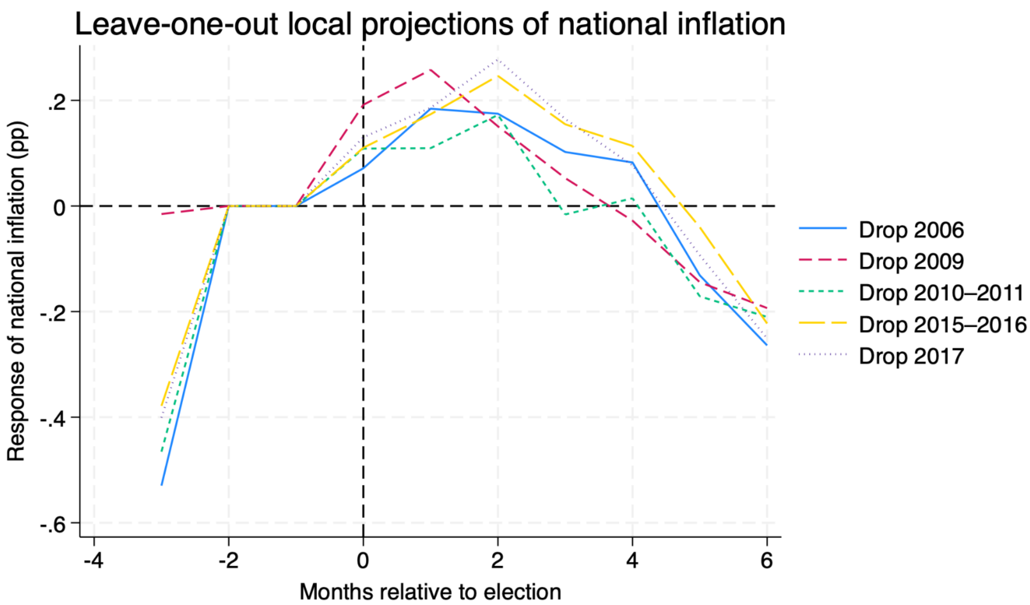
Figure 6: Stacked event-study responses by CPI component



C.4.1 Leave-one-out local projections

A natural concern with only five electoral episodes is that the average impulse response may be disproportionately influenced by one particular election. To assess this possibility, we re-estimate the symmetric local-projection specification five times, each time excluding one of the five electoral sequences from the sample. Figure 7 plots the resulting leave-one-out impulse responses of national inflation to an election shock.

Figure 7: Leave-one-out local projections of national inflation

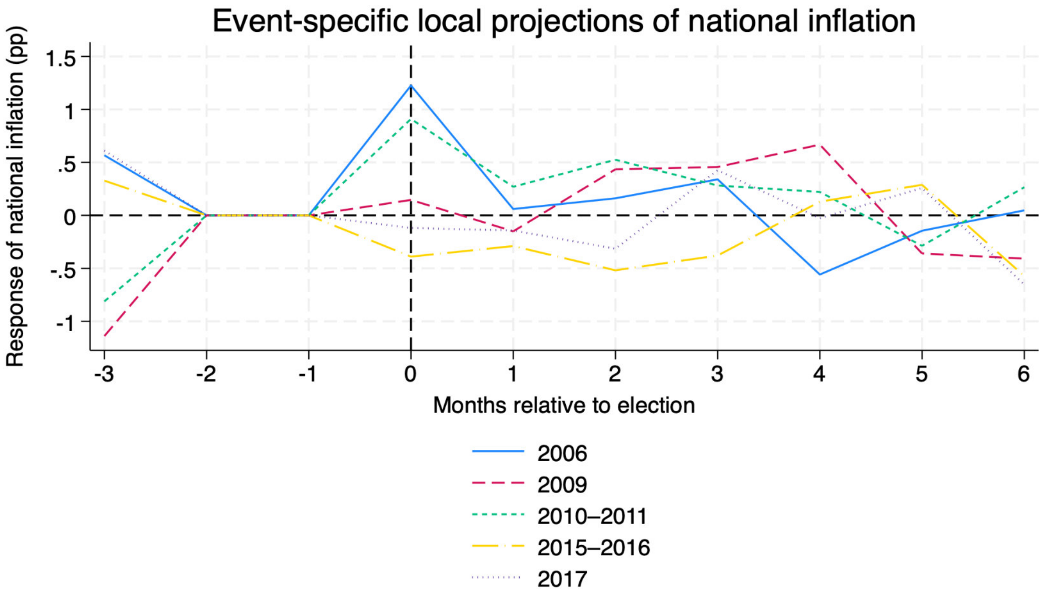


The leave-one-out profiles are remarkably similar across exclusions. In all cases, inflation increases in the months immediately following an election, peaking around horizons $h = 1$–2 before gradually returning toward baseline. Dropping any individual election shifts the magnitude slightly - reflecting expected heterogeneity across episodes - but the qualitative pattern remains unchanged. No single electoral sequence is capable of eliminating the short-run post-election inflation spike. This confirms that the baseline LP result is not driven by an outlier episode and reflects a robust average response across electoral cycles.

C.4.2 Event-specific local projections

As a complementary robustness check, we estimate event-specific LPs, allowing the response of inflation to differ across electoral sequences. We run separate LP regressions for each of the five election episodes and report the implied impulse responses in Figure 8.

Figure 8: Event-specific local projections for five electoral sequences



The event-specific responses reveal noticeable heterogeneity in magnitude, but a broadly similar temporal profile. Elections such as 2006 and 2010–2011 are associated with clear short-run increases in inflation around the election month, while other episodes (notably 2015–2016 and 2017) display more muted and sometimes slightly negative responses. In all cases, however, the effects remain short-lived: responses tend to peak between $h = 0$ and $h = 3$ and then converge back toward zero by $h = 5$–6. Taken together with the leave-one-out analysis, these results reinforce the interpretation that the baseline LP captures a coherent average pattern rather than being driven by a single unusually inflationary election.

A common feature across these local projections is the small "plateau" observed in the immediate post-election months, with responses changing little between $h = 0$ and $h = 2$. This reflects the limited month-to-month variation available within each episode and the smoothing inherent in LP estimates with short samples. The flat region is therefore best interpreted as a descriptive feature of the average short-run adjustment rather than as evidence of a prolonged inflationary impulse.