école d'économie d'aix-marseille
aix-marseille school of economics

Working Papers / Documents de travail

In Search of the Lost Exchange Rate Pass-Through in Mexico

Samuel Martinez
Daniel Ventosa-Santaularia

WP 2025 - Nr 36

In Search of the Lost Exchange Rate Pass-Through in Mexico*

Samuel Martínez† Daniel Ventosa-Santaulària‡

December 18, 2025

Abstract

Exchange rate pass-through (ERPT) is a key channel through which external shocks affect domestic inflation in open emerging market economies (EMEs). This paper estimates ERPT for Mexico using local projections with an instrumental-variable strategy to trace the cumulative effects of peso-dollar depreciations on consumer prices. The results indicate a relatively high degree of pass-through–higher than most estimates reported in the existing literature–suggesting that Mexican inflation responds more strongly to exchange-rate movements than commonly assumed. Although price subcomponents display the expected heterogeneity (with goods showing the highest sensitivity and services the lowest), the main implication is clear: exchange-rate fluctuations remain a powerful driver of inflation dynamics, underscoring the policy relevance of monitoring currency volatility in open EMEs such as Mexico.

Keywords: Inflation, Exchange rate Pass-through, ERPT, EMEs, Mexico, Local Projections
JEL Codes: E31, F14, F31.

---

*The project leading to this publication has received funding from the French government under the "France 2030" investment plan managed by the French National Research Agency (reference :ANR-17-EURE-0020) and from Excellence Initiative of Aix-Marseille University - A*MIDEX.

†samuel.martinez@alumnos.cide.edu

‡**Corresponding Author:** Aix Marseille Univ, CNRS, AMSE, Marseille, France & CIDE, División de Economía, Mexico. daniel.ventosa@cide.edu.

1 Introduction

Open economies are exposed to external shocks that require careful policy considerations. In such settings, the exchange rate channel plays a central role in the transmission of external disturbances and monetary policy to domestic prices.

These mechanisms are particularly relevant in emerging market economies (EMEs), where exchange rate fluctuations tend to be larger, more persistent, and more likely to affect inflation dynamics. In Mexico–characterized as a small open economy–the nominal exchange rate (ER) plays a key role in the transmission of monetary policy to inflation. Changes in the nominal ER influence the prices of goods and services through multiple channels, most notably through exchange rate pass-through (ERPT) (Mishkin, 2008).

In the literature, ERPT is defined as the exchange rate elasticity of domestic prices. When local prices fully absorb the effects of exchange rate depreciation or appreciation, ERPT equals one and is said to be "complete." Conversely, if prices adjust only partially to these changes, then ERPT is considered "incomplete" and its value is less than one.

Volatility in the exchange rate can introduce uncertainty for both firms and consumers in EMEs, influencing their decision-making processes. This uncertainty often leads to precautionary price increases as firms hedge against future fluctuations (see Calvo and Reinhart (2002)). In addition, exchange rate movements can alter production costs, especially in highly integrated supply chains. For instance, exchange rate depreciation raises the cost of imported goods–whether final consumer products directly counted in the consumer price index basket, or intermediate inputs later reflected in higher production costs– ultimately contributing to inflationary pressures (see (Forbes, 2015)).

We estimate the accumulated ERPT elasticity using local projections (LP) as in Carrière-Swallow et al. (2021) but we deviate by instrumenting the depreciation of the exchange rate with an exogenous measure of U.S. monetary surprises –an external factor of particular relevance for Mexico. We favor the LP approach because it is more robust to miss-specifications compared to traditional VAR methods, as it directly computes impulse responses at each horizon and thus is more flexible to capture the dynamic effects of exchange rate fluctuations.

Beyond documenting heterogeneity across price components, our estimates point to a level of pass-through that is quantitatively larger than what is usually found for Mexico. We find that a 1% depreciation of the Mexican peso, solely driven by changes in the U.S. monetary policy stance, leads to an accumulated increase in goods prices of about 1.10% after 24 months, while services prices rise by only 0.27% over the same horizon.

This pronounced difference highlights the asymmetric nature of exchange rate pass-through in Mexico, with inflationary pressures being transmitted â€¦ predominantly through tradable goods rather than services, and at levels that exceed those commonly found in previous empirical studies.[1]

This paper is structured as follows: The rest of the introduction presents a literature review; next, the econometric methodology is described in Section 2; then, the estimation results are detailed in Section 3; robustness exercises are outlined

[1](Banco de México, 2015, 2017, 2024). Note, however, that these studies estimate ERPT within a system of equations, making direct comparisons less straightforward.

in Section 4; and finally, Section 5 presents the main conclusions.

Literature Review

An important challenge for policymakers is understanding how external factors shape national economic activity, particularly in EMEs where exchange rate fluctuations are an important driver of inflation with considerable implications for monetary policy (see Mishkin, 2008; Forbes, 2015). Calvo and Reinhart (2002) emphasize that these effects are magnified in EMEs, where frequent currency swings often translate into immediate and forceful central bank measures.

Marked exchange rate shifts can affect aggregate spending and the terms of trade (see Obstfeld, 1982), heightening inflationary risks and complicating strategies to stabilize prices. A high ERPT indicates that depreciation rapidly transmits into domestic production costs, constraining a central bank's capacity to maintain price stability-especially if producers swiftly pass cost rises on to consumers. Conversely, a lower ERPT grants monetary authorities more freedom to address inflation without reacting to every exchange rate fluctuation.

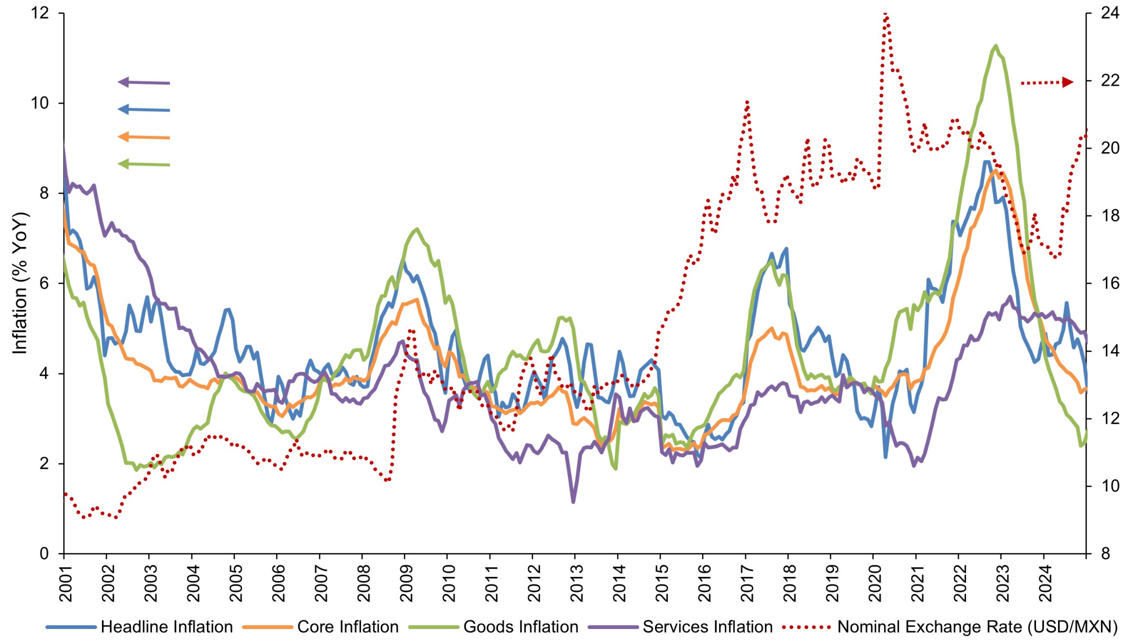
Numerous studies tie ERPT to country-specific traits, particularly the design of monetary policy. Jogrim et al. (2019) find that heightened trade openness strengthens ERPT, as imported goods have a larger presence in domestic markets. Moreover, firmly anchored inflation expectations–supported by a credible monetary framework–are viewed as among the main in tempering pass-through (see Carrière-Swallow et al., 2016; Choudhri and Hakura, 2006; Taylor, 2000). In stable, low-inflation settings, not every depreciation is seen as permanent, so firms often hesitate to adjust prices. However, persistently high inflation fosters repeated cost increases, thereby amplifying ERPT (Jašová et al., 2019; Choudhri and Hakura, 2006; Campa and Goldberg, 2005).

Research also indicates that currency swings may not yield a complete one-to-one shift in local prices. Aure and Chaney (2009) highlight "pricing-to-market" behavior, wherein exporters revise their pricing based on local demand and elasticity conditions. Likewise, Choudhri and Hakura (2015) underscore local price rigidities among competitors striving to preserve stable relative prices, thus limiting immediate pass-through and tempering short-run inflationary shocks.

Over the last decade, the Mexican ERPT literature has examined how Banco de Mexico's 2001 inflation-targeting regime reshaped domestic price behavior. Evidence suggests that this policy significantly curbed consumer price inflation, yet pronounced heterogeneity persists across different components of the National Consumer Price Index (INPC).[2] As depicted in Figure 1, certain periods of sharp depreciation–often linked to external shocks–have been accompanied by changes in inflation. This is particularly relevant given Mexico's high degree of trade openness, where import price variations can rapidly pass through to final consumers. Consequently, a deeper understanding of how exchange rate movements permeate inflation remains pivotal in crafting effective and timely policy responses, ensuring stability amid global market fluctuations. Therefore, the interaction among currency changes, domestic price formation, and policy action remains an ongoing focus for researchers examining open EMEs.

[2] See González and Saucedo, 2018; Cortés, 2013; Capistrán et al., 2012; Aleem and Lahiani, 2013.

Figure 1: Inflation and depreciation of the USD/MXN exchange rate.



Source: Authors' calculations using data from INEGI, and Banco de México.

Authors' calculations based on data from INEGI, Banco de México, and the Federal Reserve Bank of St. Louis.

2 Econometric Methods

We examine the impact of peso-dollar exchange rate depreciation on inflation across various consumer price indices (headline, core, services, and goods) using monthly data from January 2003 to December 2024. Our reduced-form specification follows standard empirical models (Carrière-Swallow et al., 2016; Capistrán et al., 2012; Campa and Goldberg, 2005) to estimate the cumulative response of inflation for each index. Specifically, we employ the local projection (LP) methodology introduced by Jordà (2005),[3] which allows us to measure the accumulated response of consumer prices in Mexico h periods ahead, following a 1 percent exchange rate depreciation. The LP estimates correspond to the coefficients β_h in the following specification:

$$p_{t+h} - p_t = \alpha_h + \beta_h \Delta ner_t + \theta_h X_t + \varepsilon_{t+h},\ h = 1, 2, \ldots, H, \tag{1}$$

where p_t is the natural logarithm of the consumer price index at time t, and Δner_t denotes the first difference of the log of the peso-dollar exchange rate. The vector X_t includes 12 lags of monthly inflation and exchange rate depreciation, Mexico's 1-month interbank equilibrium interest rate (TIIE), and indicators of economic activity. Given the close economic ties between Mexico and the U.S., we also incorporate the U.S. industrial production index, the U.S. consumer price index (both in log differences), 1-year ahead U.S. inflation expectations, and the 10-year U.S. Treasury yield. To control for commodity-price volatility, we use the IMF commodity price index (in log differences) and include the U.S. interest rate to capture capital-flow effects.

Nevertheless, the above specification may suffer from endogeneity: Higher inflation can weaken the currency's purchasing power and demand, while a depreciated exchange rate can further elevate inflation by raising the cost of imported goods. To address this issue, we use an identification strategy with exogenous instruments to estimate β_h in equation (1).

[3]For further details on the methodology, see Appendix A.

3 Identification with external instruments

We adopt an identification strategy based on external instruments, following Jordà et al. (2020), Carrière-Swallow et al. (2023), and Hernández et al. (2024), to obtain an exogenous measure of Δner_t. Specifically, we isolate the impact of exchange rate fluctuations driven by unanticipated changes in U.S. monetary policy by estimating the effects on the U.S. interest rate as the residuals obtained from projecting it onto its main determinants:

$$\Delta FFR_t = \alpha + \gamma_h x_t + \eta_t, \tag{2}$$

where ΔFFR_t is the first difference of the federal funds rate, and x_t includes the U.S. industrial production index, the U.S. Consumer Price Index (both in logarithmic differences), 1-year ahead inflation expectations, and the 10-year market yield on U.S. Treasury bonds. The resulting Fed policy shocks, $\widehat{\eta}_t$, are orthogonal to U.S. economic activity, inflationary pressures, and global financial conditions, which follows from the properties of OLS projections. We then use $\widehat{\eta}_t$ as an instrumental variable for Δner_t.

According to Jordà (2023), the formal conditions for this instrumental-variables approach to be valid rely on the typical local projection setup:

$$y_{t+h} = \alpha_h + \beta_h s_t + \gamma_h x_t + v_{t+h}, \quad h = 0, 1, 2, \ldots, H. \tag{3}$$

Let z_t be the instrument for the variable s_t. Instrument relevance requires that $\text{Cov}(z_t, s_t) \neq 0$, while instrument exogeneity is established if $\text{Cov}(z_t, v_{t+h}) = 0$

for each h–a condition typically referred to as lead-lag exogeneity.

In our setting, relevance follows from Mexico's status as a small, open economy which is influenced by U.S. monetary policy. Such policy decisions affect capital inflows shaping global investor risk-appetite, therefore affecting the demand for different kinds of assets in EMEs. Given the high liquidity of the Mexican peso, these shifts in risk appetite are reflected in movements of the MXN-USD exchange rate.

Empirically, we find that $\mathrm{Cov}(\Delta ner_t, \widehat{\eta}_t) = -0.0012$, with a t-statistic of -4.5714. Since this value is statistically different from zero, the relevance condition is satisfied.

Regarding lead-lag exogeneity required by Stock and Watson (2018), we observe that the estimated residuals $\widehat{\varepsilon}_{t+h}$ in Equation (1) are orthogonal to $\widehat{\eta}_t$, again due to the geometry of ordinary least squares. The first-stage regression of our instrumental variables approach is:

$$\Delta ner_t = \delta_0 + \psi \widehat{\eta}_t + \delta x_t + u_t. \tag{4}$$

We then estimate equation (1) by replacing Δner_t with its fitted values $\widehat{\Delta ner_t}$ from the first-stage regression. Because these fitted values incorporate all the exogenous information from $\widehat{\eta}_t$, they remain orthogonal to the error term $\widehat{\varepsilon}_{t+h}$ in the main equation. In other words, once $\widehat{\eta}_t$ has been used to predict Δner_t, any correlation with $\widehat{\varepsilon}_{t+h}$ is eliminated by construction–an outcome stemming from the geometry of ordinary least squares. This orthogonality satisfies the lead-lag exogeneity condition.

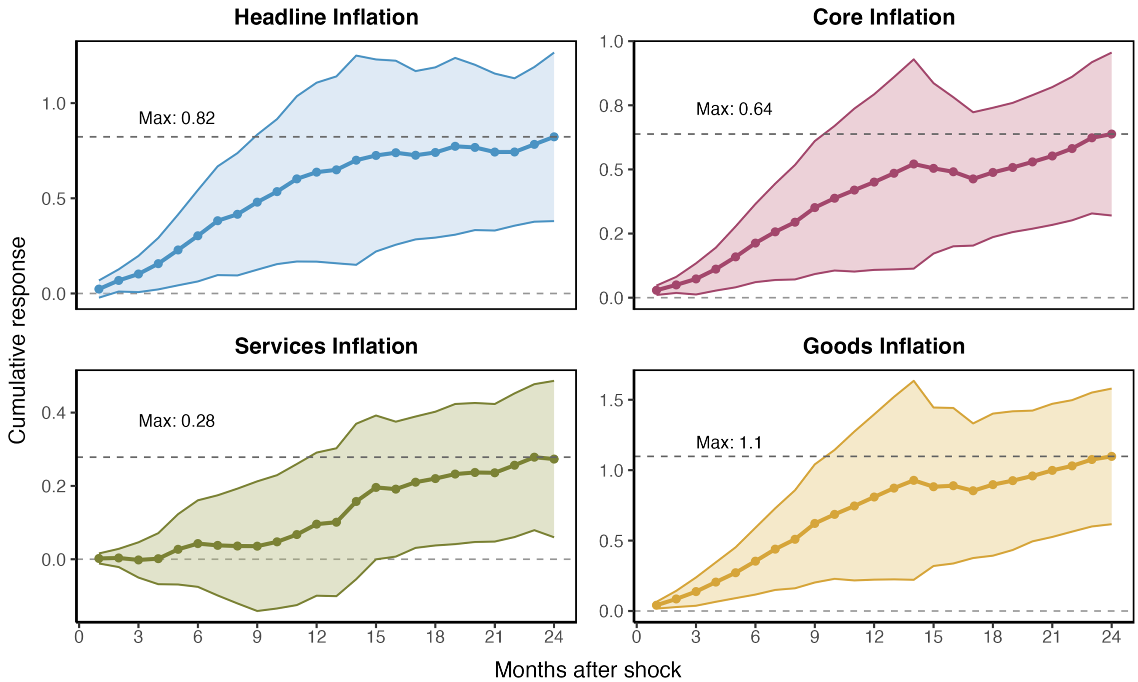
Finally, as proposed by Jordà (2005), we employ Newey–West standard errors (Newey and West, 1987) to account for heteroskedasticity and autocorrelation, ensuring robust inference under more general conditions.

4 Results

The main findings are illustrated by estimating how a 1% depreciation of the peso-dollar exchange rate affects consumer prices over a 24-month horizon, along with 90% confidence intervals. These results are presented as impulse-response functions (IRFs) in Figure 2, encompassing headline, core, services, and goods inflation.[4]

[4] Additional weak-instrument tests are also provided in Appendix B to ensure robust and reliable inference on the exchange rate pass-through (ERPT).

Figure 2: Exchange Rate Pass-Through



Source: Authors' calculations using data from INEGI, Banco de México, and the Federal Reserve Bank of St. Louis.

Several factors shape the heterogeneous price responses to exchange rate fluctuations. One critical aspect is the distinction between goods and services. Goods typically exhibit a more immediate and pronounced pass-through, as they often rely heavily on imported inputs in a small, open, and highly integrated EME. In contrast, services generally have lower import content, rendering them less sensitive to exchange rate movements.

The timing of price adjustments is also pivotal. Exchange rate variations do not always feed into production costs immediately, as medium–to long-term contracts–particularly with the United States, Mexico's principal trading partner–can dampen or delay the pass-through effect. These contractual arrangements help cushion unanticipated changes in trade conditions.

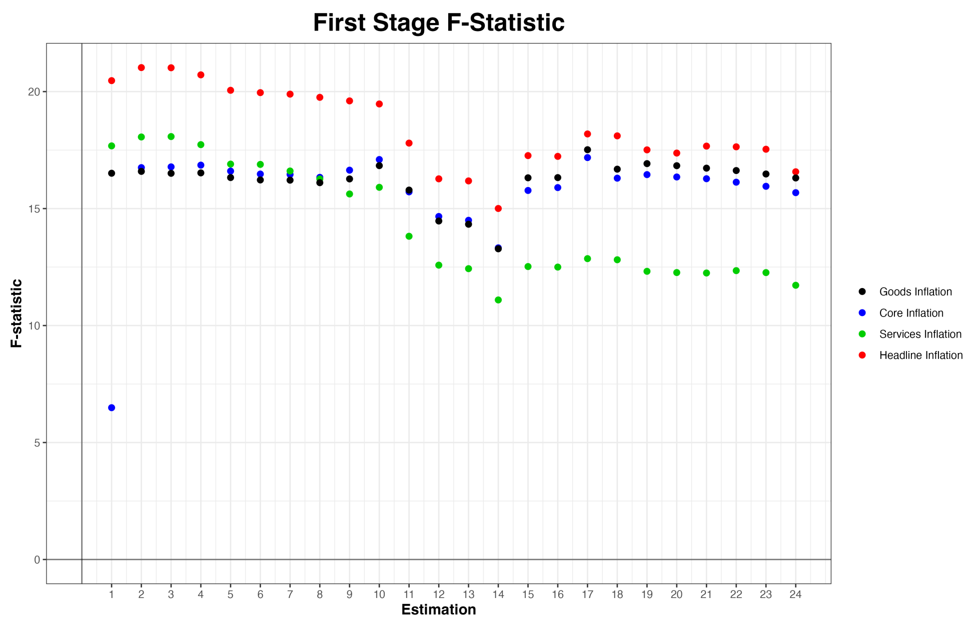
Such contractual rigidities lead to delayed price adjustments, making firms less responsive to short-lived exchange rate shifts until contracts are renewed. Moreover, inventory strategies can further postpone price revisions: Companies may rely on existing stocks before passing on higher import costs to consumers.

To gauge the strength of our instruments, we follow the common practice of performing a first-stage F-test at each forecast horizon, testing the null hypothesis that the instrument is unrelated to the endogenous variable.[5] Figure 3 (panel a) shows these results: the horizontal axis denotes the LP horizons, while the vertical axis displays the corresponding first-stage F-statistics.

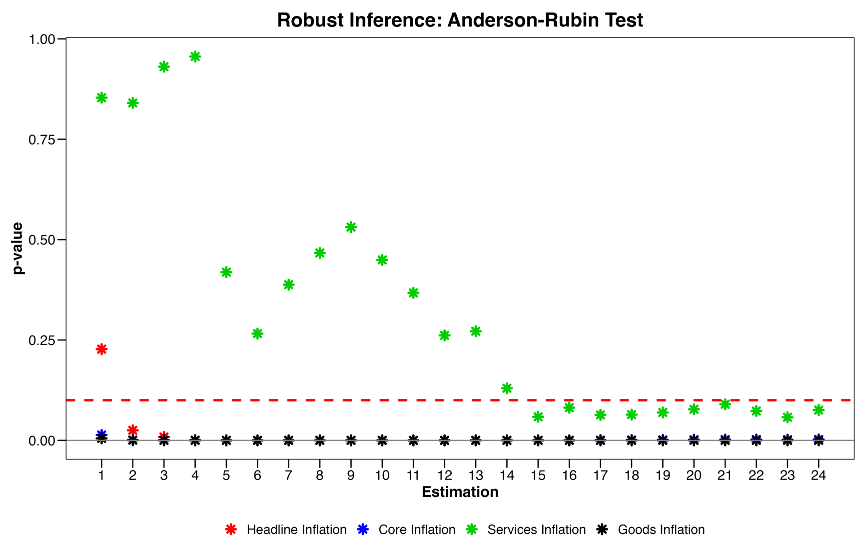
[5] IV estimation requires instruments that are sufficiently correlated with the endogenous regressor and uncorrelated with the error term. Weak instruments can distort inference in local projections (LP), leading to biased estimates of causal effects (Jordà and Mertens, 2022).

Figure 3: F-statistics: Week Instruments Test.

(a) F-statistics: Week Instruments Test



(b) Anderson-Rubin Test



Source: Authors' calculations using data from INEGI, Banco de México, and the Federal Reserve Bank of St. Louis.

For all four inflation measures (headline, core, goods, and services), the first-stage test statistic confirms that our instrument is highly relevant, except at the one-period horizon for core inflation. Thus, the instruments are sufficiently correlated with the endogenous variable to yield reliable estimates.[6]

The AR test specifically evaluates whether the estimated parameters differ significantly from zero when instruments might be weak. Figure 3 (panel b) displays its results for each price index, with a 10% significance threshold marked by the horizontal line. Points below this line indicate rejection of the null hypothesis (i.e., evidence of a statistically significant parameter), while points above it suggest that the parameter estimate is not statistically different from zero. Generally, we fail to reject the null of zero pass-through in cases where the instrument appears weak. In scenarios where the instrument proves to be strong, the exchange rate pass-through (ERPT) effect is typically significant.

5 Robustness checks

To further validate the stability of our findings, we conduct two additional robustness exercises where we restrict the dataset to pre-pandemic observations, and we replace our baseline index with an alternative measure. These adjustments allow us to verify whether our main results hold when either the sample period or the

[6] To reinforce the evidence on instrument relevance we also conduct the Anderson–Rubin (AR) test proposed by Jordà and Mertens (2022), designed for local projection inference under potentially weak instruments. Details are provided in Appendix B. The AR test confirms valid inference across the various horizons for headline (except at the first period), core, and goods inflation; for services, the effect becomes significant after roughly one year.

core indicator changes. The fundamental conclusions remain intact: Although certain coefficients show minor shifts in magnitude or significance, these deviations align with the expected differences in sample composition and the nature of the alternative metric. Overall, these outcomes reinforce the reliability of our empirical strategy, suggesting that the relationships identified are not overly sensitive to either the chosen time frame or the specific index employed. The two aforementioned exercises are detailed below:

- *Robustness 1.* Considering that our estimates span the period of the COVID-19 pandemic, we explore how our results would change if we truncated the data through December 2019.
- *Robustness 2.* We estimate the model with an alternative variable for IMF commodity prices, including the Food and Agriculture Organization (FAO) commodity index, being alternative indicators of international commodity prices.

Full details and accompanying results are provided in Appendix C, where the robustness checks are conducted for headline, core, services, and goods inflation. Truncating the data at December 2019 yields a lower pass-through estimate, likely reflecting the substantial disruptions and volatility introduced by the COVID-19 pandemic. The pandemic reshaped global supply chains, demand patterns, and price dynamics in ways that diverge from typical economic conditions. Similarly, employing the FAO commodity price index instead of the IMF index also results in a lower pass-through estimate. This difference can be attributed to the FAO index's distinct construction, coverage, and periodicity, which emphasize agricultural products and may capture a narrower range of price fluctuations. Consequently, the FAO index could reflect less volatility in non-agricultural commodities, thus altering both the measured price movements and their broader economic impact. Nevertheless, as documented in the Appendix, the fundamental relationships hold across these specifications, underscoring the overall robustness of our findings.

6 Concluding Remarks

This paper contributes to the existing literature by examining the Exchange Rate Pass-Through (ERPT) on consumer prices in Mexico, with a focus on different price indices (headline, core, services, and goods). By employing Local Projections and an identification strategy based on an instrumental variable approach, the analysis sheds light on how exchange rate fluctuations affect differently various components of the consumer price basket.

The estimates indicate that when a depreciation is entirely explained by an unexpected change in the monetary policy rate of the U.S., the ERPT to consumer prices is generally more pronounced in the goods indices, a result that aligns in direction with prior findings in EMEs. The direct influence of the exchange rate on tradable goods helps explain why prices in this category exhibit higher sensitivity. In contrast, the services indices show considerably less responsiveness, suggesting that services, being less exposed to imported inputs and global supply chain pressures, tend to be more insulated from exchange rate volatility.

From a policy perspective, these findings underscore the importance of monitoring exchange rate movements and considering the lag of the ERPT when deciding on changes to the monetary policy interest rate. This is especially true in economies where the tradable goods sector represents a substantial share of consumption as in a small open EME. Central banks and policymakers may benefit from distinguishing between ERPT to goods- and services-based inflation, tailoring their policy actions accordingly. By focusing on the categories most affected by exchange rate swings, authorities can better anticipate inflationary pressures and refine their monetary policy responses.

Although our analysis points to the robustness of the ERPT patterns in Mexico, further research could explore potential nonlinearities in pass-through dynamics or investigate whether the effects differ over time in response to changes in inflation regimes, trade agreements, or macroeconomic volatility. Furthermore, expanding the scope to incorporate sector-specific cost structures and competitive environments could provide deeper insights into the underlying mechanisms of exchange rate transmission.

Overall, this study confirms that the exchange rate remains a critical driver of inflation dynamics in small open emerging economies, particularly within the tradable sector. Recognizing these differential impacts enhances our understanding of how open economies, such as Mexico, can react to external shocks that threaten price stability.

References

Aleem, A. and A. Lahiani (2013). A threshold vector autoregression model of exchange rate pass-through in mexico. *Research in International Business and Finance 30*(1), 24–33.

Aure, R. and T. Chaney (2009). Exchange rate pass-through in a competitive model of pricing-to-market. *Journal of Money, Credit and Banking 41*(1), 151–175.

Banco de México (2015). Traspaso de variaciones en el tipo de cambio a precios en economías de latinoamérica. Informe Trimestral Julio-Septiembre 2015, Recuadro 1, 18-20.

Banco de México (2017). Evolución del traspaso del tipo de cambio a la inflación. Informe Trimestral Abril-Junio 2017, Recuadro 1, 7-13.

Banco de México (2024). Evolución del traspaso del tipo de cambio a la inflación. Informe Trimestral Abril-Junio 2024, Recuadro 5, 86-88.

Calvo, G. and C. Reinhart (2002). Fear of floating. *Quarterly Journal of Economics 107*(2), 379–408.

Campa, J. and L. Goldberg (2005). Exchange rate pass-through into import prices. *Review of Economics and Statistics 87*(4), 679–690.

Capistrán, C., R. Ibarra, and M. Ramos-Francia (2012). El traspaso de movimientos del tipo de cambio a los precios: Un análisis para mexico. *El Trimestre Económico 74*(316), 813–838.

Carrière-Swallow, Y., M. Firat, M. Furceri, and D. Jiménez (2023). State-dependent exchange rate pass-through. Technical Report 23/86, International Monetary Fund.

Carrière-Swallow, Y., B. Gruss, E. Magud, and F. Valencia (2016). Monetary policy credibility and exchange rate pass-through. Technical Report 16/240, International Monetary Fund.

Carrière-Swallow, Y., B. Gruss, E. Magud, and F. Valencia (2021). Monetary policy credibility and exchange rate pass-through. *International Journal of Central Banking 17*(3), 61–94.

Choudhri, U. and S. Hakura (2006). Exchange rate pass-through to domestic prices: Does the inflationary environment matter? *Journal of International Money and Finance 25*(4), 614–639.

Choudhri, U. and S. Hakura (2015). The exchange rate pass-through to import and export prices: The role of nominal rigidities and currency choice. *Journal of International Money and Finance 51*, 1–25.

Cortés, J. (2013). Estimating the exchange rate pass-through to prices in mexico. *Monetaria 1*(2), 287–316.

Forbes, K. (2015). Much ado about something important: How do exchange rate movements affect inflation? *The Manchester School 84*(S1), 15–41.

González, J. and E. Saucedo (2018). Traspaso depreciación-inflación en mexico:

Análisis de precios al consumidor y productor. *Revista Mexicana de Economía y Finanzas 13*(4), 525–545.

Hernández, J., D. Ventosa-Santaulària, and J. Valencia (2024). Global supply chain inflationary pressures and monetary policy in mexico. *Emerging Markets Review 58*, 101089.

Jašová, M., R. Moessner, and E. Takáts (2019). Exchange rate pass-through: What has changed since the crisis? *International Journal of Central Banking 15*(3), 27–58.

Jogrim, H., A. M. Kose, and F. Ohnsorge (2019). *Inflation in Emerging and Developing Economies: Evolution, Drivers and Policies*. Word Bank Publications.

Jordà, O. (2005). Estimation and inference of impulse responses by local projections. *American Economic Review 95*(1), 161–182.

Jordà, O. (2023). Local projections for applied economics. Technical Report 2023-16, Federal Reserve Bank of San Francisco.

Jordà, O. and K. Mertens (2022). Techniques of empirical econometrics.

Jordà, O., M. Schularick, and A. M. Taylor (2020). The effects of quasi-random monetary experiments. *Journal of Monetary Economics 112*, 22–40.

Mishkin, F. (2008). Exchange rate pass-through and monetary policy. Technical Report 13889, National Bureau of Economic Research.

Newey, W. K. and K. D. West (1987). A simple, positive semi-definite heteroskedasticity and autocorrelation consistent covariance matrix. *Econometrica 55*(3), 703–708.

Obstfeld, M. (1982). Aggregate spending and the terms of trade: Is there a laursen-metzler effect? *Quarterly Journal of Economics 97*(2), 251–270.

Stock, J. H. and M. W. Watson (2018, 05). Identification and Estimation of Dynamic Causal Effects in Macroeconomics Using External Instruments. *The Economic Journal 128*(610), 917–948.

Taylor, J. (2000). Low inflation, pass-through, and the pricing power of firms. *European Economic Review 44*(7), 1389–1408.

Appendix

A Local Projections

Let $\omega_t = (\omega_{1t}, \ldots, \omega_{jt}, \ldots, \omega_{kt})'$ be a vector of stationary random variables observed over $t = 1, \ldots, T$ periods. Assume that ω_t follows a VAR(1) process:

$$(\omega_t - \mu) = A(\omega_{t-1} - \mu) + \varepsilon_t; \;\; \varepsilon_t \sim D(0, \Omega). \tag{5}$$

We define the response of ω_{jt+h} due to a shock of size δ_i in ϵ_t as

$$\mathcal{R}_{ij}(h) = \mathbb{E}[\omega_{jt+h} \mid \varepsilon_t = \delta_i; \; \omega_{t-1}] - \mathbb{E}[\omega_{jt+h} \mid \varepsilon_t = 0; \; \omega_{t-1}] = A^h_{[j,]}\delta_i,$$

for $h = 0, \ldots, H$ and where $A^h_{[j,]}$ denotes the j-th row of the matrix A raised to the h-th power. The local projections methodology, proposed by Jordà (2005), begins by using recursive substitution:

$$(\omega_{t+h} - \mu) = A^{h+1}(\omega_{t-1} - \mu) + A^h\varepsilon_t + \ldots A^0\varepsilon_{t+h}; \text{ with } A^0 = I.$$

This suggests that the regression:

$$\omega_{jt+h} = c_{jh} + \beta_{jh+1}\omega_{t-1} + v_{jt+h}; \;\; \text{where } v_{t+h} = A^h\varepsilon_t + \cdots + A^0\varepsilon_{t+h} \tag{6}$$

for $h = 0, 1, \ldots, H$ provides an estimator of the impulse-response functions because

$$\mathcal{R}_{ij}(h) = \mathbb{E}[\omega_{jt+h} \mid \varepsilon_t = \delta_i;\ \omega_{t-1}] - \mathbb{E}[\omega_{jt+h} \mid \varepsilon_t = 0;\ \omega_{t-1}] = \beta_{jh}\delta_i,$$

which is equivalent to $A^h_{[j,]}\delta_i$ as long as the DGP coincides with that in 5.

B P-values Weak Instrument Test

Anderson–Rubin (AR) test proposed by Jordà and Mertens (2022), designed for local projection inference under potentially weak instruments.

This table presents the p-values from the AR test, which evaluates whether the instruments used in our instrumental variables setup are collectively relevant and not weak. Specifically, the AR test checks if the endogenous regressor(s) in the second-stage equation are jointly significant when projected onto the instruments. A small p-value (e.g., below 0.05) leads us to reject the null hypothesis that the coefficients on the endogenous regressor(s) are jointly zero and that the model is unidentified. From the table, we see that the p-values for all columns fall below conventional significance thresholds. Thus, we reject the null hypothesis of weak instruments, indicating that our instruments are sufficiently strong and relevant. Consequently, the estimates derived from the two-stage least squares (2SLS) procedure are less likely to suffer from the biases typically associated with weak identification. These findings bolster the credibility of our main results by demonstrating that instrument weakness does not compromise the validity of our empirical strategy.

Table 1: P-values Weak Instrument Test

Estimation	Headline	Core	Services	Goods
1	0.0000	0.0003	0.0000	0.0001
2	0.0000	0.0001	0.0000	0.0001
3	0.0000	0.0001	0.0000	0.0001
4	0.0000	0.0001	0.0000	0.0001
5	0.0000	0.0001	0.0000	0.0001
6	0.0000	0.0001	0.0000	0.0001
7	0.0000	0.0001	0.0000	0.0001
8	0.0000	0.0001	0.0000	0.0001
9	0.0000	0.0001	0.0000	0.0001
10	0.0000	0.0001	0.0000	0.0001
11	0.0000	0.0001	0.0000	0.0001
12	0.0000	0.0001	0.0000	0.0001
13	0.0001	0.0001	0.0000	0.0001
14	0.0000	0.0003	0.0005	0.0005
15	0.0000	0.0003	0.0000	0.0004
16	0.0000	0.0001	0.0000	0.0001
17	0.0000	0.0001	0.0000	0.0001
18	0.0000	0.0001	0.0004	0.0001
19	0.0000	0.0001	0.0004	0.0001
20	0.0000	0.0001	0.0001	0.0001
21	0.0000	0.0001	0.0000	0.0001
22	0.0000	0.0001	0.0000	0.0001
23	0.0000	0.0001	0.0000	0.0001
24	0.0001	0.0001	0.0000	0.0001

Source: Authors' calculations using data from INEGI, Banco de México, and the Federal Reserve Bank of St. Louis.

C Robustness test results

The IRFs presented in this Appendix show patterns that are broadly consistent with those of our baseline models . While the robustness checks (e.g., pre-COVID subsample, alternative commodity price indices) lead to minor shifts in the magnitude or timing of responses, the overall shapes and statistical significance of the IRFs remain similar. This coherence suggests that our main conclusions about the dynamic effects under study hold across different model specifications and sample periods, reinforcing the stability and reliability of our empirical findings.

C.1 Headline and core inflation

The following Tables report the impulse responses under two robustness checks for headline inflation (Table 2), and for Core Inflation (Tables 3). The first column (*Robustness 1*) restricts the sample to pre-COVID observations, thereby excluding any pandemic-related shocks. The second column (*Robustness 2*) replaces the IMF commodity price index with the FAO commodity price index. In both cases (and for both variables), the impulse responses remain qualitatively similar to those in the baseline specification, suggesting that neither the presence of COVID-era data nor the choice of commodity price index substantially alters the dynamic relationships of interest. These findings reinforce the robustness of our main results and confirm that the core insights hold under alternative sample periods and data sources.

Table 2: Robustness for Headline Inflation

		Months after shock					
Month	Baseline	Robust. 1	Robust. 2	Month	Baseline	Robust. 1	Robust. 2
1	0.0231	0.0628	0.0164	13	0.6495	0.6653	0.4416
	(0.0275)	(0.0351)	(0.0271)		(0.2984)	(0.4631)	(0.1626)
2	0.0685	0.0803	0.0523	14	0.7001	0.6962	0.4744
	(0.0354)	(0.0588)	(0.0322)		(0.3341)	(0.4915)	(0.1735)
3	0.1023	0.1094	0.0777	15	0.7250	0.7361	0.5432
	(0.0579)	(0.0917)	(0.0409)		(0.3066)	(0.4821)	(0.1815)
4	0.1567	0.1795	0.1195	16	0.7394	0.7565	0.5545
	(0.0822)	(0.1336)	(0.0676)		(0.2939)	(0.4822)	(0.1754)
5	0.2286	0.2663	0.1730	17	0.7262	0.7726	0.5519
	(0.1130)	(0.1942)	(0.0853)		(0.2688)	(0.4857)	(0.1683)
6	0.3031	0.3593	0.2313	18	0.7406	0.7655	0.5592
	(0.1459)	(0.2510)	(0.1062)		(0.2719)	(0.5100)	(0.1628)
7	0.3822	0.4428	0.2928	19	0.7732	0.7320	0.5865
	(0.1736)	(0.2917)	(0.1219)		(0.2823)	(0.4840)	(0.1523)
8	0.4161	0.5084	0.3153	20	0.7672	0.7122	0.5902
	(0.1953)	(0.3327)	(0.1416)		(0.2640)	(0.4624)	(0.1444)
9	0.4797	0.5715	0.3611	21	0.7429	0.6823	0.5733
	(0.2157)	(0.3728)	(0.1547)		(0.2506)	(0.4372)	(0.1348)
10	0.5351	0.6213	0.4026	22	0.7433	0.6965	0.5727
	(0.2315)	(0.3991)	(0.1626)		(0.2354)	(0.4233)	(0.1381)
11	0.6020	0.6313	0.4377	23	0.7833	0.7070	0.6160
	(0.2639)	(0.3728)	(0.1686)		(0.2470)	(0.4089)	(0.1408)
12	0.6373	0.6458	0.4381	24	0.8229	0.7063	0.6476
	(0.2858)	(0.4221)	(0.1621)		(0.2691)	(0.4129)	(0.1484)

Source: Authors' calculations using data from INEGI, Banco de México, and the Federal Reserve Bank of St. Louis.

Table 3: Robustness for Core Inflation

		Months after shock					
Month	Baseline	Robust. 1	Robust. 2	Month	Baseline	Robust. 1	Robust. 2
1	0.0288	0.0356	0.0209	13	0.4851	0.5689	0.3320
	(0.0111)	(0.0204)	(0.0094)		(0.2279)	(0.3714)	(0.1164)
2	0.0498	0.0760	0.0388	14	0.5211	0.6054	0.3564
	(0.0190)	(0.0391)	(0.0161)		(0.2480)	(0.3800)	(0.1214)
3	0.0731	0.1155	0.0549	15	0.5042	0.6306	0.3907
	(0.0369)	(0.0693)	(0.0266)		(0.2019)	(0.3939)	(0.1212)
4	0.1113	0.1639	0.0829	16	0.4909	0.6307	0.3878
	(0.0509)	(0.1104)	(0.0390)		(0.1767)	(0.3978)	(0.1104)
5	0.1591	0.2263	0.1177	17	0.4630	0.6363	0.3760
	(0.0722)	(0.1506)	(0.0520)		(0.1580)	(0.4124)	(0.0944)
6	0.2128	0.2891	0.1568	18	0.4882	0.6458	0.3889
	(0.0926)	(0.1820)	(0.0651)		(0.1533)	(0.4173)	(0.0930)
7	0.2567	0.3299	0.1883	19	0.5076	0.6347	0.3999
	(0.1141)	(0.2203)	(0.0787)		(0.1531)	(0.4096)	(0.0849)
8	0.2943	0.3719	0.2147	20	0.5293	0.6525	0.4164
	(0.1357)	(0.2451)	(0.0898)		(0.1583)	(0.4192)	(0.0867)
9	0.3513	0.4144	0.2562	21	0.5524	0.6458	0.4355
	(0.1576)	(0.2734)	(0.1020)		(0.1633)	(0.4140)	(0.0898)
10	0.3877	0.4572	0.2858	22	0.5814	0.6478	0.4568
	(0.1714)	(0.3023)	(0.1105)		(0.1701)	(0.4167)	(0.0931)
11	0.4195	0.4902	0.3017	23	0.6232	0.6480	0.4835
	(0.1932)	(0.3135)	(0.1143)		(0.1792)	(0.4177)	(0.0966)
12	0.4506	0.5377	0.3094	24	0.6379	0.6269	0.4959
	(0.2081)	(0.3480)	(0.1082)		(0.1931)	(0.4108)	(0.1017)

Source: Authors' calculations using data from INEGI, Banco de México, and the Federal Reserve Bank of St. Louis.

C.2 Services and Goods Inflation

We apply the same robustness checks to services inflation (Tables 4), and goods inflations (Tables 5) and the impulse-response functions remain essentially unchanged, mirroring those of the baseline model.

Table 4: Robustness for Services Inflation

	Months after shock						
Month	Baseline	Robust. 1	Robust. 2	Month	Baseline	Robust. 1	Robust. 2
1	0.0021	0.0150	0.0024	13	0.1010	0.1881	0.0522
	(0.0083)	(0.0161)	(0.0068)		(0.1226)	(0.1824)	(0.0885)
2	0.0035	0.0257	0.0039	14	0.1576	0.1958	0.0940
	(0.0151)	(0.0300)	(0.0120)		(0.1289)	(0.1729)	(0.0895)
3	-0.0019	0.0389	-0.0015	15	0.1959	0.1715	0.1386
	(0.0292)	(0.0495)	(0.0243)		(0.1193)	(0.1248)	(0.0871)
4	0.0015	0.0678	-0.0005	16	0.1912	0.1264	0.1358
	(0.0423)	(0.0764)	(0.0352)		(0.1119)	(0.0981)	(0.0846)
5	0.0271	0.1183	0.0191	17	0.2102	0.1054	0.1544
	(0.0583)	(0.1143)	(0.0475)		(0.1089)	(0.0865)	(0.0846)
6	0.0427	0.1612	0.0309	18	0.2200	0.0754	0.1566
	(0.0717)	(0.1421)	(0.0582)		(0.1109)	(0.0742)	(0.0870)
7	0.0378	0.1669	0.0246	19	0.2323	0.0489	0.1696
	(0.0831)	(0.1676)	(0.0654)		(0.1162)	(0.0706)	(0.0897)
8	0.0360	0.1863	0.0231	20	0.2367	0.0409	0.1754
	(0.0952)	(0.1900)	(0.0744)		(0.1153)	(0.0740)	(0.0890)
9	0.0357	0.1949	0.0194	21	0.2358	0.0249	0.1749
	(0.1075)	(0.2010)	(0.0824)		(0.1141)	(0.0807)	(0.0905)
10	0.0475	0.1941	0.0267	22	0.2564	0.0356	0.1913
	(0.1105)	(0.1988)	(0.0846)		(0.1191)	(0.0872)	(0.0908)
11	0.0674	0.1836	0.0368	23	0.2784	0.0473	0.2075
	(0.1170)	(0.1833)	(0.0876)		(0.1211)	(0.0967)	(0.0918)
12	0.0957	0.1849	0.0514	24	0.2732	0.0248	0.2021
	(0.1186)	(0.1856)	(0.0853)		(0.1297)	(0.1047)	(0.0941)

Source: Authors' calculations using data from INEGI, Banco de México, and the Federal Reserve Bank of St. Louis.

Table 5: Robustness for Goods Inflation

Months after shock							
Month	Baseline	Robust. 1	Robust. 2	Month	Baseline	Robust. 1	Robust. 2
1	0.0410	0.0607	0.0317	13	0.8725	0.9370	0.6229
	(0.0146)	(0.0296)	(0.0117)		(0.3938)	(0.6291)	(0.2105)
2	0.0856	0.1310	0.0666	14	0.9279	0.9935	0.6594
	(0.0350)	(0.0679)	(0.0270)		(0.4294)	(0.6759)	(0.2321)
3	0.1379	0.2009	0.1043	15	0.8823	1.0714	0.7054
	(0.0611)	(0.1094)	(0.0423)		(0.3420)	(0.7127)	(0.2326)
4	0.2054	0.2770	0.1567	16	0.8895	1.1210	0.7219
	(0.0851)	(0.1705)	(0.0598)		(0.3356)	(0.7050)	(0.2302)
5	0.2718	0.3632	0.2062	17	0.8539	1.1501	0.7057
	(0.1097)	(0.2223)	(0.0784)		(0.2905)	(0.7410)	(0.2065)
6	0.3538	0.4482	0.2680	18	0.8972	1.1950	0.7288
	(0.1442)	(0.2667)	(0.1002)		(0.3066)	(0.7671)	(0.2096)
7	0.4393	0.5237	0.3343	19	0.9256	1.2066	0.7422
	(0.1761)	(0.3161)	(0.1212)		(0.2994)	(0.8130)	(0.1886)
8	0.5093	0.5767	0.3852	20	0.9588	1.2411	0.7673
	(0.2116)	(0.3588)	(0.1435)		(0.2823)	(0.7998)	(0.1864)
9	0.6218	0.6553	0.4719	21	0.9985	1.2560	0.7995
	(0.2549)	(0.4120)	(0.1701)		(0.2874)	(0.7969)	(0.1874)
10	0.6861	0.7352	0.5267	22	1.0304	1.2374	0.8211
	(0.2782)	(0.4675)	(0.1898)		(0.2838)	(0.7642)	(0.1837)
11	0.7458	0.7973	0.5614	23	1.0757	1.2303	0.8503
	(0.3214)	(0.5118)	(0.1993)		(0.2890)	(0.7978)	(0.1827)
12	0.8094	0.8903	0.5799	24	1.0979	1.2051	0.8701
	(0.3567)	(0.5859)	(0.1956)		(0.2928)	(0.7843)	(0.1737)

Source: Authors' calculations using data from INEGI, Banco de México, and the Federal Reserve Bank of St. Louis.