Quantitative Finance and Economics

QFE, 9(4): 780–809.
DOI: 10.3934/QFE.2025027
Received: 02 May 2025
Revised: 08 August 2025
Accepted: 18 September 2025
Published: 15 October 2025

https://www.aimspress.com/journal/QFE

Research article

Spillover effects of U.S. business cycle shocks on monetary policy in Pacific Alliance Countries (PAC)

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Abstract: This paper investigated the spillover effects of U.S. structural shocks on the monetary policy of Pacific Alliance (PA) countries. We began by estimating a backward-looking Taylor rule using panel data, incorporating the U.S. output gap and policy rate as external factors to assess the sensitivity of PA central banks to U.S. economic conditions. Additionally, unlike traditional approaches that treat foreign macroeconomic variables as exogenous, we recognized that they are subject to supply and demand shocks. To address this, we employed a two-stage strategy. First, we used a structural vector autoregression (SVAR) with long-run restrictions to identify these shocks. Then, we evaluated the dynamic responses of PAC policy rates using a distributed lag model. Our findings revealed the following: (i) U.S. GDP and policy rates plays a significant role in the monetary policy reaction functions of PA central banks; (ii) the nature of the shock—whether demand or supply—results in differing responses, with PA countries generally reacting more strongly to U.S. demand shocks; and (iii) the persistence of these responses varies across countries.

Keywords: Taylor rule; central banking behavior; open economies; business cycle shocks; SVAR

JEL Codes: C32, E52, F41

1. Introduction

The U.S. economy plays a crucial role in the global economic landscape. Its influence has expanded over time due to globalization, financial integration, and trade interdependencies (Kose et al., 2017; Dées and Galesi, 2021; Bräuning and Sheremirov, 2022). A particularly relevant issue concerns the spillover effects of economic shocks originating in the U.S. on small open and emerging economies, which are more vulnerable to external fluctuations (Almansour et al., 2015; Briguglio, 2016). A substantial body of literature has examined these spillover effects from various perspectives, including the impact of U.S. uncertainty, business cycles, and monetary policy on key macroeconomic aggregates in emerging markets (Bhattarai et al., 2020; Trung, 2018; Tillman et al., 2019). One critical aspect is how monetary policy in these economies responds to external shocks, as an optimal policy response can mitigate or even offset adverse effects (Drechsel et al., 2019).

While there is some consensus on the impact of U.S. shocks on emerging economies, we identify a methodological challenge in empirical research that examines these effects, particularly in studies focusing on monetary policy responses. The conventional approach to address the effects on policy rates often treats external shocks as fully exogenous, which may lead to misinterpretation. Although U.S. shocks are exogenous to emerging economies, they are endogenous to the U.S. economy itself, as they stem from underlying supply and demand fluctuations. For example, assessing the impact of U.S. business cycle fluctuations on emerging market GDP requires identifying the underlying structural shocks. Ignoring these shocks creates an omitted variable problem, as both U.S. supply and demand shocks indirectly affect emerging markets, leading to biased estimates.

Our paper addresses this issue through a two-stage empirical strategy. First, we identify structural supply and demand shocks in the U.S. using a structural vector autoregression (SVAR) model with long-run restrictions, following Blanchard and Quah (1989). This allows us to distinguish between permanent supply shocks and demand shocks. To validate our identification strategy, we assess the historical decomposition of these shocks and examine their consistency with major economic crises from 1950 to 2024. In the second stage, we estimate the response of Pacific Alliance Countries (PAC) monetary policy rates to lagged structural shocks. The coefficients in this regression can be interpreted as impulse response functions.

Furthermore, we investigate whether PA central banks adhere to a Taylor rule framework and assess the relevance of U.S. business cycles and monetary policy in shaping their policy responses. To this end, we estimate a backward-looking Taylor rule using panel data and the generalized least squares (GLS) method with fixed effects.

Our empirical strategy also aligns with the theoretical framework of Medina and Soto (2007), who developed a DSGE model for a small, open, commodity-exporting economy. Their model highlights how external demand and terms-of-trade shocks transmit to domestic macroeconomic variables, validating the relevance of analyzing U.S. shocks in export-oriented Latin American countries.

The main contribution of our study is the application of this two-stage approach as an alternative to instrumental variable estimation in this topic, which can introduce substantial asymptotic bias if instruments are weak (Bound et al., 1995). Additionally, we focus on the Pacific Alliance (PA) countries—Peru, Chile, Colombia, and Mexico—where research on this topic remains relatively scarce. Our empirical strategy is flexible, allowing us to examine the differential and persistent effects of supply and demand shocks from the U.S.

The Pacific Alliance countries provide a compelling case study due to their adoption of shared economic policies in recent years, including inflation targeting, fiscal rules, flexible exchange rates, and openness to trade and finance. This policy mix has the potential to either mitigate or amplify the impact of external shocks on domestic economies, providing valuable evidence for countries with that characteristic (Rodríguez and Vasallo, 2023). Additionally, these countries maintain strong trade and financial ties with the U.S., making them highly exposed to fluctuations in the U.S. economy. In contrast, a more diverse sample with varying monetary frameworks and policy approaches could introduce significant heterogeneity, making it harder to draw clear conclusions. By concentrating on the Pacific Alliance Countries, this study ensures that the observed effects primarily capture the influence of external shocks rather than country-specific factors, enhancing the reliability of the analysis.

Our findings reveal the following: (i) U.S. GDP and policy rates play a significant role in the monetary policy reaction functions of PA central banks; (ii) the nature of the shock—whether demand or supply—results in differing responses, with PA countries generally reacting more strongly to U.S. demand shocks; and (iii) the persistence of these responses varies across countries. Peru, Chile, and Mexico show acyclical reactions to supply shocks, indicating minimal policy adjustments, whereas Colombia responds more assertively. In the case of demand shocks, all four countries tighten monetary policy, with Chile and Mexico exhibiting similar peak responses. However, Mexico's adjustment is more persistent, while Chile's effect diminishes more rapidly. These variations underscore the differing levels of exposure and policy sensitivity to external shocks.

The remainder of this paper is structured as follows: Section 2 presents a review of empirical and theoretical literature. Section 3 outlines the empirical model and methodology. Section 4 describes the data and presents the results of our Taylor rule estimation and the identification of U.S. supply and demand shocks. Section 5 concludes with a discussion of our key findings and policy implications. Finally, Section 6 presents the robustness checks.

2. Literature review

Economic literature provides several explanations for how U.S. shocks are transmitted to emerging economies. For instance, Kandil (2009) showed that U.S. GDP growth has a significant impact on Latin American emerging markets' GDP and inflation, with trade being a key transmission channel. Higher U.S. imports can stimulate consumption and investment in emerging markets, while financial integration can amplify these effects. These factors contribute to inflationary pressures, prompting monetary policy responses in PA countries.

Fluctuations in the U.S. business cycle also influence Federal Reserve policy, generating spillover effects on emerging markets. Maćkowiak (2007) found that U.S. monetary policy shocks have rapid and strong effects on exchange rates and interest rates in emerging markets. This aligns with the seminal works of Calvo et al. (1993) and Calvo and Reinhart (2002), who argued that "fear of floating" leads emerging market central banks to adjust monetary policy in response to exchange rate fluctuations. Additionally, Rey (2015) and Miranda-Agrippino and Rey (2015) highlighted the Fed's role in driving global financial cycles, affecting asset prices in both advanced and emerging economies. These findings are further supported by Bräuning and Ivashina (2019), who showed that during a typical U.S. monetary easing cycle, borrowers in emerging markets experience a 32-percentage-point greater incr ease in foreign bank lending than those in developed markets, followed by a rapid credit contraction once U.S. monetary policy tightens. In summary, international trade and financial channels play a crucial role in the transmission of U.S. shocks to emerging markets.

Our research contributes to the extensive literature on U.S. spillover effects on emerging markets. Previous studies have analyzed the impact of structural supply and demand shocks on macroeconomic aggregates in emerging economies, emphasizing their significant influence on output, inflation, and interest rates (Ocampo and Ojeda-Joya, 2022; Feldkircher and Huber, 2015; Feldkircher and Tondl, 2020). These works provide valuable insights into the effects of U.S. shocks and their transmission mechanisms. They highlight that U.S. shocks play a crucial role in shaping monetary policy responses and that both financial and trade channels facilitate the global propagation of these shocks. Specifically, we extend the work of Canova (2005), who analyzed the effects of U.S. supply and demand shocks on Latin American countries. However, our study additionally examines how domestic monetary policy responds to these external shocks. We also draw on Cavallo and Ribba (2015), who investigated how common macroeconomic shocks affect business cycle fluctuations in Euro area countries using a structural VAR approach. Their findings underscore the importance of distinguishing between common and country-specific shocks.

Our research is closely related to the aforementioned studies; however, a key distinction is that when identifying monetary policy shocks, they typically treat them as exogenous (except for Canova, 2005). In contrast, we not only employ different identification strategies but also consider these shocks as endogenous, focusing exclusively on the monetary policy response.

More recent contributions at the frontier of the literature have explored the role of uncertainty shocks, credit cycles, inflation targeting, and the informational content of monetary policy meetings in shaping macroeconomic outcomes in emerging markets (Trung, 2018; Bräuning and Ivashina, 2019; Gai and Tong, 2022; Chión-Chacón and Alvarez, 2025). These studies suggest that sentiments of optimism or pessimism conveyed through policy communications can be a key determinant in amplifying spillover effects. Moreover, the financial vulnerabilities of emerging economies (including market imperfections and weak inflation expectation anchoring) can significantly influence the magnitude of external shocks (Ahmed et al., 2021).

3. Research methodology

The empirical strategy in this research consists of estimating a back-looking Taylor rule with panel data for PAC by a generalized least squares (GLS) estimator with cross-section weights to correct for heteroskedasticity problems. Also, our model was estimated under fixed effects in a transversal unit[1].
where mpr_t is the monetary policy rate; π_t is the level of inflation in period t; $\bar{y}_t$ denotes the domestic output gap; Δy_{it} is domestic GDP growth; Δy^*_{it} is the U.S. real GDP growth; $\overline{y^*_t}$ is the U.S. output gap; and mpr^*_t is the U.S. monetary policy rate. The model is specified as follows:

[1] A random effects model assumes that unobserved country-specific heterogeneity is uncorrelated with the explanatory variables, which may not hold in the context of inflation-targeting (IT) economies. In the case of the Pacific Alliance countries, institutional factors such as central bank credibility, the degree of independence, historical inflation patterns, and the formation of inflation expectations tend to be persistent over time. These structural characteristics are likely correlated with key macroeconomic variables such as GDP and inflation, violating the exogeneity assumption required for random effects estimation. Consequently, employing a fixed effects model is preferable, as it accounts for time-invariant heterogeneity and mitigates potential estimation biases.

$$mpr_{it} = c_0 + c_1 mpr_{it-1} + c_1 \pi_{it} + c_2 \overline{y_{it}} + c_3 \overline{y_{it}^*} \quad (1)$$

$$mpr_{it} = d_0 + d_1 mpr_{it-1} + d_2 \pi_{it} + d_3 \overline{y_{it}} + d_4 mpr_{it}^* \quad (2)$$

$$mpr_{it} = g_0 + g_1 mpr_{it-1} + g_2 \pi_{it} + g_3 \Delta y_{it} + g_4 \Delta y_{it}^* \quad (3)$$

$$mpr_{it} = e_0 + e_1 mpr_{it-1} + e_2 \pi_{it} + e_3 \Delta y_{it} + e_4 mpr_{it}^* \quad (4)$$

where $i = 1,2,3,4$ corresponds to the cross-sectional units (countries), and t represents the time dimension. For both the U.S. and PAC, we use real GDP growth and the output gap as proxies for economic activity.

We estimate Equations (1)–(4) under the premise that the U.S. business cycle may be highly correlated with the monetary policy rate in emerging markets. This correlation arises because the policy rate typically reacts to economic conditions (Taylor, 1993). To mitigate potential multicollinearity among explanatory variables, we do not include both U.S. variables in the same regression specification.

A common empirical challenge when estimating structural equations such as Taylor-type rules is the simultaneity problem, which can introduce bias, particularly in domestic variables. However, our empirical strategy is robust for several reasons. First, while an exogenous monetary policy shock raises the interest rate, which in turn dampens demand and output, these effects typically unfold over time, reducing concerns about simultaneity bias. Second, our primary objective is not to establish causal relationships or identify structural parameters but rather to provide empirical evidence on the statistical significance of key relationships. This distinction is crucial in justifying our choice of estimation strategy.

Finally, there is empirical evidence suggesting that Taylor rule estimates obtained via two-stage least squares (2SLS) often exhibit greater bias than those derived from ordinary least squares (OLS) due to the issue of weak instruments (Carvalho et al., 2021). Given this limitation, GLS[2] estimates provide a more reliable and appropriate estimation method for our analysis.

To decompose the demand and supply shocks of the U.S., we use an SVAR approach to identify historical structural shocks on the U.S. economy. The empirical strategy is the long-run restriction, proposed by Blanchard and Quah (1989). The structural model is set as follows:

$$B_0 Y_t = c + B_i \sum_{i=1}^{p} Y_{t-1} + e_t;\ e_t \sim WN\ (0, D) \quad (5)$$

where e_t are the uncorrelated structural shocks, assumed to follow a normal distribution with zero mean and diagonal var-cov matrix D. The objective is to identify the structural parameter from the reduced form:

$$Y_t = d + A_i \sum_{i=1}^{p} Y_{t-i} + u_t; u_t \sim WN(0, \Sigma) \quad (6)$$

[2] In our model, the presence of autocorrelation leads to inefficiency in the estimators. For this reason, the GLS estimator becomes BLUE.

where u_t are the correlated shocks, thereby Σ is a non-diagonal var-cov matrix. We can standardize the structural shock through $z_t = D^{-\frac{1}{2}}e_t \sim WN\ (0, I)$. Thus, we can obtain the correlated shocks as a linear combination of structural shock $u_t = B_0^{-1}D^{\frac{1}{2}}z_t = Sz_t$, where the dynamic is captured by B_0 matrix. We can identify the structural parameters imposing parametric restrictions on B or S if these matrices are triangular. On the other hand, our approach does not follow that strategy; instead, we impose parametric restriction on the sum of structural shock over time. To show that, we solved the reduced VAR model in VMA (∞):

$$A(L)Y_t = d + u_t$$

$$Y_t = A(L)^{-1}d + A(L)^{-1}u_t = \phi + \psi(L)u_t$$

where $\psi(L) = A(L)^{-1}$. The extensive form of the VMA model is:

$$Y_t = \phi + \psi_0 u_t + \psi_1 u_{t-1} + \cdots$$

$$\psi_0 = I$$

This, the sum of all structural shocks over time is described by the F matrix,

$$F = \psi(1) = \psi_0 + \psi_1 + \psi_2 + \psi_3 + \psi_4 \ldots$$

$$F = \begin{bmatrix} f_{11} & f_{12} \\ f_{12} & f_{22} \end{bmatrix}$$

Our strategy is to set $f_{12} = 0$ (as Blanchard and Quah, 1989),

$$Y_t = \phi + \begin{bmatrix} f_{11} & 0 \\ f_{12} & f_{22} \end{bmatrix} B_0^{-1} D^{\frac{1}{2}} \begin{bmatrix} z_1 \\ z_2 \end{bmatrix}$$

where z_1 is the supply shock, and z_2 is the demand shock. This identification relies on the assumption that demand shocks affect output only in the short run but not in the long run. This approach allows us to recover structural shocks by exploiting their distinct long-run implications for output. In particular, we follow the strategy proposed by Blanchard and Quah (1989), which restricts demand shocks from having permanent effects on output, while allowing supply shocks to do so.

We allow the model to estimate the long-run effects of the identified structural shocks on inflation, which may lead to puzzling results in some cases. We acknowledge this issue and address it empirically by estimating two models: one imposing $f_{22} = 0$ and the other $f_{22} \neq 0$, both while maintaining a recursive structure. In both cases, the response of inflation to a demand shock converges[3]. Therefore, to estimate a less restrictive model, we opt for the specification without exclusion restrictions. Consequently, it is important to recognize that, by definition, the structural shocks correspond to permanent supply shocks and transitory demand shocks.

To obtain the IRF, we follow Kilian (2009) and Dedola et al. (2017) strategy. We use a linear regression model with the structural shocks extracted in the previous model. The regression is:

$$x_{it} = a_0 + b_i \sum_{j=0}^{12} \epsilon_{t-j} + u_{1t} \tag{7}$$

[3] We refer to "convergence" when the response of inflation is approximately 0.009. This effect is estimated by both models.

$$x_{it} = b_0 + c_i \sum_{j=0}^{12} \gamma_{t-j} + u_{2t} \tag{8}$$

In models (7) and (8), u_{1t} and u_{2t} denote potentially correlated error terms. The variable x_{it} includes the monetary policy rate as well as other domestic macroeconomic aggregates such as real GDP growth, nominal exchange rates, inflation, and sovereign risk spreads (EMBI). The impulse response functions correspond to the estimated coefficients b_i and c_i over a 12-period horizon. To account for possible correlation in the error terms when conducting inference, we implement the Wild Recursive Block Bootstrap method. The variables ϵ_t and γ_t represent the structural demand and supply shocks from the U.S., respectively. The subscript i indexes the country, while t denotes time.

We argue that this approach to obtaining IRFs is robust. The issue of endogeneity, arising from correlation between structural shocks and local residuals, is mitigated because U.S. structural demand and supply shocks are assumed to be exogenous with respect to emerging economies. For instance, a positive technology shock in the U.S. that increases aggregate supply is largely unrelated to domestic monetary policy innovations. Similarly, a large-scale fiscal stimulus that boosts U.S. demand does not originate from, nor is contemporaneously caused by, domestic developments in emerging markets. While these U.S. shocks may generate endogenous responses in emerging economies, such as changes in interest rates, the exogeneity assumption ensures that parameter estimates remain unbiased. Table 1 presents a correlation-based metric, showing that the correlation between U.S. structural shocks and domestic residuals is very close to zero, supporting this identification strategy. Additionally, Table 2 presents formal hypothesis tests that do not reject the exogeneity assumption, thereby reinforcing the validity of our identification strategy.

Table 1. Correlations between structural shocks and errors of regression.

	Structural supply shocks	Structural demand shocks
Peru	$3.16\ x\ 10^{-17}$	$2.36\ x\ 10^{-17}$
Chile	$-5.71\ x\ 10^{-17}$	$2.08\ x\ 10^{-17}$
Colombia	$2.98\ x\ 10^{-16}$	$2.53\ x\ 10^{-16}$
Mexico	$1.13\ x\ 10^{-16}$	$4.85\ x\ 10^{-17}$

Note(s): correlations of regression 5 and 6.

Table 2. Endogeneity test.

Wu–Haussman test; Ho: structural shocks are exogenous										
	Structural demand shock					Structural supply shocks				
	mpr	GDP	Infl	ER	EMBI	mpr	GDP	Infl	ER	EMBI
Peru	0.52	0.68	0.32	0.19	0.09*	0.75	0.33	0.59	0.91	0.06*
Chile	0.26	0.06*	0.07*	0.78	0.41	0.68	0.68	0.40	0.32	0.06*
Colombia	0.44	0.72	0.56	0.35	0.07*	0.44	0.11	0.34	0.55	0.10
Mexico	0.08*	0.11	0.77	0.48	0.24	0.95	0.10	0.89	0.12	0.92

Notes: Reported values correspond to p-values from the Wu–Hausman endogeneity test. Instruments include 2–5 lags of the respective structural shocks. For Chile, in the EMBI and inflation specifications, an additional lag of the dependent variable was used as an instrument.

Similarly, when assessing the effects of these structural shocks on other domestic macroeconomic aggregates, we assume that there is no feedback from emerging market variables to the identified U.S. demand and supply shocks. This assumption is reasonable given that the emerging economies in our sample account for a relatively small share of global trade, financial markets, and output, implying that their economic fluctuations are unlikely to influence the U.S. economy in a meaningful way.

In general, Table 2 presents the results of the Wu–Hausman test applied to each estimated equation in order to assess the validity of the exogeneity assumption. The test examines whether the structural U.S. shocks, previously identified, are uncorrelated with the regression error terms. At the 5% significance level, we do not reject the null hypothesis of exogeneity in any case. This outcome reinforces the credibility of our identification strategy, indicating that the structural demand and supply shocks from the U.S. can be treated as exogenous to the macroeconomic dynamics of emerging market economies.

3.1. Data and samples

In our analysis, we collected the following variables for the U.S. economy: real GDP (measured in billions of chained 2017 dollars) and the Consumer Price Index (CPI) for all urban consumers (base period: 1982–1984 = 100). For the Pacific Alliance (PA) countries, we compiled data on monetary policy interest rates, headline CPI, and real GDP expressed in domestic currency. Nominal exchange rates are included as bilateral rates against the U.S. dollar, reflecting the value of each country's currency in terms of USD. Additionally, we incorporate the Emerging Market Bond Index (EMBI) as a measure of sovereign spreads. This spread serves as a proxy for perceived credit risk and external borrowing costs in each country.

All variables are expressed at quarterly frequency. Data sources include the International Monetary Fund (IMF), the Federal Reserve Economic Data (FRED) provided by the St. Louis Fed, and the official websites of the respective central banks.

The U.S. data span the period from 1947Q1 to 2019Q4. For the PA countries, monetary policy rate data are available from 2007Q1 to 2019Q4 for Peru and Chile, from 2005Q1 for Colombia, and from 2002Q1 for Mexico. The remaining macroeconomic variables for PA countries are available from 2004Q1 to 2019Q4.

We restrict our analysis to the pre-pandemic period, as the extraordinary magnitude and nature of the COVID-19 shock present severe econometric challenges.

3.2. Unit root test

As is common in time series analysis, we assess whether the series in our sample is stationary. To this end, we perform unit root tests considering the variables as a panel for three reasons. First, panel unit root tests offer higher statistical power by exploiting both cross-sectional and time-series dimensions, thereby reducing the risk of falsely failing to reject non-stationarity (Im et al., 2003; Levin et al., 2002). Second, these tests account for heterogeneity across countries or sectors, providing more reliable results compared to individual time-series tests (Maddala and Wu, 1999). Moreover, they help mitigate small-sample biases and allow for cross-sectional dependence, which is particularly relevant in macroeconomic analyses (Pesaran, 2007).

Table 3. Unit root test for data panel.

	Common unit root process		Individual unit root process	
Variables	Levin, Lin, and Chu	Breitung t stats	ADF-Fisher Chi-square (χ^2)	IPS
$mpr_{i,t}$	0.11	−2.18**	14.64*	−1.56*
$\bar{y}_{it}$	−0.15	−2.21**	−22.01***	−2.97***
π_{it}	−1.80**	−5.53***	37.15***	−4.80***
$\bar{y}^*{}_{it}$	−0.69	−2.81***	17.62**	−2.43***
Δy_{it}	0.87	−3.06***	21.26***	−2.67***
Δy^*_{it}	−0.72	−2.44***	15.37*	−2.00**
mpr^*_{it}	0.30	−2.60***	4.25	0.34

Note(s): The asterisks denote statistical significance at the 1% (***), 5% (**), and 10% (*) levels. All variables were tested with individual intercept and trend.

Table 4. Unit root and stationarity tests for the U.S.

	Augmented Dickey–Fuller test (ADF)			Phillips–Perron test (PP)			KPSS test	
	None	Intercept	Trend and intercept	None	Intercept	Trend and intercept	Intercept	Trend and intercept
cpi_t	3.29	3.90	−2.86	8.35	3.33	−2.63	2.00	0.43
$Log\ y_t$	7.28	−2.13	−1.61	10.29	−2.15	−1.30	2.04	0.40
Δcpi_t	−1.63*	−2.12	−2.08	−2.66***	−3.83***	−3.81***	0.32***	0.29
$\Delta Log\ y_t$	−1.51	−3.59***	−4.06***	−2.96***	−3.49***	−3.07	0.51**	0.03***
1%	−2.59	−3.51	−4.07	−2.59	−3.50	−4.06	0.74	0.22
5%	−1.94	−2.89	−3.46	−1.94	−2.89	−3.46	0.46	0.15
10%	−1.61	−2.58	−3.16	−1.61	−2.58	−3.16	0.35	0.12

Note(s): The significance levels are (*) at 90%, (**) at 95%, and (***) at 99%. The null hypothesis of the unit root tests states the presence of a unit root. For the stationarity test (KPSS), the null hypothesis indicates that the series is stationary. Data are in quarterly frequency and cover the period from 2002Q1 to 2024Q3. The optimal lag length for the ADF test was determined using the Schwarz Information Criterion (max. 14). For the PP and KPSS tests, the Bartlett kernel spectral estimation method and Newey–West bandwidth selection were applied.

Table 3 shows the results considering individual and common unit root processes. In all cases, there is statistical evidence that the process does not have a unit root. Also, for the U.S. variables included in the SVAR model, we conduct conventional unit root tests for time series: the augmented Dickey-Fuller (ADF) test, the Phillips-Perron (PP) test, and the Kwiatkowski-Phillips-Schmidt-Shin (KPSS) test. The choice of these tests is based on their respective sensitivities. The ADF and PP tests may have low power when the autoregressive parameter is close to one, potentially failing to reject the null hypothesis of a unit root even when the series is stationary. In contrast, the KPSS test is more robust in the presence of structural breaks, providing an additional perspective on stationarity. Notably, the null hypothesis in the ADF and PP tests assumes the presence of a unit root (non-stationarity), whereas in the KPSS test, the null hypothesis assumes stationarity. This distinction allows for a more comprehensive assessment of the time series properties. However, it is important to acknowledge that these tests are not entirely reliable, as their performance depends on sample size, structural breaks, and model specification. Nonetheless, they provide a useful indication of the underlying behavior of the series. The results of these tests, presented in Table 4, indicate that the U.S. variables are stationary in first differences.

Unit root test results for the domestic macroeconomic aggregates of each country are presented in Tables 11–14 (see Appendix). These results suggest that the variables, when included in the model in first differences, appear to be stationary. Given this transformation, we do not conduct additional structural break tests for these Latin American series, as differencing is generally sufficient to reduce the risk of spurious regression and simplifies the empirical strategy. This approach is common in empirical work on emerging markets, where detecting structural breaks can be particularly challenging due to limited data availability and the presence of multiple overlapping shocks, such as commodity price cycles, political instability, or global financial volatility (Del Tedesco Lins, 2025).

However, for U.S. real GDP, which is transformed using logarithms and plays a key role in the identification of global demand shocks, we perform a unit root test with one endogenous structural break. Results from the Zivot-Andrews test (Table 5) fail to reject the null hypothesis of a unit root with a single break, with the estimated break date occurring in 2006Q2. This break is economically plausible, as it anticipates the contraction associated with the 2008 Global Financial Crisis, a period of major disruption in global output, investment, and trade. To assess whether this structural break compromises the identification of shocks or the model's dynamic behavior, we conduct a robustness check based on subsample estimations. The results (see Figure 10) confirm that the system remains stable and that the dynamic responses are preserved across subsamples, reinforcing the credibility of our baseline model.

Table 5. Unit root test allowing for a structural break for the U.S.

Zivot-Andrews test				
		Intercept	Trend	Both
Log (rgdp)		−3.69	−3.73	−3.76
Critical values	1%	−5.34	−4.80	−5.57
	5%	−4.93	−4.42	−5.08
	10%	−4.58	−4.11	−4.82

Notes: *, **, and *** denote rejection of the null hypothesis at the 10%, 5%, and 1% significance levels, respectively. The structural break date is identified as 2006Q2. Sample period: 1949Q3–2019Q4. The null hypothesis assumes the presence of a unit root with a one-time endogenous structural break. The maximum lag length used is 4.

4. Results

4.1. Cyclical synchronization and volatility between the U.S. and P.A.C

This section conducts data analysis on business cycles based on simple correlations. This preliminary analysis aims to identify how important the U.S. business cycles are to PAC countries.

Table 6 shows that the correlation between the business cycles of PAC countries and the U.S. has increased over time. In all cases, the values of the correlation coefficient are higher in the most recent period (1996Q2–2024Q2) compared to the earlier subsamples, suggesting a greater co-movement of business cycles. This trend is particularly noticeable in Chile, Colombia, and Peru, where the correlation has gradually risen across subsamples.

The table also indicates that the relative volatility of PAC countries' business cycles, measured as the ratio of their standard deviation to that of the U.S., has generally increased over time. For Peru, Chile, and Colombia, this ratio is higher in the latest period compared to the earlier ones, suggesting that fluctuations in these economies have become more pronounced relative to the U.S. business cycle. Mexico, which already exhibited a high correlation in the earlier periods, has maintained this strong relationship while also showing an increase in relative volatility. These patterns suggest that the connection between PAC economies and the U.S. business cycle has evolved over time, with both correlation and relative volatility increasing in most cases.

Table 6. Correlation of the PAC countries' business cycles with the U.S. business cycle

Subsamples	(1996Q2–2009Q1)		(1996Q2–2019Q4)		(1996Q2–2024Q2)	
	$\rho(i, us)$	$\frac{\sigma_i}{\sigma_{us}}$	$\rho(i, us)$	$\frac{\sigma_i}{\sigma_{us}}$	$\rho(i, us)$	$\frac{\sigma_i}{\sigma_{us}}$
Peru	0.20	3.02	0.24	3.31	0.54	3.67
Chile	0.50	1.45	0.57	1.61	0.72	1.83
Colombia	0.36	1.45	0.48	1.40	0.68	1.92
Mexico	0.81	1.68	0.82	1.80	0.87	2.07
U.S.	1	1	1	1	1	1

Note(s): i denotes the business cycles of a specific country. U.S. corresponds to the business cycle of the U.S. ρ is the correlation coefficient, and σ is the standard deviation. Statistics were calculated for the business cycles obtained by HP filter application to the log of the real GDP series for each country.

Figure 1 presents the rolling correlation between the business cycles of the U.S. and the PAC countries. The left panel shows the correlation based on the cycle extracted from the logarithm of real GDP, while the right panel reflects the cycle derived from real GDP growth.

Between 2003 and 2013, the correlation remained elevated across most PAC countries, which may be associated with increasing financial and trade integration, as well as the spillover effects of major global events such as the 2009 financial crisis and the European sovereign debt crisis. During this period, U.S. monetary expansion and capital outflows toward emerging markets may have strengthened business cycle synchronization through financial and trade linkages.

However, from 2013 to 2019, a declining trend in correlation is observed. This reduction could reflect asymmetric shocks affecting the U.S. and PAC economies differently. The fall in commodity prices in 2016 had a significant impact on the terms of trade for PAC economies, while the U.S. economy remained relatively resilient due to strong domestic demand. Likewise, the trade tensions between the United States and China in 2018 may have indirectly affected PAC economies through China, which is a key trading partner for these countries, thereby weakening their synchronization with the U.S. business cycle. Additionally, monetary policy divergence—where the Federal Reserve gradually tightened its stance while PAC central banks maintained more accommodative policies—along with stronger macroprudential measures in emerging markets, may have further contributed to the observed desynchronization.

Following the COVID-19 pandemic, the correlation rises sharply and remains at high levels. While the initial shock was global and not exclusive to the U.S., the sustained high correlation in the post-pandemic period suggests that U.S. macroeconomic conditions continue to play a growing role in shaping Latin American business cycles. This renewed alignment may reflect both the global economic recovery and the increasing influence of U.S. monetary and financial conditions on emerging markets.

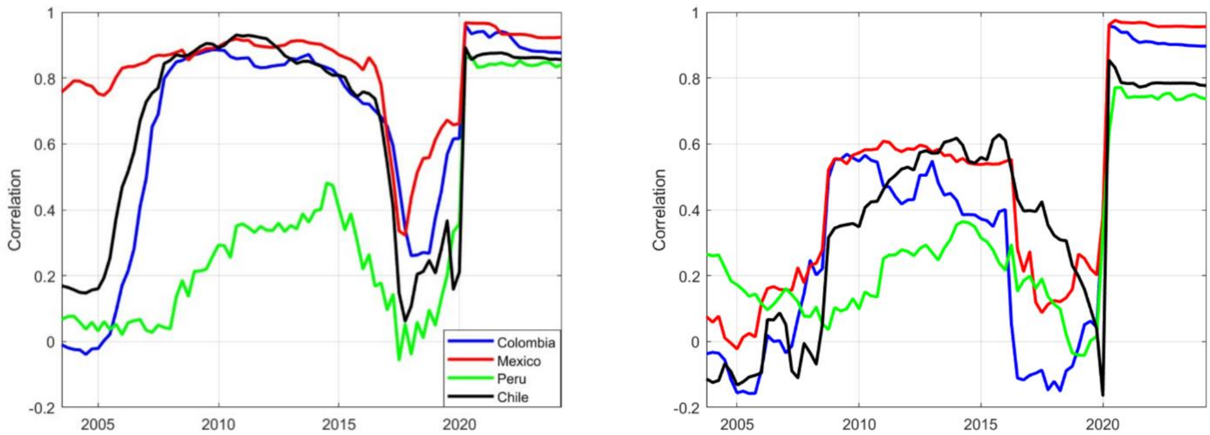

Figure 1. Rolling correlation of U.S. and PAC business cycles. The estimates were made using a 30-quarter window. The business cycles for all countries were extracted using the HP filter. The left graph presents the business cycle obtained by applying the HP filter to the logarithm of real GDP, while the right graph shows the cycle based on real GDP growth.

4.2. Monetary policy reaction function

Table 7 presents the estimation results of a Taylor-type rule using OLS panel regressions for the Pacific Alliance countries. The findings reveal a consistent pattern of interest rate smoothing behavior, as indicated by the highly significant and sizeable coefficients on the lagged policy rate, which range from 0.81 to 0.85 across all specifications. This suggests that central banks in the region adjust their policy rates gradually over time, in line with standard monetary policy practices in emerging markets.

Inflation appears as a key determinant of the policy rate, with positive and statistically significant coefficients throughout the different model variations. This highlights that price stability remains a central objective for monetary authorities. Domestic measures of economic activity, such as output growth and the output gap, also exhibit significant effects on interest rate decisions, indicating responsiveness to the domestic business cycle.

Importantly, external variables related to U.S. economic conditions also enter the Taylor rule with statistically significant coefficients. Both the U.S. output gap and U.S. output growth have a positive effect on policy rates in the PAC, reflecting how global demand shocks influence monetary policy in the region. Moreover, the U.S. policy rate is positively and significantly associated with domestic policy rates, underscoring the importance of international financial conditions, particularly U.S. monetary policy, as a key external driver of domestic policy rate setting.

Taken together, the results indicate that central banks in the Pacific Alliance conduct monetary policy by responding not only to inflation and domestic economic fluctuations but also to global macroeconomic developments, particularly those originating in the United States. The high R-squared values across specifications (all equal to 0.94) suggest that the model captures a substantial portion of the variation in interest rate movements, reinforcing the empirical relevance of the included variables and the robustness of the estimated Taylor-type rule. To further validate these findings, robustness checks were conducted (see Tables 9 and 10), which confirm the overall stability of the model and the persistence of the main empirical patterns across different specifications.

Table 7. Estimated Taylor rule with OLS panel data for PAC.

Dependent variable	r_{it}			
	(1)	(2)	(3)	(4)
Intercept	0.43***	−0.67***	0.39	−0.30***
	(0.10)	(0.11)	(0.10)	(0.11)
r_{it-1}	0.85***	0.85***	0.83***	0.81***
	(0.02)	(0.02)	(0.02)	(0.03)
π_{it}	0.28***	0.17***	0.30***	0.17***
	(0.05)	(0.03)	(0.05)	(0.03)
Δy_t	-	0.10***	-	0.12***
		(0.01)		(0.01)
Δy_t^*	-	0.10***	-	-
		(0.02)		
$\bar{y}_{it}$	0.16***	-	0.17***	-
	(0.03)		(0.02)	
$\bar{y}^*{}_{it}$	0.06*	-	-	-
	(0.03)			
r_{it}^*	-	-	0.07***	0.08***
			(0.02)	(0.02)
R-squared	0.94	0.94	0.94	0.94
SE regression	0.53	0.52	0.52	0.53
F-statistic	511.86***	562.01***	535.27***	513.16***

Notes: Asterisks denote statistical significance at the 1% (*), 5% (**), and 10% (***) levels. The table reports the OLS estimation of the Taylor rule with balanced data panel for PAC. Estimates obtained using generalized least squares (GLS) with cross-section weights to correct for heteroskedasticity. These regressions feature country-fixed effects. $\bar{y}_{it}$ denotes the output gap, while Δy_{it} refers to real GDP growth for each country. $\bar{y}^*{}_{it}$, Δy_{it}^*, and r_{it}^* correspond to the U.S. output gap, real GDP growth, and monetary policy rate, respectively. The data spans from 2004Q1 to 2019Q4.

4.3. Quantifying the evolution of the demand and supply shocks of the U.S.

The impulse response functions shown in Figure 2 represent the responses of U.S. GDP and inflation to structural supply and demand shocks. These results align well with macroeconomic theory, which validates our identification strategy as theoretically sound.

For the response of GDP to structural shocks, we see that a positive supply shock leads to a sustained increase in GDP. The impact peaks between the fourth and eighth periods, with GDP rising by approximately 1.5% before gradually tapering off but remaining positive. This behavior is consistent with the theoretical expectation that supply shocks, such as technological innovations or decreases in production costs, enhance output by reducing marginal costs, thereby stimulating production without directly exerting inflationary pressures. The prolonged positive effect on GDP indicates that supply shocks generate lasting improvements in productive capacity, which is consistent with neoclassical growth theory.

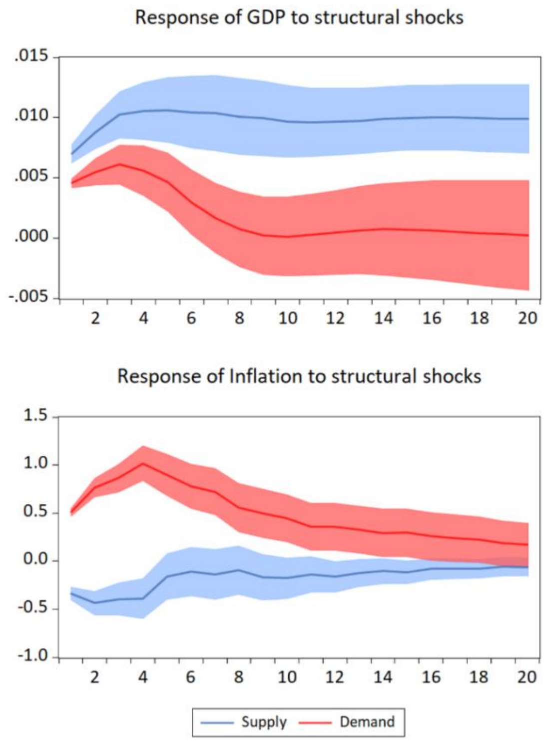

Figure 2. Impulse response functions of the U.S. economy. The graph shows the responses to 1 SD of supply and demand shocks. The confidence interval was constructed with asymptotic simulations at 95% confidence.

In contrast, a positive demand shock generates an initial increase in GDP, peaking around 0.5%, but this effect quickly dissipates after 6 periods. The temporary nature of the demand-driven GDP response is consistent with the idea that demand shocks, while boosting short-term output, do not permanently affect real GDP in the long run due to supply constraints and the economy's return to its potential output. This rapid decay in the effect of demand shocks reflects the standard assumption that demand policies, such as fiscal stimulus or monetary expansion, have short-lived impacts on output as the economy reverts to equilibrium.

The response of inflation to structural shocks further reinforces the coherence of our identification approach. Supply shocks exhibit a deflationary effect, as expected, with inflation decreasing slightly for the first few periods. This reflects the fact that positive supply shocks lower production costs and exert downward pressure on prices. Over time, the effect diminishes, but inflation remains slightly below zero, suggesting that supply-side improvements keep inflation contained without triggering strong inflationary pressures.

On the other hand, the inflationary response to demand shocks is much more pronounced. A positive demand shock leads to a sharp rise in inflation, peaking around 1% before gradually decreasing over time. This result aligns with the theoretical expectation that demand shocks, which increase aggregate demand relative to supply, exert upward pressure on prices. The peak inflationary response occurs early, reflecting the rapid transmission of demand-side pressures into price levels. The gradual reduction in inflation following the peak also makes theoretical sense, as the economy absorbs the demand shock and returns to its long-term potential.

In sum, the dynamics of GDP and inflation in response to supply and demand shocks are entirely consistent with established economic theory. Supply shocks lead to sustained increases in output and slight deflationary pressure, while demand shocks generate temporary output increases and more immediate inflationary effects. These results provide a solid theoretical foundation for our identification strategy, confirming that it captures the expected macroeconomic responses to these types of shocks.

In Figure 3, showing the decomposition of U.S. structural supply and demand shocks, several significant historical events stand out, corresponding to well-known macroeconomic disruptions in U.S. economic history. The shocks offer insight into the underlying causes of various periods of economic instability.

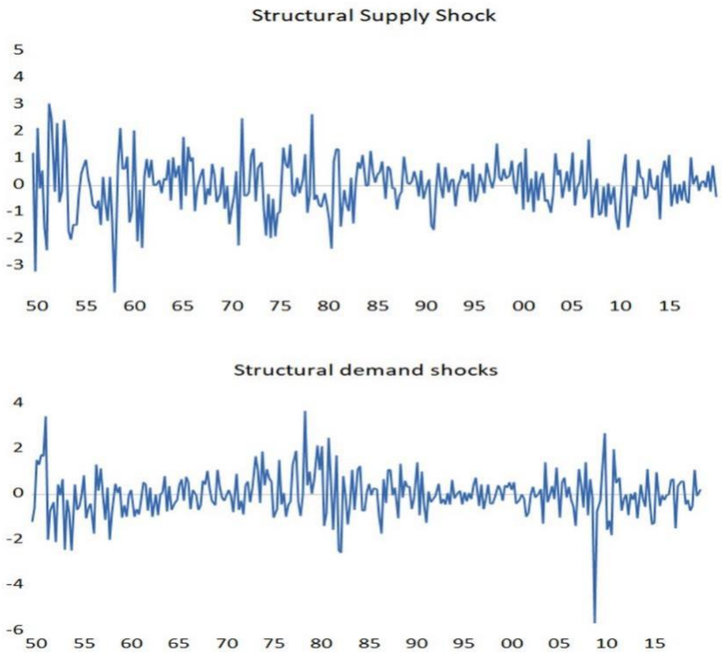

Figure 3. Decomposition of U.S. structural shocks. Historical evolution of U.S. structural shocks (1950Q1–2019Q4).

In the early 1950s, both supply and demand shocks show high volatility, likely reflecting the post-World War II adjustment period and the Korean War (1950–1953). This period was marked by shifts in production capacity and government spending, which would explain the larger supply shocks.

Similarly, the spikes in demand shocks around the 1950s and 1960s may be linked to the strong post-war consumer spending and the boom associated with rebuilding. The 1970s are a prominent period where supply shocks dominate. The large positive and negative spikes align with the oil crises of 1973 and 1979. These were classic supply-side disruptions where sharp increases in oil prices led to higher production costs across industries, resulting in inflation and stagflation, a scenario where inflation rises while output growth slows or contracts. During the Volcker disinflation in the early 1980s, supply shocks show notable fluctuation. This period saw the Federal Reserve aggressively raising interest rates to combat inflation, which had built up throughout the 1970s. This led to recessions in the early 1980s, as reflected in the volatility of the supply shocks.

The most striking feature in the demand shocks is the significant negative spike in 2008, corresponding to the Global Financial Crisis. This was primarily a demand-driven crisis, where collapsing financial markets and falling consumption and investment led to a sharp contraction in output. The extreme negative demand shock during this period indicates a rapid fall in aggregate demand as financial conditions worsened globally. Following the financial crisis, there is evidence of a recovery in demand, but still considerable volatility. The demand shock graph shows that the recovery was slow and uncertain, potentially reflecting the uneven pace of recovery and the lagging effects of fiscal and monetary interventions that sought to restore stability. During this time, supply shocks remained relatively stable, but demand shocks fluctuated, which could be associated with ongoing adjustments after the 2008 crisis, as well as other global developments such as the European debt crisis. The relatively low supply shocks reflect more stable production conditions, but demand-side fluctuations highlight the lingering effects of global demand disruptions and the Fed's accommodative monetary policies during this period.

The decomposition of supply and demand shocks reveals a clear relationship between historical events and their impact on the U.S. economy. Supply shocks have been especially dominant during periods of external disruptions like oil crises, while demand shocks are more evident during financial crises and their aftermath. Figure 4 supports that evidence and shows that, in general, the U.S. business cycle has been predominantly driven by supply shocks.

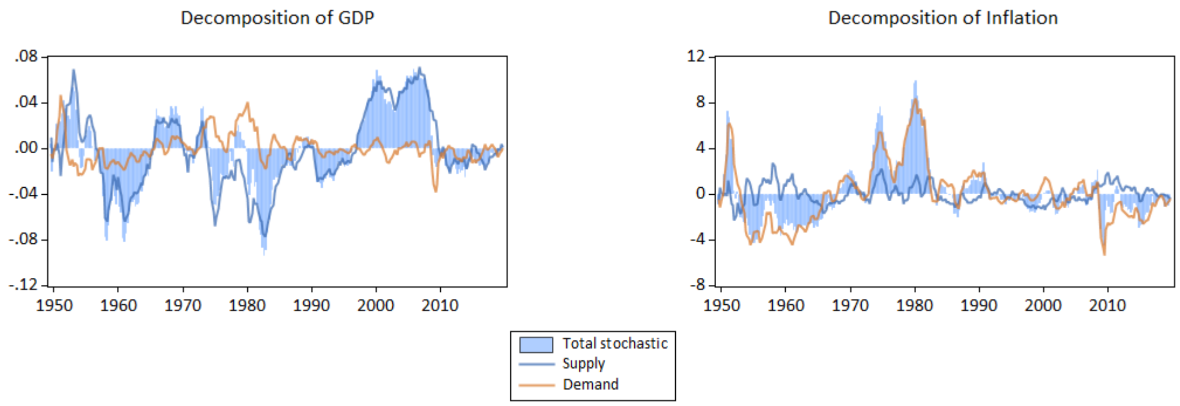

Figure 4. Historical decomposition of U.S. shocks.

4.4. Dynamic response of monetary policy rates to U.S. structural shocks

The impulse response functions (IRFs) in Figure 5 illustrate how the monetary policy rates of Peru, Colombia, Chile, and Mexico respond to a structural supply shock originating from the United States. These functions are represented with the solid blue lines, while the red dashed lines indicate the

95% confidence intervals. Below is a detailed analysis of the magnitudes, dynamics, persistence, and cross-country comparisons. Figure 8 presents the impulse response functions for each country, illustrating the transmission mechanisms following a U.S. supply shock.

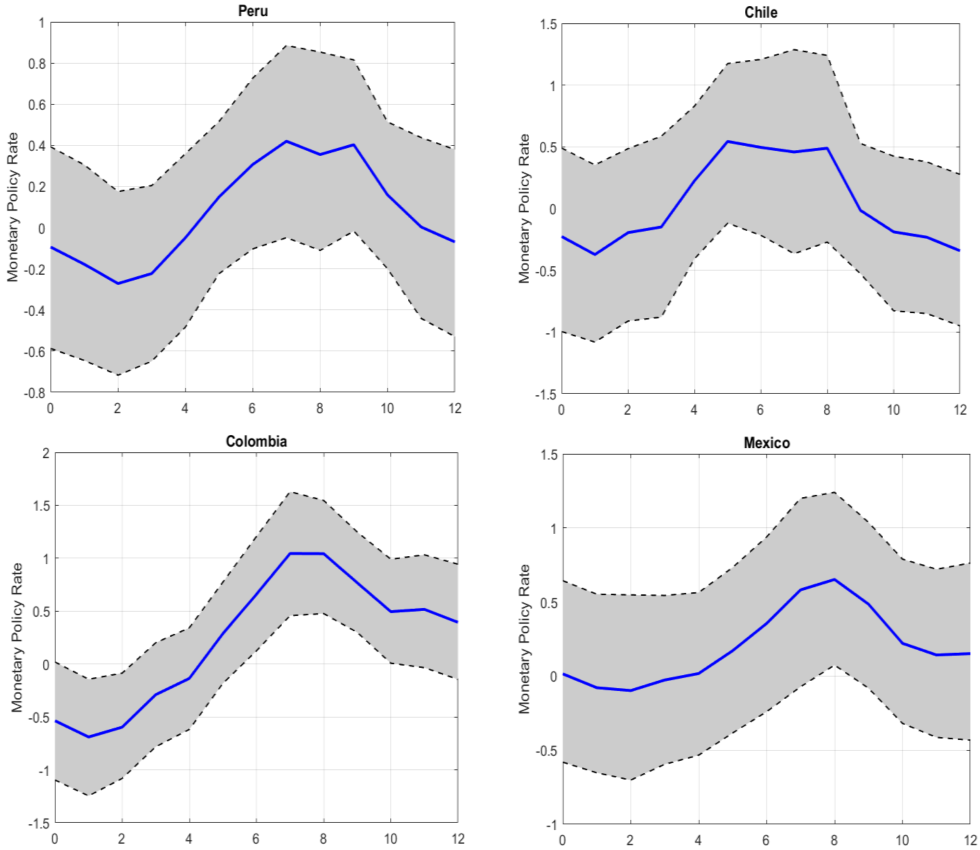

Figure 5. IRF of monetary policy rates to U.S. structural supply shocks. The solid blue line represents the estimated impulse response function (IRF), while the shaded area denotes the confidence bands. Confidence intervals were constructed using the wild bootstrap with 2000 replications, employing the Rademacher distribution, following the methodology of Gonçalves and Kilian (2004). The confidence bands correspond to the 2.5th and 97.5th percentiles of the bootstrapped distribution. The x-axis indicates the horizon in quarters, while the y-axis shows the response of the monetary policy rate to the structural shock.

In Peru, the impulse response function of the monetary policy rate to a U.S. supply shock remains initially slightly negative, stabilizing around zero in the first few periods. From the fourth period onward, the response turns positive, peaking around the seventh period at approximately 0.3%. However, this reaction is not statistically significant at conventional levels throughout the horizon, indicating an overall acyclical monetary policy response. As shown in Figure 8, this result aligns with the transmission patterns observed across key macroeconomic variables. Real GDP does not exhibit a significant reaction, and the nominal exchange rate also remains largely unaffected, with confidence bands encompassing zero at all horizons. The inflation rate, however, falls significantly on impact, yet the effect fades rapidly and becomes statistically insignificant by the third period. Similarly, the EMBI spread shows no significant response across the impulse-response horizon. These dynamics suggest that the U.S. supply shock exerts limited influence on the Peruvian economy through both real and financial channels. The absence of persistent movements in inflation, output, or the exchange rate helps explain the lack of an active monetary policy adjustment, as the external shock fails to generate sustained macroeconomic imbalances that would warrant a policy response.

In Colombia, the monetary policy rate responds to a U.S. supply shock by initially decreasing by about 0.5% during the first two quarters, before reversing direction and increasing significantly, peaking at approximately 1.5% in the seventh period. This procyclical response aligns with the dilemma central banks face when dealing with supply shocks, as noted by Végh et al. (2017) and Ocampo and Ojeda-Joya (2022). The broader dynamics captured in Figure 8 show that output and the nominal exchange rate exhibit no significant reaction throughout the IRF horizon, with confidence bands including zero. In contrast, inflation falls significantly on impact, potentially creating space for initial monetary easing. However, by the fifth period, inflation begins to rise, which could explain the subsequent tightening of the policy rate. This sequence suggests that the Colombian central bank initially adopts an accommodative stance in response to disinflationary pressures but later adjusts rates upward to contain emerging inflation risks. Additionally, the EMBI declines significantly on impact, suggesting a temporary improvement in perceived sovereign risk or investor sentiment. This aligns with the observed stability of the nominal exchange rate, which likely helped ease external pressures and allowed for an initially accommodative monetary stance. However, from the seventh period onward, the EMBI turns positive, indicating a deterioration in external financing conditions or a reassessment of sovereign risk. This shift may coincide with the rise in inflation observed around the same time, potentially prompting a tighter monetary policy response by the central bank.

For Chile, the IRF indicates that the monetary policy rate initially declines by approximately 0.5%, similar to the pattern observed in Colombia. The rate then rises steadily, peaking around the sixth period at approximately 1%. Although this response is smaller in magnitude than Colombia's, the IRFs are not statistically significant in all periods. Importantly, the broader transmission mechanisms captured in the model show that real GDP, the nominal exchange rate, EMBI, and inflation do not exhibit statistically significant responses to the U.S. supply shock at any point in the horizon. This suggests that the lack of a consistent reaction in domestic macroeconomic variables may help explain the relatively muted and acyclical policy rate response, possibly reflecting a non-accommodative stance by the Chilean central bank.

In Mexico, the IRF indicates a very mild initial decline in the monetary policy rate, followed by a gradual increase that peaks at approximately 0.6% around the seventh period. Although the overall magnitude of the response is smaller than that observed in Colombia and Chile, the general pattern is consistent across the three countries. The monetary policy reaction can be characterized as acyclical in the short run, as the IRFs are not statistically significant throughout most periods. This is consistent with the absence of significant responses in output and the nominal exchange rate. However, inflation falls on impact and subsequently turns positive in the fifth and sixth periods, which may explain the upward adjustment in the policy rate during that phase. The reaction becomes statistically significant in the eighth period, consistent with the lagged rise in inflation, suggesting that the central bank reacts with some delay to these emerging pressures. Additionally, the EMBI becomes slightly positive between the fifth and seventh periods before declining, indicating a temporary rise in perceived sovereign risk.

Figure 6 depicts how the monetary policy rates of Peru, Colombia, Chile, and Mexico respond to a structural aggregate demand shock from the United States. As with the previous IRFs, the solid blue lines represent the estimated response of the policy rates, while the red dashed lines show the 95% confidence intervals. Below is an analysis of the magnitude, dynamics, persistence, and cross-country comparisons. Figure 9 presents the impulse response functions for each country, illustrating the transmission mechanisms following a U.S. supply shock.

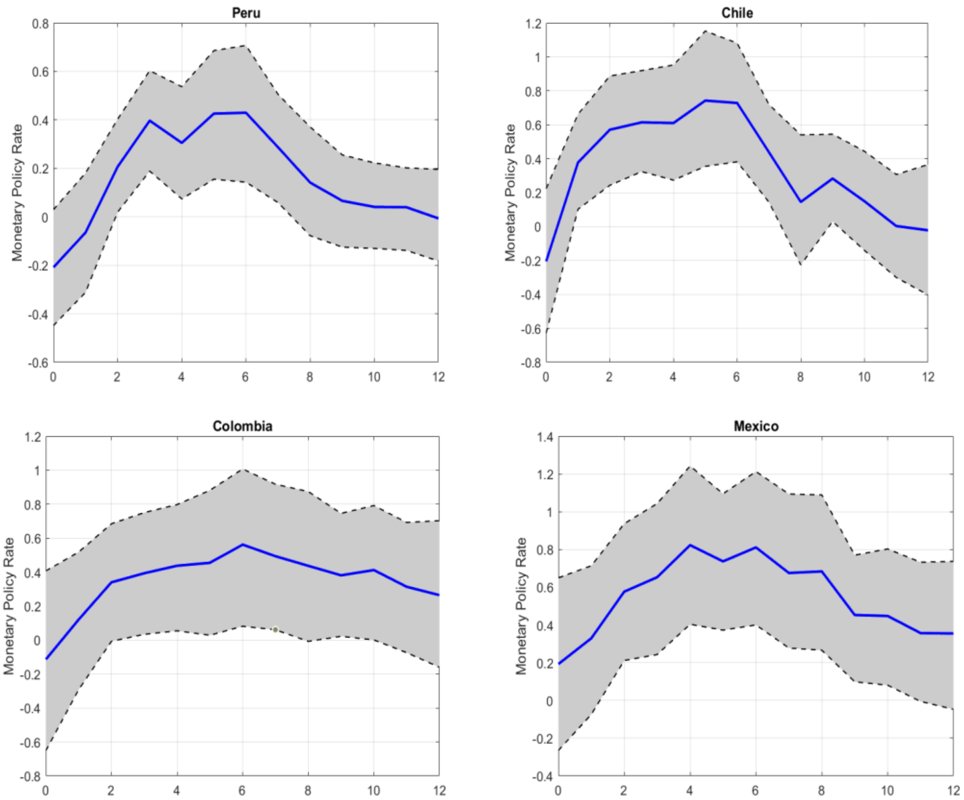

Figure 6. IRF of monetary policy rates to U.S. structural demand shocks. The solid blue line represents the estimated impulse response function (IRF), while the shaded area denotes the confidence bands. Confidence intervals were constructed using the wild bootstrap with 2000 replications, employing the Rademacher distribution, following the methodology of Gonçalves and Kilian (2004). The confidence bands correspond to the 2.5th and 97.5th percentiles of the bootstrapped distribution. The x-axis indicates the horizon in quarters, while the y-axis shows the response of the monetary policy rate to the structural shock.

In Peru, the monetary policy rate exhibits a positive and moderate response to the U.S. demand shock. While the initial response is slightly negative, it is not statistically significant. The policy rate begins to rise steadily from the second period, reaching a peak of approximately 0.45% in the fourth period, and then gradually declines while remaining above zero for the rest of the forecast horizon.

This trajectory suggests a persistent, though mild, tightening stance by the central bank. Notably, inflation does not exhibit a statistically significant response, and the EMBI remains largely unaffected, indicating limited upward pressure on prices and perceived sovereign risk. However, a mild appreciation of the nominal exchange rate is observed, potentially reflecting improved terms of trade or capital inflows driven by the external shock. Output displays a brief positive reaction, consistent with a temporary boost in activity. Altogether, the monetary tightening appears precautionary, possibly aimed at containing potential inflationary pressures or maintaining macroeconomic stability in light of currency movements, rather than responding to immediate domestic imbalances.

In Colombia, the monetary policy rate exhibits a steady increase, peaking at approximately 0.55% around the fifth period and remaining above zero throughout the 12-period horizon. This indicates a moderately persistent tightening stance in response to the U.S. demand shock. The output response is positive and statistically significant on impact, suggesting an immediate boost in economic activity. However, this effect fades quickly, becoming statistically insignificant in subsequent periods. Around the eleventh period, output registers a significant decline before returning to non-significant levels, pointing to a delayed contraction possibly reflecting external demand normalization or policy tightening effects. Inflation does not display a significant response across the horizon, and the EMBI remains broadly unchanged, indicating stable inflation expectations and sovereign risk perceptions. The exchange rate rises sharply on impact, although this effect dissipates by the second period. The observed monetary tightening may reflect a precautionary response by the central bank to contain potential overheating or second-round effects, despite the absence of sustained inflationary pressures.

In Chile, the monetary policy rate exhibits a delayed but strong response to the U.S. demand shock, beginning to rise after the first period and peaking at approximately 0.8% around the fifth period. This timing suggests that the Chilean central bank does not react immediately, possibly reflecting a wait-and-see approach to assess the persistence of external conditions before tightening. After reaching its peak, the policy rate declines gradually and becomes statistically insignificant by the eighth period. Inflation rises significantly on impact, consistent with the initial tightening, but becomes statistically insignificant from the second period onward. Output does not display any significant response throughout the impulse-response horizon, suggesting limited real activity spillovers. The nominal exchange rate rises significantly on impact and up to the second period, after which the response loses statistical significance. Meanwhile, the EMBI falls in the initial periods (up to around the fifth), signaling a temporary improvement in perceived sovereign risk. Overall, the response reflects a flexible and reactive monetary stance, with a short lag and temporary tightening to address immediate inflationary pressures, followed by normalization as external effects fade.

In Mexico, the monetary policy rate increases to around 0.8% by the fifth period in response to the U.S. demand shock. The magnitude is smaller than in Chile but larger than in Peru and Colombia, indicating that Mexico's central bank takes a moderate but decisive stance in response to external demand shocks. Like the other countries, the response is persistent, although the rate starts to decline after the fifth period. The output shows a significant increase in impact, but the effect becomes statistically insignificant by the second period, suggesting only a short-lived boost in real activity. Inflation, in contrast, falls significantly on impact before turning insignificant in the following period, possibly reflecting temporary disinflationary effects or a terms-of-trade improvement. The nominal exchange rate rises sharply in the first period, with the effect losing significance after the second period. Meanwhile, the EMBI does not display any significant response, suggesting no major shifts in sovereign risk perception.

These results can be summarized as follows: The responses of monetary policy rates to U.S. supply and demand shocks are quite similar across Peru, Colombia, Chile, and Mexico, reflecting the similarity in their economic structures and policy frameworks. Peru, Chile, and Mexico exhibit acyclical responses to U.S. supply shocks, suggesting limited monetary policy adjustments due to offsetting exchange rate effects and non-accommodative stances, while Colombia reacts more aggressively. Regarding demand shocks, all countries tighten monetary policy, with Chile and Mexico displaying similar peak magnitudes. However, Mexico's response is more persistent, while Chile declines more sharply after the peak. These differences highlight varying degrees of exposure and policy responsiveness to external shocks.

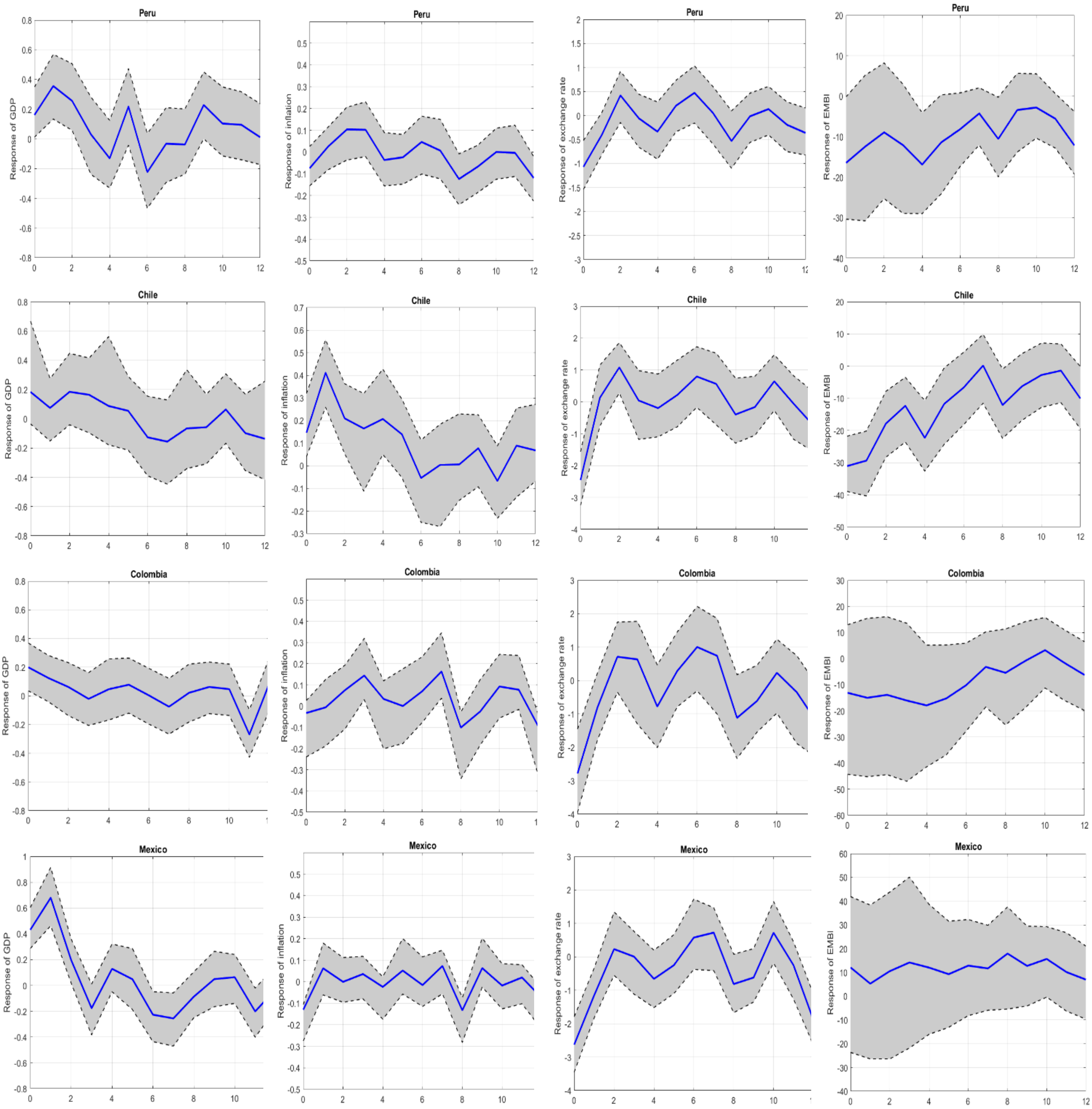

Figure 7. IRF of macroaggregates to U.S. structural demand shocks. The solid blue line represents the estimated impulse response function (IRF), while the shaded area denotes the confidence bands. Confidence intervals were constructed using the wild bootstrap with 2000 replications, employing the Rademacher distribution, following the methodology of Gonçalves and Kilian (2004). The confidence bands correspond to the 2.5^{th} and 97.5^{th} percentiles of the bootstrapped distribution. The x-axis indicates the horizon in quarters, while the y-axis shows the response of the monetary policy rate to the structural shock.

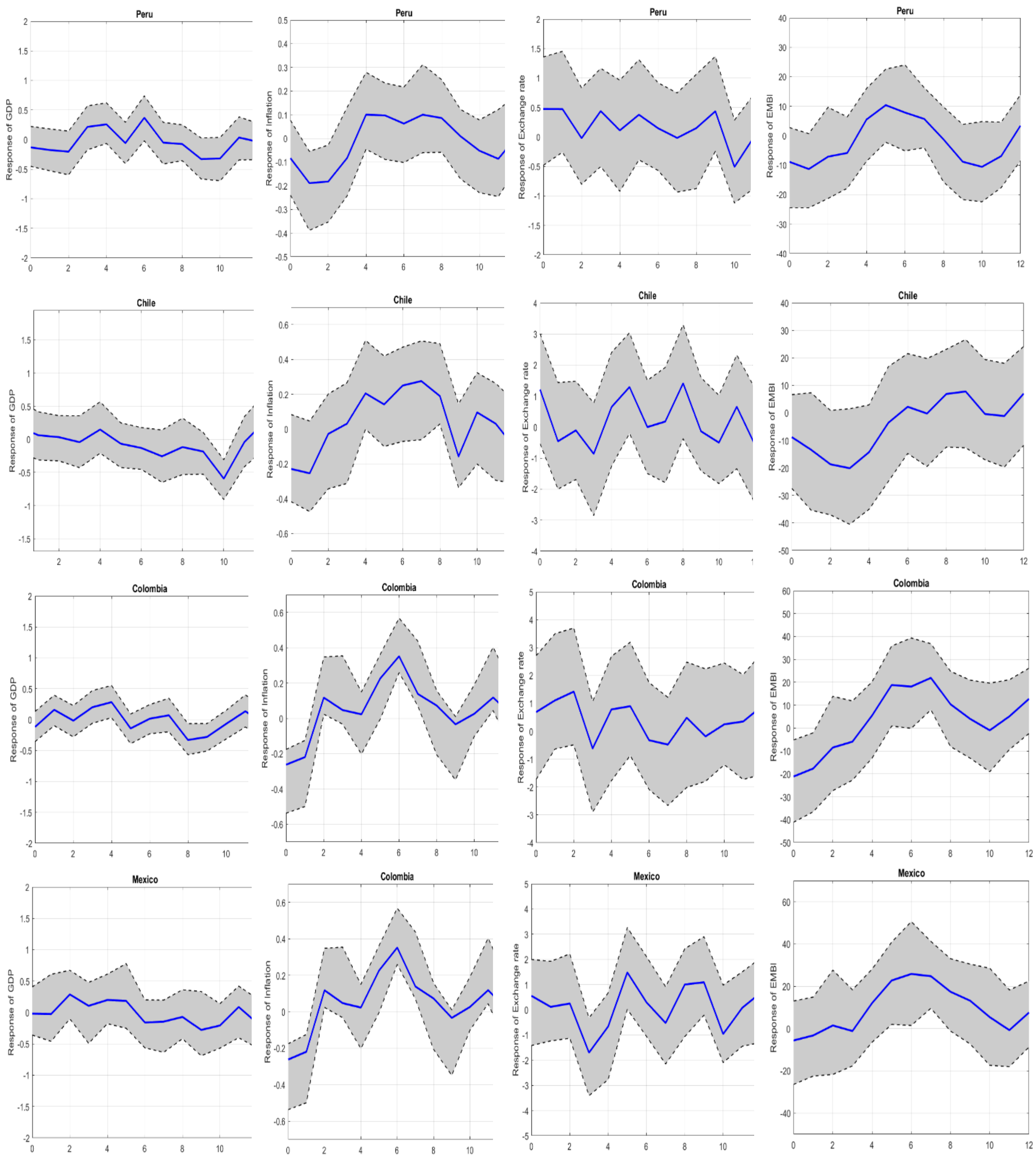

Figure 8. IRF of macroaggregates to U.S. structural supply shocks. The solid blue line represents the estimated impulse response function (IRF), while the shaded area denotes the confidence bands. Confidence intervals were constructed using the wild bootstrap with 2000 replications, employing the Rademacher distribution, following the methodology of Gonçalves and Kilian (2004). The confidence bands correspond to the 2.5th and 97.5th percentiles of the bootstrapped distribution. The x-axis indicates the horizon in quarters, while the y-axis shows the response of the monetary policy rate to the structural shock.

5. Discussion

This study contributes to the literature on international monetary transmission by showing that U.S. structural shocks, both supply and demand, stem from endogenous dynamics rather than being purely exogenous forces for emerging markets. Using a structural VAR framework with long-run restrictions, we identify these shocks and trace their effects across Pacific Alliance countries. The findings challenge traditional models that treat foreign variables as exogenous, with the observed variation in response persistence underscoring the importance of incorporating country-specific features such as trade exposure, financial openness, and exchange-rate regimes into open-economy frameworks. Furthermore, our results suggest that the supply shock dilemma posed by Végh et al. (2017), wherein monetary authorities face a trade-off between stabilizing output and inflation, does not appear to hold for most Pacific Alliance countries, with the notable exception of Colombia. This finding raises questions about the generalizability of the dilemma and suggests that certain institutional frameworks, such as inflation targeting with flexible exchange rates, may mitigate its relevance.

From a policy perspective, our results suggest that central banks should distinguish between the origin and type of U.S. shocks when setting interest rates. Countries highly exposed to U.S. demand shocks, such as Mexico, may require stronger buffers or macroprudential tools to safeguard monetary autonomy. Conversely, the muted reaction to supply shocks in Chile, Colombia, and Peru supports the stabilizing role of inflation-targeting regimes with flexible exchange rates. These institutional arrangements appear effective in anchoring expectations and limiting volatility.

Nonetheless, this study faces certain limitations. Although the identification strategy relies on long-run restrictions—which can be sensitive to alternative model specifications—we conducted a robustness check that supports the validity of our results. Additionally, our analysis focuses solely on the monetary policy rate, omitting other potential transmission channels such as financial markets or fiscal policy responses. Future research could address these limitations by implementing alternative identification schemes such as sign-restricted VARs, incorporating high-frequency data, exploring nonlinear dynamics or regime changes, and extending the analysis to a broader set of emerging economies to assess the external validity of our findings.

6. Robustness check

6.1. Sign restriction for the U.S. model

To evaluate the results obtained from the impulse response functions of the VAR model for the United States (Figure 2), we implemented the following robustness strategies. First, we applied different transformations to the variables; in particular, both the CPI and real GDP were expressed as annualized quarterly growth rates. Second, we reduced the number of lags in the VAR specification to two, in order to test the sensitivity of the results to lag selection. Finally, structural shocks were identified using the sign restrictions approach, for which we employed the following Givens rotation matrix:

$$Q = \begin{bmatrix} \cos\theta & -\sin\theta \\ \sin\theta & \cos\theta \end{bmatrix}$$

The sign restrictions are presented in Table 8. Overall, the results suggest that the findings are robust to alternative identification strategies, lag length choices, and variable transformations.

Table 8. Sign restrictions

Variable/shock	Supply	Demand
GDP	+	+
Inflation	-	+

Figure 9 shows that the responses of the macroeconomic variables to structural shocks remain consistent in terms of sign, persistence, and significance. In particular, supply shocks have a positive impact on GDP, and their influence persists over time. Likewise, demand shocks generate positive but short-lived effects on output, with the impact dissipating around the fifth quarter. Inflation responds negatively to supply shocks on impact, with the effect fading by approximately the fifteenth quarter. Demand shocks, in turn, lead to a positive and highly persistent response in inflation in the short term. The persistence and direction of these responses are consistent with the baseline model, indicating that the system's dynamics are preserved and that the structural shocks are properly identified.

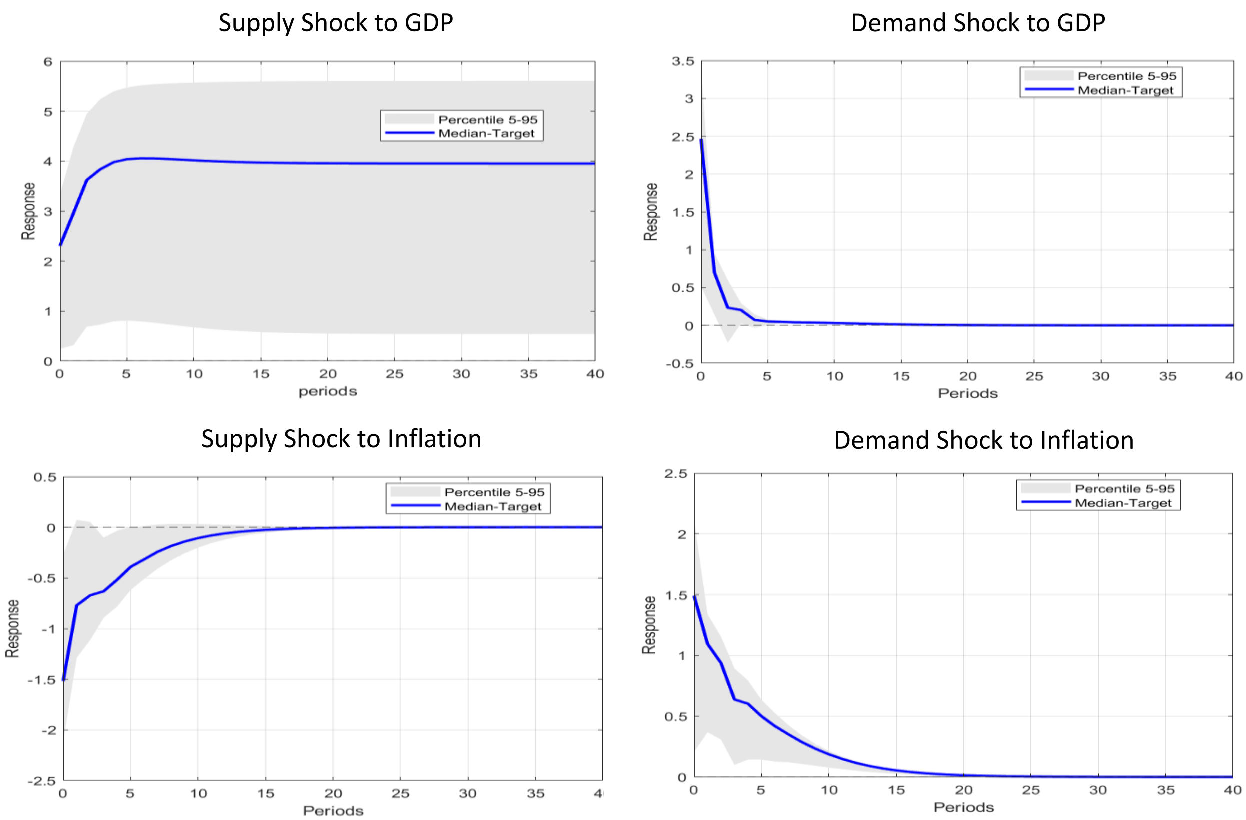

Figure 9. Impulse response functions (IRFs) for the U.S. under a sign-restriction identification scheme. Based on 10,000 simulations, 47.13% of the draws satisfy the imposed sign restrictions. The median-target approach follows Fry and Pagan (2011). Reported are the median responses along with the 5th and 95th percentiles from the accepted draws.

6.2. Taylor rule

As an additional robustness exercise, we estimate a forward-looking version of the monetary policy rule using a regression framework. In the first specification (Table 9), we include one-period-ahead inflation π_{t+1} as a proxy for rational expectations, replacing the contemporaneous or lagged inflation term used in the baseline model. This approach allows us to assess whether the estimated response of the policy interest rate is sensitive to the assumption of forward-looking behavior by monetary authorities.

In a second specification (Table 10), we extend the forward-looking component by including expectations of key domestic macroeconomic variables: inflation, real GDP growth, and the output gap. These expectations are constructed using their respective values at $t+1$, assuming rational expectations. External variables from the U.S. remain included to control for global shocks.

Table 9. Estimated Taylor rule with forward-looking inflation.

Dependent variable	r_{it}			
	(1)	(2)	(3)	(4)
intercept	0.43***	−0.67***	0.41***	−0.34***
	(0.11)	(0.11)	(0.11)	(0.11)
r_{it-1}	0.88***	0.85***	0.86***	0.80***
	(0.02)	(0.02)	(0.02)	(0.02)
π_{it+1}	0.14***	0.20***	0.15***	0.22***
	(0.05)	(0.03)	(0.05)	(0.03)
Δy_t	-	0.08***	-	0.10***
		(0.01)		(0.01)
Δy_t^*	-	0.09***	-	-
		(0.02)		
$\bar{y}_{it}$	0.18***	-	0.19***	-
	(0.03)		(0.02)	
$\bar{y}^*{}_{it}$	0.04	-	-	-
	(0.03)			
r_{it}^*	-	-	0.06***	0.08***
			(0.02)	(0.02)
R-squared	0.94	0.94	0.94	0.94
SE regression	0.53	0.52	0.52	0.53
F-statistic	511.86***	562.01***	535.27***	513.16***

Notes: Asterisks denote statistical significance at the 1% (*), 5% (**), and 10% (***) levels. The table reports the OLS estimation of the Taylor rule with balanced data panel for PAC. Estimates obtained using generalized least squares (GLS) with cross-section weights to correct for heteroskedasticity. These regressions feature country-fixed effects. $\bar{y}_{it}$ corresponds to the output gap of each country. $\bar{y}^*{}_{it}$ and r_{it}^* are the output gap and monetary policy rate of U.S., respectively. The data spans from 2004Q1 to 2019Q4.

Across both specifications, the main qualitative results hold. The estimated responses of the policy rate to macroeconomic fundamentals remain consistent in sign, magnitude, and statistical significance. These results reinforce the baseline findings and hold under an alternative specification.

Table 10. Estimated forward-looking Taylor rule.

Dependent variable		r_{it}		
	(1)	(2)	(3)	(4)
intercept	0.34***	−0.75***	0.27**	−0.36***
	(0.11)	(0.12)	(0.11)	(0.12)
r_{it-1}	0.90***	0.86***	0.89***	0.80***
	(0.02)	(0.02)	(0.02)	(0.03)
π_{it+1}	0.15***	0.22***	0.17***	0.24***
	(0.05)	(0.02)	(0.05)	(0.03)
Δy_{t+1}	-	0.06***	-	0.08***
		(0.01)		(0.01)
Δy^*_{t+1}	-	0.10***	-	-
		(0.01)		
$\bar{y}_{it+1}$	0.12***	-	0.14***	-
	(0.03)		(0.03)	
$\bar{y}^*{}_{it}$	0.07**	-	-	-
	(0.03)			
r^*_{it}	-	-	0.05**	0.09***
			(0.02)	(0.02)
R-squared	0.94	0.94	0.94	0.94
SE regression	0.53	0.52	0.52	0.53
F-statistic	511.86***	562.01***	535.27***	513.16***

Notes: Asterisks denote statistical significance at the 1% (*), 5% (**), and 10% (***) levels. The table reports the OLS estimation of the Taylor rule with balanced data panel for PAC. Estimates obtained using generalized least squares (GLS) with cross-section weights to correct for heteroskedasticity. These regressions feature country-fixed effects. $\bar{y}_{it}$ corresponds to the output gap of each country. $\bar{y}^*{}_{it}$ and r^*_{it} are the output gap and monetary policy rate of the U.S., respectively. The data spans from 2004Q1 to 2019Q4.

6.3. Stability of the VAR model

To assess the robustness of the estimated VAR model, we first verify the system's stability by inspecting the companion matrix's characteristic roots. As shown in Figure 10, all roots lie within the unit circle in the complex plane, confirming the dynamic stability of the model over the full sample.

Next, we examine whether the impulse response functions (IRFs) are sensitive to potential structural breaks in the data. Specifically, we split the sample based on a break date identified by the Zivot-Andrews test applied to U.S. real GDP, which points to a structural break around 2006, consistent with the global financial crisis. We estimate the IRFs over two subsamples: 1943q3–2006q1 and 2006q2–2019q4.

As shown in Figure 11, the IRFs across the two subsamples remain qualitatively similar in terms of sign and direction, suggesting that the transmission mechanism is broadly preserved. However, as expected, confidence intervals are wider in the subsample estimates due to the reduced number of observations.

The only noticeable difference is in the response to inflation shocks: while the sign of the response is consistent across subsamples, the persistence is notably lower in the second period

(2006Q2–2019Q4). This decline in persistence may reflect changes in the monetary policy stance of the Federal Reserve following the financial crisis, including a shift toward unconventional policy tools and greater transparency, which could have altered market expectations and the propagation of inflation shocks.

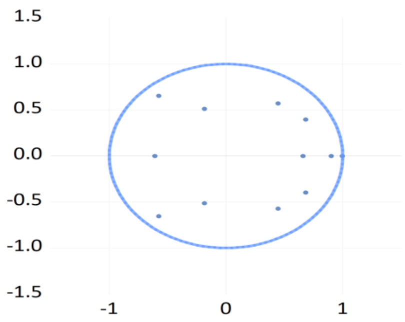

Figure 10. Inverse roots of the characteristic polynomial.

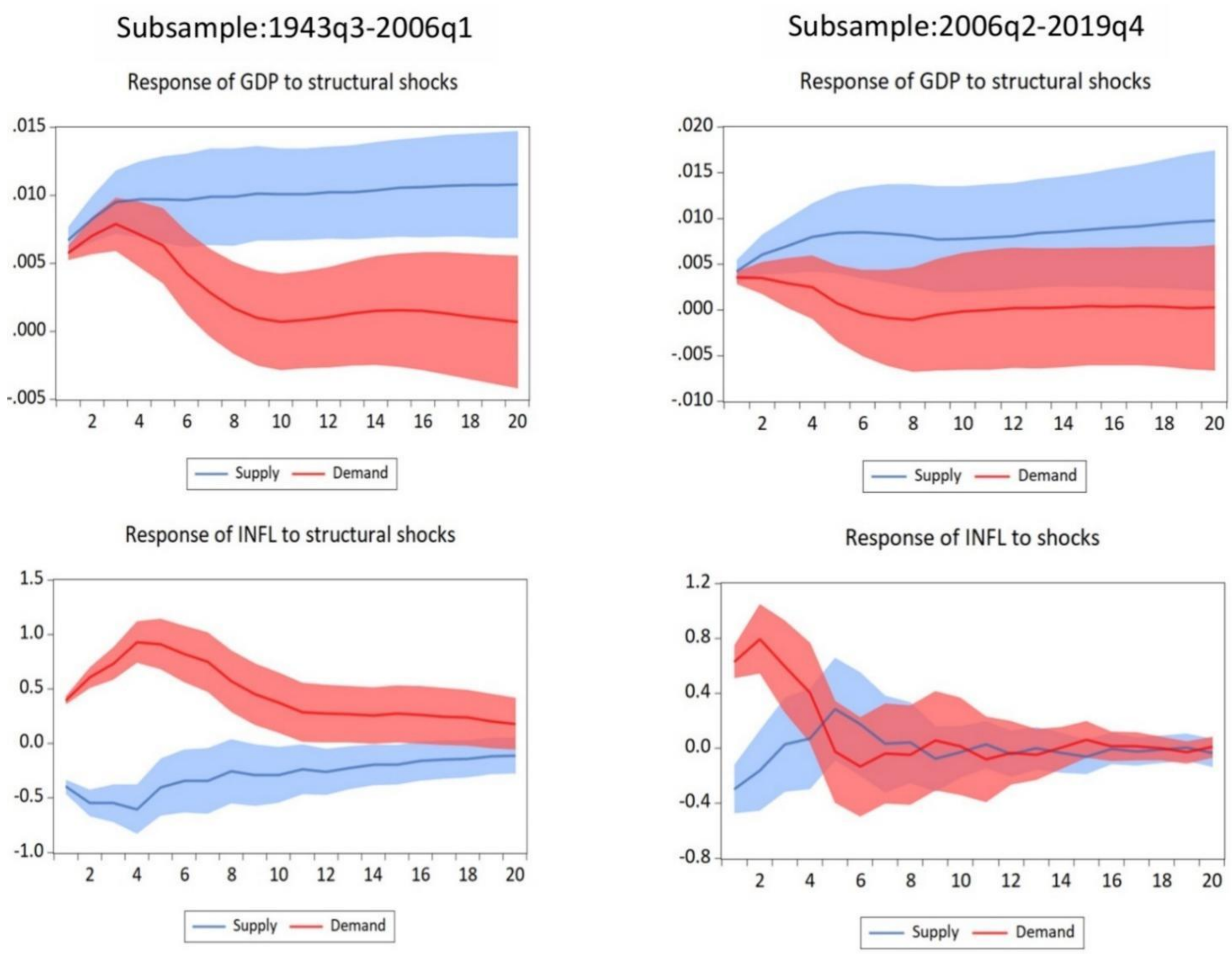

Figure 11. Subsample IRF analysis: Pre- and post-break dynamics. The graph shows the responses to 1 SD of supply and demand shocks. The confidence interval was constructed with asymptotic simulations at 95% confidence.

7. Conclusions

This paper provides new evidence on how U.S. structural supply and demand shocks propagate through the Pacific Alliance countries' monetary policy frameworks. By combining a backward-looking Taylor rule with SVAR-identified shocks and distributed-lag analysis, we document both the magnitude and persistence of spillovers, highlighting significant cross-country heterogeneity. Our work demonstrates the value of treating foreign shocks as endogenous to their home economy and underscores the need for tailored policy responses in emerging markets. Ultimately, these insights contribute to more informed macroeconomic modeling and practical guidance for central banks operating in an increasingly interconnected global environment.

Author contributions

Both authors contributed equally to the conceptualization, methodology, formal analysis, investigation, writing of the original draft, and writing, review, and editing. All authors have read and approved the final manuscript.

Use of AI tools declaration

The authors declare they have not used Artificial Intelligence (AI) tools in the creation of this article.

Conflict of interest

The authors declare no conflict of interest in this paper.

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